ONDERZOEKSRAPPORT NR 9320

Persistence in Foreign Exchange Rates

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* This paper is an updated version of working paper D-93-02 of the Center for International Business Education and Research, University of South Carolina.

ABSTRACT

Prior studies have investigated the presence of a unit root in nominal exchange rates. This paper extends those studies by investigating the long-run behavior of seven daily spot rates from 1974 to 1992 using univariate and multivariate persistence measures. Our results indicate that in certain periods, and for some currencies, the long-run behavior deviates from a pure random walk. For most currencies, the mean-reverting component has increased over time. The long-run impact of an unexpected change in one currency on another reflects both the impact of the ERM and the increase in coordinated intervention after the Louvre Accord.

The stochastic properties of nominal foreign exchange rates have been the subject of much research interest. Mussa (1979) states as his first empirical regularity that spot rates approximately follow a random walk. Theoretical support for this conjecture is found in the general asset market theory of exchange rate determination.¹ Empirical support for the pure random walk behavior of exchange rates is found in Giddy and Dufey (1975), Cornell (1977) and Meese and Rogoff (1983), among others. These studies find that a pure random walk model is an adequate description of exchange rate series which cannot be improved upon by more complex time series models such as higher-order ARIMA or exponential smoothing models. Meese and Rogoff (1983) even find that the forecasting accuracy of a pure random walk model is at least as good as the accuracy of structural models which allow for a variety of macroeconomic determinants of exchange rates.

Other studies, however, emphasize that while spot rates *approximately* follow a random walk, they may not do so *exactly*. For example, Baillie and McMahon (1989) find evidence of small but significant serial correlations in weekly spot rate series in the 1970's. Meese and Rogoff (1983) contend that exchange rates probably do not follow an exact random walk, and that sampling errors may be the reason why more complicated time series models do not outperform the simple random walk in fit and/or forecasting accuracy. According to Hakkio (1986) and Ahking and Miller (1987), a sufficient condition for nominal exchange rates to follow a random walk under the asset market approach is that the market fundamentals which determine the exchange rates are also random walks. However, market fundamentals do not follow a random walk in general (Ahking and Miller, 1987). In addition, Levich (1979) points out that activities of the central banks can have a decisive

effect on the efficiency of the foreign exchange market.

Recently, a lot of attention has been devoted to whether nominal exchange rates are stationary; i.e. whether or not they have a unit root. Meese and Singleton (1982) show that weekly spot rates in the late 1970's are non-stationary. For daily spot rates in the 1980's, Baillie and Bollerslev (1989), Hakkio and Rush (1989) and Barnhart and Szakmary (1991) also report strong evidence of non-stationarity.²

In this paper, we are concerned not only with the *presence* of a unit root, but also with its *quantitative importance*. For example, the presence of a unit root implies that part of an unexpected change in the exchange rate is permanent and continues to influence the future levels of the exchange rate, not only in the near future but also in the long run. Regular unit root tests do not assess the magnitude of this retained portion, which determines by how much the long-run forecast of an exchange rate should be updated after an unexpected event. Similarly, any unit root process can be written as the sum of a random walk (or long-run) and a stationary (or mean-reverting) component. Unit root tests reveal the presence of this long-run component, but do not measure its relative contribution to the overall variability of the exchange rate series. The quantitative importance of the permanent component of a stochastic process is often called its *persistence*, and a number of operationalizations of the persistence concept have been proposed in the economics literature. In this paper, we apply some of these measures to seven daily spot rates (relative to the U.S. dollar) from 1974 to 1992, and assess their stability across different regimes of monetary policies and international agreements on currency coordination.

The potential contributions from our analyses are fourfold. First, our analyses can

provide additional insights into the appropriateness of alternative specifications that have been used to test the efficiency of foreign exchange markets. Indeed, Cochrane (1988, 1991) argues that if the retained portion of an unexpected change turns out to be small (or similarly, if the variability in the long-run component is limited), asymptotic theory based on stationarity may be more appropriate than theories based on unit roots. Hence, even though a unit root is present, a model in the levels should then be used rather than a model in the differences.

Second, univariate persistence measures quantify how much the long-run behavior of an exchange rate deviates from that of the random walk benchmark. For example, the best long-run forecast of a random walk process is its current value, and an unexpected one-unit increase in the current level results in a one-unit increase of the long-run forecast; hence, 100% of the initial shock is retained. For other unit root processes, the magnitude of this retained portion may only be 90%, and even though the exchange rate will then be close to a random walk, its long-run forecasts will be different. This deviation suggests the existence of market inefficiencies that affect the long-run evolution of that exchange rate. As such, our analyses add a long-run dimension to previous empirical studies (see e.g. Giddy and Dufey, 1975; Cornell, 1977; Meese and Rogoff, 1983) which only considered short-term deviations between a pure and an approximate random walk model.

Third, we assess the long-run impact of unexpected changes in the seven exchange rates both in a univariate and a multivariate framework. Univariate estimates use only the information contained in the series' own past, whereas multivariate calculations also include the historical patterns of other currencies as well as their interrelationships. Hence,

multivariate persistence calculations should result in more efficient estimates of the long-run impact of an unexpected change in the own currency (Pesaran, Pierse and Lee, 1993). However, they also allow us to quantify the long-run impact on one currency (e.g. the German mark) of a shock to another currency (e.g. the French franc). Our multivariate estimates add a long-run dimension to previous work by Madura and Nosari (1984) and Soenen (1988), who investigated the instantaneous pairwise correlations between multiple currencies. The persistence approach to study long-run interrelationships also complements some of the recent cointegration analyses of the foreign exchange market (see e.g. Baillie and Bollerslev, 1989; Hakkio and Rush, 1989; MacDonald and Taylor, 1991).

Finally, prior studies have suggested that countries may affect nominal exchange rates through their monetary policies and through central bank intervention in the foreign exchange market. The total sample period is divided into three subperiods, each characterized by a distinct shift in U.S. monetary policy and renewed coordinated central bank intervention across countries. We assess the stability of the persistence estimates across those different time periods. As such, some first insights are obtained into the sensitivity of both the long-run properties of exchange rates and their interrelationships to certain major changes in monetary policy and coordinated foreign exchange market intervention.³

The remainder of the paper is structured as follows. Section I introduces the univariate and multivariate persistence measures used in the empirical part of the paper. The data set is described in Section II. Results are provided in Section III and conclusions drawn in Section IV.

I. Methodology

I.A. Univariate persistence measures

The effect of an unexpected change in a time series is said to be "persistent" if it continues to influence the level of the series for an indefinite period into the future. Two commonly used persistence measures are Campbell and Mankiw's (1987a,b) parametric $A(1)$ measure and Cochrane's (1988) nonparametric V measure. The $A(1)$ measure describes what portion of an unexpected one-unit shock to a series affects its long-run level. For example, assume the German mark (S_{GM}) is non-stationary and follows a random walk process given by:

$$(1) \quad S_{GM,t} = S_{GM,t-1} + u_t ,$$

where u_t is white noise. The corresponding infinite-shock representation is

$$(2) \quad S_{GM,t} = u_t + u_{t-1} + u_{t-2} + u_{t-3} + \dots .$$

It is clear that in this case 100% of a shock at time $t-k$ influences all future values of S_{GM} .

Alternatively, suppose the over-time behavior of the French franc (S_{FF}) follows the

ARIMA(0,1,1) process

$$(3) \quad (1-L)S_{FF,t} = \Delta S_{FF,t} = (1-0.4L)u_t ,$$

where Δ is the difference operator and L the lag operator ($L^j S_t = S_{t-j}$). The infinite-shock representation for this process is given by

$$(4) \quad \begin{aligned} S_{FF,t} &= S_{FF,t-1} + u_t - 0.4u_{t-1} \\ &= u_t + 0.6u_{t-1} + 0.6u_{t-2} + 0.6u_{t-3} + \dots . \end{aligned}$$

In this non-stationary process, only 60% of the initial shock affects the future values of S_{FF} .

Thus, even though both S_{OM} and S_{FF} have a unit root, only S_{OM} follows an exact random walk process, and the long-run impact of a random shock to this series is almost twice as large as it is for S_{FF} . For more complex data-generating processes, the derivation of the infinite-shock representation of the series may become quite complicated. Campbell and Mankiw (1987a,b), however, developed a simple procedure to derive the magnitude of the retained fraction from an ARMA representation of the first-differenced series. Consider the following univariate ARIMA representation for an exchange rate S_t :

$$(5) \quad \Phi(L) \Delta S_t = \Theta(L) u_t ,$$

where $\Phi(L)$ and $\Theta(L)$ are finite order polynomials in L . Equation (5) can also be written in its infinite-shock representation:

$$(6) \quad \Delta S_t = \frac{\Theta(L)}{\Phi(L)} u_t = A(L) u_t = (1 + A_1 L + A_2 L^2 + A_3 L^3 + \dots) u_t .$$

With this representation, the impact of a unit unexpected change in period t on the *growth rate* in period $t+k$ (i.e. on ΔS_{t+k}) is A_k . It can be shown that the total impact of the shock on the *level* of the exchange rate in $t+k$ (i.e. S_{t+k}) equals $1 + A_1 + A_2 + \dots + A_k$ (Campbell and Mankiw, 1987a). Thus, the long-run impact of the shock on the level of the exchange rate is equal to the sum of the moving-average coefficients in the infinite-shock representation of the first-differenced series, often denoted as $A(1)$. As can easily be seen in equation (6), this sum can also be estimated as the ratio $\Theta(1)/\Phi(1)$, where $\Theta(1)$ ($\Phi(1)$) is the sum of the moving-average (autoregressive) coefficients of the ARMA representation of ΔS_t .

For a stationary series, $A(1)$ equals zero: an unexpected change at time t is only a temporary deviation from the series' trend or mean, and will not affect the long-run

expectation.⁴ For a random walk, $A(1)$ equals one.⁵ This result reflects the fact that the best long-run forecast of a random walk is its current value. Therefore, a current increase of one unit causes a one-unit increase in the long-run forecast. For series which are neither stationary nor a pure random walk, $A(1)$ can take on any other value greater than zero. If $A(1)$ is greater than one, an unexpected increase today will typically be reinforced by other positive changes in the future, and the series will continue to diverge from its pre-shock expectation. Series with a persistence measure between zero and one show a mean-reverting tendency in that they try to return to the pre-shock expected level. However, they never get there completely, even in the far future, because of the permanent component in their behavior.

A somewhat different but, as will be illustrated below, related persistence operationalization has been proposed by Cochrane (1988). Any unit root series can be written as the sum of a random walk plus a stationary process, in which the random walk carries the permanent part of the series, and the stationary process carries the temporary part.⁶ The ratio of the variance in the random walk part to the variability in the one-period changes reflects the relative importance of the random walk, or permanent, component to the behavior of the series.⁷ This variance ratio, denoted as V , can be shown to equal the limiting value (i.e. for k going to infinity) of

$$(7) \quad V_k = \frac{1}{k+1} \frac{\text{var}(S_{t+k+1} - S_t)}{\text{var}(S_{t+1} - S_t)}.$$

V can be approximated by choosing a finite k value. The building blocks of equation (7) are (1) the variance of the long differences, which capture the long-run movements in the series'

behavior, and (2) the variance of the one-period changes. For a stationary series, the variance of the long differences does not depend on k and V_k approaches zero for large k . For a random walk, the variance of the $(k+1)$ lagged difference is $(k+1)$ times the variance of the once-lagged difference; thus, V_k will equal one. V_k can also be written as a function of the series' autocorrelation coefficients:

$$(8) \quad V_k = 1 + 2 \sum_{j=1}^k \left(1 - \frac{j}{k+1}\right) \rho_j ,$$

where ρ_j is the j th autocorrelation coefficient of ΔS_t . This measure is purely data driven, and does not require any model specification (e.g. no ARIMA model is assumed as in Campbell and Mankiw's procedure). One can therefore refer to V_k as a nonparametric persistence operationalization. Although $A(1)$ and V_k have the same value for a random walk (i.e. one) and for a stationary process (i.e. zero), this equality does not hold in general. Still, there is a mathematical relationship between the two measures (Campbell and Mankiw, 1987b). In the limit, as k approaches infinity,

$$(9) \quad A(1) = \sqrt{V / (1 - R^2)} ,$$

where V is the limiting value of V_k , and R^2 is the proportion of the variance of ΔS_t explained by equation (5). Thus, $[A(1)]^2$ is bounded from below by V , and the more predictable the differenced process, the greater the disparity between the two measures. Equation (9) can be used to obtain a nonparametric estimate of Campbell and Mankiw's $A(1)$ measure.

Univariate persistence measures have been used to analyze the long-run behavior of macro-economic variables (e.g. Campbell and Mankiw, 1987a,b; Cochrane, 1988; Jaeger and Kunst 1990), and to measure the long-run components in real exchange rates (e.g. Huizinga,

1987). The variance ratio given in equation (7) is also used by Liu and He (1991) in a test statistic to determine the significance of low-order correlations in the increments of five weekly nominal exchange rates. In their analyses, k is only allowed to vary between two and sixteen, and hence, only short-run deviations from a pure random walk process are considered. Following Huizinga (1987), our persistence calculations focus on the long-run movements of the nominal exchange rates (as captured by the variance of their long differences), which is indicated by Liu and He (1991, p. 783) as an important area for future research.

Finally, the long-run dimension of our analyses is also evident from the fact that Cochrane's V measure corresponds to a simple normalization of the spectral density at zero frequency.⁸ Similarly, the multivariate extension of Cochrane's V measure, which we use to capture the long-run co-movements between exchange rates, corresponds to a normalization of the cross-spectrum at zero frequency.

I.B. Estimation of univariate persistence

A parametric estimate of $A(1)$ for a particular exchange rate can be obtained by fitting low-order ARMA models to its first difference (see equation (5)). $\hat{A}(1)$ is then equal to the ratio of the sum of the moving average estimates to the sum of the autoregressive estimates:

$$(10) \quad \hat{A}(1) = \hat{\Theta}(1) / \hat{\Phi}(1) .$$

Each combination of p and q (the order of the autoregressive and moving average polynomials) provides a slightly different approximation to the infinite shock representation in equation (6), and thus will give a somewhat different estimate of $A(1)$. Rather than selecting one specific ARMA(p,q) model, we assess the robustness of the $A(1)$ estimates to

different values of p and q .⁹

For $p > 0$, $A(1)$ is a nonlinear function of $B = (\phi_1, \phi_2, \dots, \phi_p, \theta_1, \theta_2, \dots, \theta_q)$ and its approximate variance can be computed using the delta method (Jaeger and Kunst, 1990):

$$(11) \quad Var[A(1)] \approx \left[\frac{\partial f(B)}{\partial B} \right]' Var(B) \left[\frac{\partial f(B)}{\partial B} \right] .$$

To estimate V_k , the population autocorrelations in equation (8) are replaced with the sample autocorrelations $(\hat{\rho}_j)$. $\hat{V}_k$ can also be computed recursively using the following difference equation (Pesaran and Samiei, 1991)

$$(12) \quad \hat{V}_k = \frac{2k}{k+1} \hat{V}_{k-1} - \frac{k-1}{k+1} \hat{V}_{k-2} + \frac{2}{k+1} \hat{\rho}_k , \quad k=3,4,\dots$$

with initial values

$$(13) \quad \hat{V}_1 = 1 + \hat{\rho}_1 \quad \text{and} \quad \hat{V}_2 = 1 + 4/3 \hat{\rho}_1 + 2/3 \hat{\rho}_2 .$$

The standard error of $\hat{V}_k$ is given by (Campbell and Mankiw, 1987b)

$$(14) \quad SE(\hat{V}_k) = \hat{V}_k / \sqrt{3/4 \ T/(k+1)} ,$$

where T is the sample size. When implementing Cochrane's procedure, one must decide on the value of k , i.e. the number of autocorrelations in equation (8). Ideally, a value of k that minimizes some criterion function (e.g. that results in a minimum mean squared error) should be selected. However, this requires knowledge of the true underlying process for the different exchange rates, knowledge which would make the entire persistence analysis redundant (Huizinga, 1987). In practice, k should be large enough to minimize the bias resulting from excluding higher-order autocorrelations, but should not approach the sample

size T either. Indeed, $\hat{V}_k$ is identically zero for $k=T-1$ (Campbell and Mankiw, 1987b). Studies typically report $\hat{V}_k$ for several values of k ,¹⁰ and (1) assess the robustness of the findings to the choice of k , and/or (2) compare the obtained values with the ones that would have been obtained if the process followed a pure random walk process. While, as discussed above, V equals one for a random walk process, the *estimated* V_k have been shown to converge to $(T-k)/T$ rather than to one. Put differently, $\hat{V}_k$ is a consistent estimate of V , but is subject to a small sample bias.

A consistent nonparametric estimate of $A(1)$ for window size k is given by

$$(15) \quad \hat{A}_k(1) = \sqrt{\hat{V}_k / (1 - \hat{\rho}_1^2)} ,$$

where $\hat{\rho}_1^2$ is used as a conservative estimate of R^2 .

Section III reports the results of these univariate persistence measures for seven daily spot rates. A drawback of univariate persistence measures is that their information set is limited to the series' own realized values. Taking the German mark as an example, $\hat{A}(1)$ and $\hat{V}_k$ use only the past values of S_{GM} . However, useful information may also be contained in the present and past values of other currencies. For example, Soenen (1988) finds evidence of contemporaneous relationships between different currencies, and a number of studies have identified the existence of a long-run or cointegrating relationship between exchange rates (see e.g. Baillie and Bollerslev, 1989; MacDonald and Taylor, 1991).

When also incorporating the information contained in other currencies, more efficient estimates of the long-run impact of a shock in the currency itself may be obtained.

However, the multivariate calculations also provide an estimate of the impact of the shock in a given currency on the long-run value of another currency. The persistence approach to studying long-run interrelationships between exchange rates complements previous work by Baillie and Bollerslev (1989) and MacDonald and Taylor (1991), who find evidence of a long-run relationship between, respectively, all the currencies in our sample, and the three ERM currencies in our sample. These studies, however, only identify the *existence* of a long-run equilibrium relationship; they do not describe its *nature*. In this respect, one can either study the coefficients of the cointegrating vector, or one can estimate the multivariate persistence measures, which may be more readily interpretable than the cointegrating coefficients (see e.g. Dekimpe, 1992; Lütkepohl and Reimers, 1992).

I.C. Multivariate extension of $A(1)$

Consider the following vector moving average (VMA) representation of a system of two non-stationary currencies, S_{OM} and S_{FF} , where for ease of exposition we have omitted all deterministic components:

$$(16) \quad \begin{bmatrix} \Delta S_{GM,t} \\ \Delta S_{FF,t} \end{bmatrix} = \begin{bmatrix} 1 & 0 \\ 0 & 1 \end{bmatrix} \begin{bmatrix} u_{GM,t} \\ u_{FF,t} \end{bmatrix} + \begin{bmatrix} a_{11}^1 & a_{12}^1 \\ a_{21}^1 & a_{22}^1 \end{bmatrix} \begin{bmatrix} u_{GM,t-1} \\ u_{FF,t-1} \end{bmatrix} + \dots ,$$

where a_{12}^k indicates the impact on $\Delta S_{OM,t}$ of a one-unit unexpected change in $S_{FF,t-k}$. Using a similar logic as in the univariate case, the long-run impact of a unit shock in $S_{FF,t-k}$ on the level of $S_{OM,t}$ is $(0 + a_{12}^1 + a_{12}^2 + a_{12}^3 + \dots + a_{12}^k)$. This sum can be interpreted as a multivariate extension of Campbell and Mankiw's $A(1)$ measure. A deficiency of this method is that the contemporaneous effect on ΔS_{GM} of a shock in S_{FF} is zero, which does not conform with the empirical findings in Soenen (1988). One solution to this problem,

implemented by Dekimpe (1992) and Evans (1989), is to use a Cholesky decomposition to transform the original specification into a system with uncorrelated errors. The resulting model specification is recursive, and allows for instantaneous effects in one direction. A disadvantage, however, is that the use of the Cholesky decomposition requires the imposition of a causal ordering between the exchange rates. For example, we must be able to say that an unexpected change in $S_{GM,t}$ has an impact on $S_{FF,t}$, but that a shock in $S_{FF,t}$ does not have a contemporaneous effect on $S_{GM,t}$ (or vice versa). However, there is no theory that would support this type of causal ordering between currencies, and contemporaneous effects may well exist in both directions. Because of these problems with the multivariate version of Campbell and Mankiw's $A(1)$, we use a multivariate extension of Cochrane's V measure in the empirical section of the paper.

I.D. Multivariate extension of V_k

Recall that V_k measures the relative importance of a series' permanent or long-run component by the ratio of the variance of the random walk component to the variance of the one-period changes. The multivariate version uses the variance-covariance matrix of the random walk components of all the currencies in the system. The diagonal elements of this variance-covariance matrix are the variances of the long differences for each currency (see equation (7)), and the off-diagonal elements are a measure of association between the currencies' long-run movements. To obtain a scale-free measure for the persistence of fluctuations in $S_{GM,t}$ caused by a unit shock in $S_{FF,t}$, the corresponding element in the variance-covariance matrix can be normalized by the unconditional variance of ΔS_{FF} .

Pesaran et al. (1993) show that the variance-covariance matrix of the respective

random walk components, denoted as $VC(\tau_j)$, equals¹¹

$$(17) \quad VC(\tau_j) = A(1) \Sigma A(1)' ,$$

where $A(1)$ is the sum of the moving-average components in a vector moving average model,¹² and Σ is the residual variance-covariance matrix. Contemporaneous effects are already reflected in Σ , and hence no causal ordering is needed in this methodology. The Appendix contains the procedure for estimating the asymptotic standard errors.

Pesaran et al. (1993) show that the existence of a cointegrating relationship between different currencies imposes certain restrictions on the associated persistence measures. Indeed, irrespective of the origin of a shock to the system, the underlying cointegrating relationship should be preserved. In terms of a bivariate model, if S_{OM} and S_{FF} are cointegrated, then they should not diverge from each other after a shock to either spot rate. Pesaran et al. show that if r cointegrating relationships exist between M series, the persistence matrix $P = [P_{ij}]$, where P_{ij} is the long-run impact of a unit shock to j on the level of i ($i, j = 1, 2, \dots, M$), estimated from an unrestricted model in the differences will be of rank $M-r$, reflecting the constraints imposed by the cointegrating relationship.

II. Data section

The data set consists of the natural log of daily spot rates for seven currencies expressed per U.S. dollar: the Canadian dollar, French franc, German mark, Italian lire, Japanese yen, Swiss franc and British pound. The spot bid rates were collected every business day at noon by the Federal Reserve System. Our sample period starts on 3 September 1974 and ends on 27 May 1992, excludes banking holidays and has 4,446 observations per series.¹³

Prior studies suggest that countries can influence nominal exchange rates through their

monetary policies (Humpage, 1991) and that the stochastic properties of exchange rates may depend upon both monetary policies and economic forces (e.g. McFarland, Pettit and Sung, 1982; Akgiray and Booth, 1988). There is also evidence that coordinated central bank intervention in the foreign exchange market sometimes may have an impact on exchange rates that is significantly different from intervention by one central bank alone (e.g. Belongia, 1992; Dominguez and Kenen, 1992). To get some insights into whether the long-run properties of exchange rates are also affected by changes in monetary policies and increased coordinated intervention, we divided the sample into three subperiods: (1) from 3 September 1974 to 9 October 1979, (2) from 10 October 1979 to 19 September 1985, and (3) from 20 September 1985 to 27 May 1992. The first subperiod starts in 1974, when most of the countries shifted to a flexible exchange rate regime. In October 1979, the starting point for the second period, the Federal Reserve Bank adopted a new anti-inflationary program based on a monetary policy that controlled M1 instead of interest rates. This date has been associated with a structural change and a shift in the relation between economic variables and therefore is a common cut-off point for examining economic time series over several periods (e.g. Liu and He, 1991). The third subperiod starts with the Plaza Accord, where the Group of Ten signed an international cooperative agreement to control the volatility of world currency markets and to establish currency target zones.

III. Empirical results

III.A. Unit root tests

Before estimating the persistence measures, we tested whether the seven spot rates are non-stationary using standard Dickey-Fuller (DF) (1981) and Augmented Dickey-Fuller (ADF)

tests. Five tests were run on both the entire sample and the three subperiods: (1) without drift or trend, (2) with drift, (3) with drift and trend, (4) with day-of-the week dummy variables, and (5) with day-of-the-week dummy variables and trend. Day-of-the-week dummy variables were included to capture possible seasonal effects in daily spot rates.¹⁴ The ADF tests for each of the five cases included 1, 6, 12, 24, 36 and 48 lags. In most instances, the results were not affected by the number of lagged differences in the ADF tests. In those few cases where this did affect our results, we used conventional significance tests on the additional coefficients to determine the need to augment an initially more restricted specification. A similar practice is found in Evans (1989), Perron (1990), and Dekimpe (1992).

For the entire period, the null hypothesis of a unit root could never be rejected. The trend, drift and seasonal dummy variables were not significant, and therefore we relied on the results of the first test. Allowing for these deterministic components did not change our conclusions, however. In the subperiods, the drift was significant for some exchange rates, in which case the results of the tests which allowed for a drift or seasonal dummy variables were used. In the first and second subperiods, the trend was never significant and the unit root null hypothesis could never be rejected.

The test results for the third subperiod, on the other hand, indicated that only the Canadian dollar and the pound had a unit root. The absence of a unit root in the other currencies was due to the time interval at the beginning of the third period, from 20 September 1985 (the Plaza Accord) to 18 February 1987 (the Louvre Accord), during which the value of the U.S. dollar declined sharply and the trend was significant. When this time

interval was excluded from the third subperiod, all spot and forward rates were again found to have a unit root. In what follows, the third subperiod is restricted to start on 19 February 1987.

We also tested for the presence of a second unit root, which was rejected in all periods.¹⁵ Overall, our results concur with those reported in Baillie and Bollerslev (1989), Barnhart and Szakmary (1991) and Copeland (1991).¹⁶

III.B. Univariate V_k estimates

Figure 1 presents the plots of the $\hat{V}_k$ values for the pound spot rate for lags one to half of the sample size and for the random walk benchmark using data from, respectively, the entire period, and the first, second and third subperiods. The values of interest are for k ranging from $T/3$ to $T/2$, where T is the sample size.¹⁷ The random walk benchmark, as discussed in Section I, is $(T-k)/T$.

Please insert Figure 1 about here

One recurring pattern in the $\hat{V}_k$ plots is that they are hump shaped: they first rise above one, and then fall.¹⁸ Huizinga (1987), who found a similar pattern for real exchange rates, argues that the hump-shaped pattern is due to positive serial correlations at low lags which are more than fully offset by negative serial correlations at higher lags. The value of k at which $\hat{V}_k$ starts to fall, however, is not a reliable indicator of when the autocorrelations of ΔS_t first become negative since the weights of the autocorrelations change as k increases (see equation (8)). Similarly, the point where $\hat{V}_k$ first falls below one is not a reliable indicator of the minimum time horizon in which the spot rate exhibits mean-reverting

behavior.¹⁹

While all plots are hump-shaped, not all fall below the random walk benchmark. Panel A of Figure 1 indicates that for the entire period (September 1974 to May 1992), the $\hat{V}_k$ approach the random walk benchmark. In the first subperiod (September 1974 to October 1979), on the other hand, the $\hat{V}_k$ are always substantially larger than one, as is apparent from panel B. Hence, unexpected changes were typically reinforced by other positive changes and the exchange rate continued to diverge from its pre-shock expectation. Panel C indicates that during the second subperiod (October 1979 to September 1985), the permanent component becomes less pronounced, and the $\hat{V}_k$ are very close to the random walk benchmark. Finally, panel D shows that in the third subperiod (February 1987 to May 1992), $\hat{V}_k$ falls below $(T-k)/T$, indicating the presence of a mean reverting component.²⁰ The graphs of the nonparametric $A(1)$ estimates for the pound follow the same general pattern.²¹

The graphs of the $\hat{V}_k$ for the French franc, the mark, the lire, the yen and the Swiss franc (available upon request) all exhibit a similar pattern, i.e. they indicate that their mean-reverting component becomes stronger in the second subperiod and even stronger in the third subperiod.²² Still, there are also some differences among the respective $\hat{V}_k$. For the mark and the Swiss franc, the $\hat{V}_k$ in the first subperiod are close to the random walk benchmark. For the French franc, the lire and the yen, on the other hand, both the entire period and the first subperiod have a stronger permanent component than a random walk process. An exception to this pattern of increasing mean reversion over time is the Canadian dollar. Its persistence estimates are above the random walk benchmark for the entire period and for the first and third subperiods. Its mean-reverting component is most pronounced in the second

subperiod.

In summary, although the unit root tests indicate that all spot rates are non-stationary, the results of the univariate persistence estimates illustrate that the series' long-run properties vary across spot rates and across time. There is evidence in Figure 1 that, especially in certain periods and for certain currencies, the *long-run* behavior of exchange rates deviates from a random walk. These results extend prior studies which test whether the *short-run* behavior of exchange rates follows a random walk (e.g. Giddy and Dufey, 1975; Cornell, 1977; Meese and Rogoff, 1983; Liu and He, 1991). The presence of a stronger mean-reverting component in the third subperiod is consistent with Baillie and Humpage (1992), who investigate the efficacy of post-Louvre Accord intervention, and find that the Federal Reserve and the German and Japanese central banks seemed to intervene in defense of specific targets for the mark and the yen relative to the U.S. dollar. However, it seems that the intervention could not reverse movements in errant exchange rates, and that the central banks would set a new target if deviations from the old target persisted. Furthermore, the central banks have cut back their intervention since the post-Louvre Accord period. These observations by Baillie and Humpage (1992) are consistent with our results in that even though the mark and yen exhibit a strong mean-reverting component in the third subperiod, there still is a permanent component in their behavior; i.e. there is no complete mean reversion.

III.C. Univariate $A(1)$ estimates

Table 1 presents the $A(1)$ estimates for the pound spot rate.²³ For each period, a total of 30 ARIMA($p,1,q$) models (i.e. for $p=0,1,\dots,5$ and for $q=1,\dots,5$) were estimated.²⁴

Maximum-likelihood estimation procedures which explicitly allow for roots on the unit circle were used. We also obtained persistence estimates for seasonally adjusted data by analyzing the residuals from a prior regression of ΔS_t on day-of-the-week dummy variables. The results were very similar to Table 1.

Please insert Table 1 about here

Our results indicate that the $\hat{A}(1)$ are quite robust with respect to the choice of p and q . A vast majority of the estimates lies between 1.03 and 1.11 for both the entire period and all three subperiods. The few exceptions all correspond to models whose moving-average polynomials have a root close to one, and which are therefore characterized by both unstable estimates and large standard errors. Overall, there is little evidence of overdifferencing, and the general picture that arises from Table 1 is an $\hat{A}(1)$ which is very close to one.

These ARIMA-based $\hat{A}(1)$ are quite different from both the $\hat{V}_k$ and the nonparametric $\hat{A}(1)$. This discrepancy between the different measures is not uncommon.²⁵ A possible explanation is that low-order ARIMA models basically capture the short-run properties of the data-generating process, which are then extrapolated to infer the series' long-run properties. Obviously, if the true long-run properties differ from the short-term characteristics of the process, misleading results may be obtained. The ARIMA models, where the maximum p and q equal 5, ignore high-order autocorrelations, which may or may not be significant. The V_k estimates, on the other hand, use all the autocorrelations up to lag k (which may be very large). A large number of small high-order autocorrelations can offset a few smaller-order autocorrelations, which may explain the difference between the two persistence

operationalizations. For the pound, for example, the $\hat{V}_k$ graphs suggest there is a long sequence of small positive autocorrelations in the first subperiod (panel B of Figure 1) and a long sequence of small negative autocorrelations in the third subperiod (panel D). However, none of the lower-order autocorrelations are significant, and all ARIMA models, as well as the $\hat{A}(1)$ derived from them, suggest a random walk process (see panels B and D in Table 1).

In summary, different results are obtained for the $A(1)$ and V_k measures. For some currencies, the $\hat{V}_k$ indicate a clear deviation from random walk behavior, but this is not true for the $\hat{A}(1)$. These differences are basically due to their different treatment of the insignificant higher order autocorrelation coefficients. This issue is discussed in more detail in the following Section which compares multivariate persistence estimates based on, respectively, a restricted and unrestricted VAR model.

III.D. Multivariate persistence estimates

Following Lee et al. (1992) and Pesaran et al. (1993), we derived the multivariate persistence estimates from both restricted and unrestricted VAR models. The selected orders for the VAR models were 2 and 10. The $VC(\tau)$ were calculated according to equation (17). The restricted VAR models were constructed by dropping all regressors whose coefficients had a t -ratio of less than one in absolute value. For the unrestricted models, Ordinary Least Squares (OLS) was used since no efficiency gain could be obtained by using a systems approach. For the restricted specifications, an iterative Seemingly Unrelated Regressions (SUR) technique was used.

Cochrane (1988) advocates using high-order autocorrelation terms in persistence

estimations *even if* they are insignificant. Pesaran et al. (1993), on the other hand, argue that the use of unrestricted models leads to poorly determined persistence measures, and that estimates based on a restricted specification are both more efficient and more reliable. The persistence estimates obtained from the restricted VAR(2) models, presented in Table 2, are similar to those obtained from the unrestricted VAR(2) specification. However, as expected, the asymptotic standard errors are smaller for the estimates based on the more parsimonious model. The diagonal elements are the multivariate version of Cochrane's V_k measure: they indicate the long-run impact of an unexpected change in the own spot rate, taking into account all possible channels of influence, including the effect of contemporaneous shocks in other currencies that are induced by the shock of interest. The off-diagonal element (i,j) is an estimate of the effect of a shock in spot rate j on the long-run value of spot rate i .

Please insert Table 2 about here

With respect to the diagonal elements in Table 2, the univariate pattern of increased mean reversion in the third subperiod is not always observed in the multivariate estimates. For most currencies (e.g. French franc, mark, Swiss franc), the 95% confidence intervals for the diagonal elements include one in all subperiods. These findings are more in line with the univariate $A(1)$ estimates than with some of the univariate V_k estimates. This is not surprising, since a VAR(2) model also captures the lower-order auto- and cross-correlations.

The following findings emerged from the VAR(10) models, which are available from the authors upon request. The estimates derived from the *unrestricted* VAR(10) specification are larger than those from the VAR(2) model, which suggests that the multivariate results

may exhibit a similar hump-shaped pattern as the univariate results.²⁶ However, the associated standard errors for the multivariate estimates became quite large. For example, the persistence estimate of the lire based on the restricted VAR(2) model for the entire period is 1.081 with standard error equal to 0.020; the corresponding estimate based on the unrestricted VAR(10) model is 1.273 with standard error 0.083. Overall, the estimates from the *restricted* VAR(10) approach were very similar (1.101 with standard error of 0.061 for the lire) to the ones obtained for the VAR(2) specification. As a consequence, depending on the treatment of the insignificant effects, somewhat different findings emerge. So far, no consensus has emerged in the econometrics literature on which approach is preferred. Following Pesaran et al. (1993), we use the restricted specifications to discuss the off-diagonal elements. We focus on the estimates of the restricted VAR(2) model, which are very similar to those of the restricted VAR(10) model.

The following observations can be made with respect to these off-diagonal elements. First, they tend to be smaller than the diagonal elements in the same column, and their 95 % confidence intervals usually lie below one (some exceptions can nevertheless be observed, such as the estimates describing the long-run impact on the Swiss franc, which are sometimes larger than one). This implies that, on average, a shock has a larger impact on the own long-run evolution than on the long-run value of the other currencies.

Second, the Canadian dollar is affected less by shocks to any of the other currencies. For example, the long-run impact of a shock originating in the other currencies on the value of the Canadian spot rate ranges for the entire period between 0.054 and 0.102 (row 1 of panel A), while the long-run impact on the French franc of shocks originating in the other

currencies ranges between 0.652 and 0.924 (row 2). Also, in the first and third subperiods, the impact of a shock originating in the Canadian dollar on the long-run value of the other currencies is very small (column 1 in panels B and D). For the second subperiod, on the other hand, these shocks have a much larger impact than those originating in, for example, the mark or the French franc (see columns 1, 2 and 3 in panel C). These results again illustrate the distinct behavior of the Canadian dollar, which was already evident from our univariate results. More research is needed to explain this unique behavior.

Third, the off-diagonal elements are not symmetric. For example, for the entire period, the long-run impact of a shock to the Swiss franc on the French franc 0.726, while the long-run impact of a shock to the French franc on the Swiss franc is 1.012. This asymmetry is caused by differences in the applied normalization factor, which in turn reflect differences in the volatility of the respective exchange rates.

Fourth, the more coordinated the currencies and the stronger the economic links between their respective countries, the larger the corresponding off-diagonal persistence estimates. The persistence estimates between European currencies are typically larger than the persistence estimates of a shock in either the yen or the Canadian dollar on European currencies (or vice versa). In the third subperiod, for example, the long-run impact on the European currencies of a shock in the yen ranges between 0.700 and 0.816, while the impact on the yen of shocks to the European currencies ranges between 0.631 and 0.753. Among the European currencies, whose economies and currencies have stronger ties to each other than to Japan, the persistence estimates are between 0.805 and 1.131.

Of the five European currencies in our sample, three participate in the ERM: the

mark, French Franc and lire. The U.K. participates in the ERM only as of 1990. Currencies in the ERM are allowed to fluctuate freely within a band around bilateral central rates. Once the currencies fluctuate outside this band, the central banks are committed to intervene to bring the exchange rate again within the band. For the mark and the French franc, this band is rather narrow, and has been fixed at 2.25% above and below the central rate. We therefore expect that any shock will have a similar long-run effect on these two currencies. This is consistent with the over-time pattern of the long-run impact of a shock in the mark on, respectively, the mark itself, the french franc, and the lire: in the first subperiod, before the ERM was established (March 1979), the long-run impact of such a shock on the mark (0.992) was quite different from the long-run impact on the other two currencies (0.681 and 0.435, respectively). In the second subperiod, these numbers are closer together, and by the third subperiod, they are very close (1.058, 1.010, 0.956). The same pattern is observed for a shock originating in the French franc and the lire. The fluctuation bands for the lire (until 1990) were set at 6% above and below the central rate, and this greater latitude is reflected also in the persistence estimates: row four in panel C (which measures the impact of different shocks on the lire) differs more from row 3 (the impact on the mark) than the latter does from row 2 (the impact on the French franc). The importance of the ERM for the long-run behavior of the different currencies was also illustrated by MacDonald and Taylor (1991), who found a unique cointegrating vector between the three ERM currencies, but not between three non-ERM currencies (Canadian dollar, yen and pound) using monthly observations from 1979 to 1988.

As a last observation from Table 2, note that the off-diagonal estimates between the

members of the Group of Seven, except for the Canadian dollar, are larger in the third than in the second subperiod. Also, note that the differences between rows 2, 3, and 4 are much smaller in the third than in the second subperiod. This result is consistent with the increased coordinated intervention between these countries.

IV. Summary and conclusions

This paper uses univariate and multivariate persistence measures to investigate the long-run properties of seven daily nominal spot rates relative to the U.S. dollar from 1974 to 1992. Univariate persistence estimates measure the quantitative importance of the unit roots that are typically found in exchange rate series. Two measures that quantify the long-run impact of an unexpected change in the current value of the exchange rate are discussed: Campbell and Mankiw's (1987a,b) $A(1)$ and Cochrane's (1988) V measure. The latter measure indicates that, especially for certain time periods, the long-run behavior of some currencies deviates from the behavior implied by a pure random walk model. Our results therefore add a long-run dimension to previous research that tested whether the short-run behavior of exchange rates is well-approximated by a random walk model. For most currencies and in most time periods, the long-run component turns out to be substantial, suggesting the appropriateness of using a model in the differences when testing the efficiency of the foreign exchange market. We do find, however, that the mean-reverting component of most currencies has increased over time. The presence of a stronger mean reversion after the 1987 Louvre Accord is consistent with the increased *and* coordinated intervention in the foreign exchange market in that period.

The multivariate calculations, based on Pesaran et al. (1993), allow us to quantify the

long-run impact on a currency of a shock to either the currency itself or to any of the other currencies. The results suggest that the long-run impact of shocks originating in another currency is similar for the ERM currencies. Furthermore, the cross-effects also reflect certain changes in U.S. monetary policy, and the increase in coordinated intervention after the Louvre Accord. This persistence approach to studying long-run interrelationships between currencies extends previous cointegration analyses which tested for the existence of long-run equilibrium relationships.

Several areas for future research remain open, however. From a statistical point of view, more research is needed on the most appropriate treatment of the insignificant higher-order autocorrelations and/or cross-correlations. For those unit root processes where a substantial mean-reverting component is present, as is the case in our third subperiod, guidelines are needed on *when* the theory based on stationarity becomes more appropriate than unit root theories. From a finance point of view, more work needs to be done to determine the underlying causes of the observed long-run effects and the shifts in these effects over time. This paper suggests some intuitive explanations for the observed patterns. More explicit conclusions could be drawn by incorporating explanatory factors, such as changes in market fundamentals, into the model. Such a model is beyond the scope of this paper, and we refer to Lee et al. (1992) for an illustration of the modelling issues involved. Another potential extension of the paper is to include not only spot rates but also forward rates in the VAR model. If the forward rate is an unbiased predictor of the future spot rate, the long-run impact of any shock should be the same on a currency's spot rate as on its forward rate. Baillie and Bollerslev (1989) test whether the cointegrating vector between

spot and forward rates equals $[0,1]$ in a bivariate framework. The multivariate persistence methodology provides an alternative testing procedure for this hypothesis, and can easily incorporate several currencies simultaneously.

Appendix

This appendix presents the derivation of an expression for the asymptotic variance of the multivariate persistence estimates. The results are presented for an unrestricted VAR model. Similar expressions can be derived for a restricted model,²⁷ and can be obtained from the authors upon request.

Consider the following vector-autoregressive model

$$(A.1) \quad \Pi(L) \Delta \vec{S}_t = \vec{u}_t ,$$

where $\Pi(L)$ is equal to $(I_M - \Pi_1 - \dots - \Pi_p)$, I_M is an identity matrix of order M , Π_i is an $(M \times M)$ matrix of unknown coefficients, and Σ is the variance-covariance matrix of the residuals. First, we derive the asymptotic variance of the elements $VC(\tau_i)$ (see equation 17). For ease of notation, we denote $VC(\tau_i)$ by Q :

$$(A.2) \quad Q = A(1) \Sigma A(1)' ,$$

where $A(1) = [\Pi(1)]^{-1}$. Note that the elements of Q (i.e. Q_{ij} ; $i, j = 1, \dots, M$) correspond to the unscaled persistence estimates. As indicated in the text, the scaled or normalized persistence estimates (indicated by P_{ij}) are obtained by dividing the elements in the j th column of Q by $\text{Var}(\Delta S_{j,t})$. Hence, standard errors for P_{ij} are given by the ratio of the standard error of Q_{ij} to the variance of $\Delta S_{j,t}$. Stacking all observations, we can rewrite equation (A.1) as

$$(A.3) \quad \Pi(L) \Delta SS' = U' ,$$

where $\Delta SS'$ is equal to the $(M \times T)$ matrix $[\Delta \vec{S}_1, \dots, \Delta \vec{S}_T]$, and U' is equal to the $(M \times T)$ matrix $[\vec{u}_1, \dots, \vec{u}_T]$. After writing equation (A.3) in vector form, we obtain

$$(A.4) \quad \text{Vec}(\Delta SS) = (I_M \otimes \Delta SS_0) \vec{\Theta} + \text{Vec}(U) ,$$

where ΔSS_0 is equal to the $(TxMP)$ matrix $[\Delta SS_{.1}, \dots, \Delta SS_{.p}]$, and where $\vec{\Theta}$ is equal to $\text{Vec}([\Pi_1, \dots, \Pi_p]')$. For an unrestricted VAR model, $\vec{\Theta}$ is of order PM^2 . All elements of Q are scalar functions of $\vec{\Theta}$, and we denote this functional relationship by $Q_{ij}(\vec{\Theta})$. Hence, if the variance-covariance matrix of $\vec{\Theta}$ is known, one can apply the delta method to derive the asymptotic variance of $Q_{ij}(\vec{\Theta})$. In matrix notation:

$$(A.5) \quad \text{Avar}\{\text{Vec}[Q(\vec{\Theta})]\} = \frac{\partial \text{Vec}[Q(\vec{\Theta})]}{\partial \vec{\Theta}'} \text{Avar}(\vec{\Theta}) \left[\frac{\partial \text{Vec}[Q(\vec{\Theta})]}{\partial \vec{\Theta}'} \right]'$$

The variances of interest (i.e. $\text{Var}[Q_{ij}(\vec{\Theta})]$) are given by the corresponding diagonal elements of the $(M^2 \times M^2)$ matrix $\text{Avar}\{\text{Vec}[Q_{ij}(\vec{\Theta})]\}$. After some algebra,²⁸ it can be shown that²⁹

$$(A.6) \quad \text{Avar}(\vec{\Theta}) = S \otimes [(\Delta SS_0)'(\Delta SS_0)]^{-1} ,$$

and

$$(A.7) \quad \begin{aligned} \frac{\partial \text{Vec}[Q(\vec{\Theta})]}{\partial \vec{\Theta}'} = & - [A(1) \otimes A(1)] * [2/(T-MP)] * [I_M \otimes U'(\Delta SS_0)] \\ & + [I_M \otimes A(1)\Sigma] * [A(1) \otimes \{\Pi(1)\}^{-1}] * [I_M \otimes S_p] \\ & - [A(1)\Sigma \otimes I_M] * [\{\Pi(1)\}^{-1'} \otimes \{\Pi(1)\}^{-1}] * \left[\frac{\partial \text{Vec}\Pi(1)}{\partial \vec{\Theta}'} \right] , \end{aligned}$$

$$(A.8) \quad \text{where } S = \frac{U'U}{T-MP} , \text{ and } S_p = [I_M, I_M, \dots, I_M] .$$

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Notes

- * We acknowledge Cynthia Cummings for computer programming assistance. An earlier version of this paper was presented at the 1992 Financial Management Association meeting.
- 1. According to this theory, which assumes no government intervention, exchange rates are relative prices of assets determined in organized markets under rational expectations (see e.g. Mussa, 1979). These prices are closely linked to the market's expectations of the future price of the asset, and new information that changes these expectations is assumed to be reflected immediately in the current price. Since new information arrives randomly, exchange rates move in an unpredictable manner. Put differently, deviations from expected values are assumed to be completely random, and the foreign exchange market is said to be efficient. Indeed, if significant serial correlations were present in exchange rate changes, market participants would be able to predict a significant portion of these changes, and could earn high profits with relatively little risk.
- 2. An exception is the study by Whitt (1992), which is discussed in somewhat more detail in a following section.
- 3. Sephton and Larsen (1991) do not explicitly consider changes in monetary policy, but provide some indirect evidence for the fact that the long-run interrelationships between currencies may have changed over time. Indeed, they show that the presence of a long-run equilibrium between the currencies in our sample depend on the time frame used in the cointegration tests.

4. Suppose $\Phi(L)S_t = \Theta(L)u_t$, where S_t is stationary. To apply equation (6), take the first difference of S_t : $\Phi(L)(1-L)S_t = \Theta(L)(1-L)u_t = \Theta^*(L)u_t$. Hence,
 $\Delta S_t = \{\Theta^*(L)/\Phi(L)\} u_t$. But since $\Theta^*(1)=0$, it follows that $A(1) = \Theta^*(1)/\Phi(1) = 0$.
5. If S_t follows a random walk, $\Delta S_t = u_t$. By equation (6), $A(L)=1$ and thus $A(1)=1$.
6. See Stock and Watson (1988) for a mathematical proof.
7. It can be shown that even though this decomposition in a random walk and stationary part is not unique, the variance ratio of interest is not affected by the selected decomposition (Cochrane, 1988).
8. See Dekimpe (1992) or Pesaran et al. (1993) for a more detailed description of this relationship.
9. Jaeger and Kunst (1990) and Pesaran and Samiei (1991) adopt a similar approach.
10. See e.g. Campbell and Mankiw (1987a,b), Huizinga (1987), Cochrane (1988), Pesaran and Samiei (1991), and Dekimpe (1992).
11. It can be shown that this measure is directly related to the cross-spectrum at zero frequency (Pesaran et al., 1993).
12. $A(1)$ can also be estimated consistently by estimating a VAR model, computing the sum of the corresponding autoregressive coefficients $(I - \Pi_1 - \Pi_2 - \dots - \Pi_p)$ and inverting the resulting matrix (Lee et al., 1992; Pesaran et al., 1993).

13. Following Copeland (1991), banking holidays (182 days) are excluded from the sample. Only the lire had a missing observation; all other spot rates were complete. This one missing observation was estimated as the average of the lire spot rates preceding and following the missing observation.
14. The need to test for day-of-the-week effects in exchange rate data has been documented by, among others, McFarland, Pettit and Sung (1982), and Baillie and Bollerslev (1989).
15. To test for the presence of a second unit root, we follow Pesaran and Samiei (1991) and replace the first differences (ΔS_t) in the DF and ADF tests with the second differences ($\Delta^2 S_t$). Pesaran and Samiei point out that it may not be entirely appropriate to run the tests on first as well as on second differences, but that this may not be a serious problem if the test statistics are well below their critical values, as was the case in our analyses. The results of all unit root tests are available from the authors upon request.
16. Baillie and Bollerslev (1989) apply the Phillips-Perron (1988) tests which allow for time-dependent heteroskedasticity. Barnhart and Szakmary (1991) use the sequential Dickey-Pantula (1987) scheme, which first tests for three, then two and finally one unit root. Copeland (1991) uses the same DF and ADF tests as we did. Recently, Whitt (1992) adopted a Bayesian approach to test for unit roots in five monthly spot rates, and found evidence of stationarity in some of them. We adopt the more frequently used classical approach for the following reasons. First, the results of Bayesian unit root tests have been shown to be very sensitive to the choice of the

prior distribution. This was already evident from Whitt's discussion, where depending on the prior weight attached to stationarity, different results were found for the French franc. Also in the special issue of the *Journal of Applied Econometrics* on Bayesian unit root tests (October-December 1991), this sensitivity to the choice of priors was repeatedly deplored (see e.g. Phillips, 1991; Stock, 1991), and it was shown that for several macroeconomic time series similar conclusions were obtained as with the classical approaches provided that other priors were used. Second, Phillips (1991) showed that the Sims (1988) test used by Whitt is characterized by a substantial bias towards finding stationarity, which may explain Whitt's conclusions. Indeed, it was shown that in a time series context, flat priors do *not* reflect ignorance about the potential parameter values, but downweight the unit root and explosive alternatives.

17. See Campbell and Mankiw (1987a,b).
18. Using weekly nominal exchange rates, Liu and He (1991) observe the onset of this hump-shaped pattern since their variance ratio typically increases when k increases from two to sixteen.
19. See Huizinga (1987) for a detailed discussion on these issues.
20. As can easily be seen in equation (14), the estimated standard errors tend to become quite large: keeping V_k constant, the standard errors increase as k increases. When interpreting these graphs, attention is therefore not focused on the significance of an individual V_k estimate, but on (1) the overall pattern of successive V_k estimates, as

well as on (2) the similarities/differences of these patterns across different currencies and time periods. A similar approach is adopted by Huizinga (1987), who emphasizes the "information content" of long strings of (albeit small) autocorrelations with the same sign, and of the similarity of these strings across real exchange rates.

21. Because of space constraints, these graphs are not included but are available from the authors upon request.
22. Given that all spot rates are expressed per U.S. dollar, their behavior may be determined primarily by unexpected U.S. events, which may explain why the pattern of persistence estimates is similar for most currencies.
23. We restrict ourselves to the pound to illustrate Campbell and Mankiw's $A(1)$ measure. The estimation procedure and the derivation of the associated standard errors is computationally involved, and a complete analysis of all series would require the estimation of 7 (the number of series) $\times$ 4 (the number of periods) $\times$ 30 (the number of models) ARIMA models. Moreover, for reasons discussed before (i.e. the need to impose a causal ordering on the spot rates when using a multivariate extension of the $A(1)$ measure), we emphasize the multivariate extension of the V measure in our subsequent analyses.
24. The order of the moving average component was always at least one to be able to detect overdifferencing.
25. Cochrane (1988), for example, found no evidence of a large random walk component in GNP, while Campbell and Mankiw (1987a,b) reported parametric $A(1)$ values

larger or equal to one. Dekimpe (1992) and Pesaran and Samiei (1991), on the other hand, obtained more robust persistence estimates: their parametric and nonparametric $A(1)$ estimates are almost identical.

26. One could of course increase the number of lags even further, in which case we expect to see a similar decrease in the estimates, and maybe also an increased mean reversion in the third subperiod, as in the univariate case. However, these models would be severely over-parameterized (each additional lag results in the estimation of 49 extra coefficients, and the use of 200 lags would require the estimation of 9800 coefficients). As a consequence, almost all estimates may well have a t -statistic of less than one, in which case restricted higher-order models would not change much. Also, computer memory constraints precluded the computation of standard errors for unrestricted models of higher order.
27. For example, in the restricted specifications of the multivariate estimates, we restrict all coefficients with a t -statistic of less than one in absolute value to zero.
28. Useful suggestions from Kevin Lee and Richard Pierse are greatly appreciated.
29. These expressions differ from the formulas given in Pesaran et al. (1993), which only considers the variance of the diagonal elements of the persistence matrix, and which uses a different normalization factor. Indeed, our normalization factor (i.e. $\text{Var}(\Delta S_{j,t})$) does not depend on $\vec{\Theta}$. Pesaran et al., on the other hand, use the conditional variance as normalization factor, which depends on $\vec{\Theta}$ and which therefore affects the partial derivatives. As indicated in Pesaran et al., both

normalization factors can be used, but our normalization has the appealing property that it results in a straightforward generalization of the univariate measure proposed by Cochrane (1988).

Table 1: Univariate $A(1)$ Estimates for the British Pound Spot Rate

Panel A: Entire period (3 September 1974 - 27 May 1992; 4446 observations).

	q=1	q=2	q=3	q=4	q=5
p=0	1.067 (0.015)	1.063 (0.022)	1.053 (0.027)	1.040 (0.031)	1.086 (0.035)
p=1	1.064 (0.021)	1.057 (0.031)	1.050 (0.028)	1.041 (0.028)	1.081 (0.038)
p=2	1.054 (0.023)	1.066 (0.023)	1.056 (0.016)	1.057 (0.022)	1.071 (0.030)
p=3	1.049 (0.028)	1.058 (0.018)	1.043 (0.011)	1.269* (0.172)	1.060 (0.025)
p=4	1.039 (0.028)	1.044 (0.029)	1.049 (0.019)	1.372* (0.327)	1.372* (0.266)
p=5	1.082 (0.039)	1.072 (0.032)	1.061 (0.025)	1.372* (0.299)	NA

Panel B: First subperiod (3 September 1974 - 9 October 1979; 1278 observations).

	q=1	q=2	q=3	q=4	q=5
p=0	1.041 (0.028)	1.054 (0.040)	NA	1.025 (0.059)	1.065 (0.065)
p=1	1.057 (0.056)	1.633* (0.562)	1.072 (0.042)	1.052 (0.052)	1.585* (0.428)
p=2	1.057 (0.051)	1.627* (0.534)	1.064 (0.043)	1.047 (0.064)	1.593* (0.298)
p=3	1.627* (0.555)	1.021* (0.031)	1.017 (0.014)	1.056 (0.035)	1.539* (0.381)
p=4	1.051 (0.053)	1.046 (0.064)	1.055 (0.036)	1.657* (0.584)	1.642* (0.550)
p=5	1.590* (0.447)	1.600* (1.015)	1.065 (0.046)	1.644* (0.534)	NA

Table 1 continued

Panel C: Second subperiod (10 October 1979 - 19 September 1985; 1490 observations).

	q=1	q=2	q=3	q=4	q=5
p=0	1.045 (0.026)	1.075 (0.037)	1.078 (0.047)	1.069 (0.054)	1.109 (0.061)
p=1	1.089 (0.069)	1.077 (0.040)	1.524* (1.081)	1.503* (1.010)	1.509* (0.946)
p=2	1.080 (0.042)	1.037 (0.023)	1.059 (0.027)	1.071 (0.026)	1.093 (0.057)
p=3	1.554* (1.279)	1.077 (0.027)	1.089 (0.038)	NA	1.519* (0.839)
p=4	1.086 (0.123)	1.086 (0.043)	1.520* (0.841)	1.101 (0.060)	1.063 (0.033)
p=5	1.094 (0.060)	1.118 (0.035)	1.091 (0.053)	1.072 (0.033)	1.157 (0.094)

Panel D: Third subperiod (19 February 1987 - 27 May 1992; 1768 observations).

	q=1	q=2	q=3	q=4	q=5
p=0	1.096 (0.027)	1.065 (0.041)	1.017 (0.050)	1.032 (0.057)	1.086 (0.064)
p=1	1.084 (0.036)	1.035 (0.069)	1.021 (0.053)	1.212* (0.221)	1.081 (0.067)
p=2	1.029 (0.054)	1.034 (0.022)	1.064 (0.031)	1.056 (0.043)	1.017 (0.052)
p=3	1.017 (0.047)	1.064 (0.034)	1.059 (0.037)	1.197* (0.296)	1.205* (0.284)
p=4	1.072 (0.077)	1.022 (0.046)	1.048 (0.044)	1.055 (0.032)	1.080 (0.048)
p=5	1.085 (0.071)	1.041 (0.051)	1.211* (0.244)	1.042 (0.037)	NA

Note: Standard errors are reported in parentheses. NA indicates that the estimates did not converge within 50 iterations. * indicates that the sum of the moving-average coefficients was close to zero. Spot rates are taken from the International Finance data tape of the Federal Reserve Bank.

Table 2: Multivariate Persistence Estimates Using a Restricted VAR(2) Model

Panel A: Entire period (3 September 1974 - 27 May 1992; 4446 observations).

Impact on	Source of the shock						
	CD	FF	GM	IL	JY	SF	BP
CD	1.124 (0.034)	0.092 (0.006)	0.091 (0.005)	0.096 (0.007)	0.054 (0.004)	0.076 (0.004)	0.102 (0.007)
FF	0.639 (0.036)	1.017 (0.018)	0.924 (0.012)	0.960 (0.017)	0.652 (0.015)	0.726 (0.012)	0.745 (0.015)
GM	0.652 (0.036)	0.953 (0.012)	1.015 (0.010)	0.960 (0.014)	0.684 (0.015)	0.785 (0.010)	0.768 (0.014)
IL	0.595 (0.040)	0.858 (0.016)	0.832 (0.012)	1.081 (0.020)	0.582 (0.015)	0.647 (0.012)	0.677 (0.017)
JY	0.337 (0.024)	0.584 (0.014)	0.593 (0.014)	0.582 (0.015)	1.055 (0.026)	0.511 (0.014)	0.497 (0.012)
SF	0.729 (0.039)	1.012 (0.016)	1.061 (0.014)	1.009 (0.020)	0.796 (0.021)	1.035 (0.018)	0.811 (0.018)
BP	0.707 (0.045)	0.747 (0.015)	0.747 (0.014)	0.760 (0.018)	0.557 (0.013)	0.584 (0.013)	1.086 (0.027)

Panel B: First subperiod (3 September 1974 - 9 October 1979; 1278 observations).

Impact on	Source of the shock						
	CD	FF	GM	IL	JY	SF	BP
CD	1.216 (0.077)	0.005 (0.014)	0.019 (0.015)	-0.017 (0.027)	-0.097 (0.021)	0.005 (0.010)	0.018 (0.015)
FF	0.026 (0.062)	0.964 (0.046)	0.681 (0.035)	0.545 (0.029)	0.450 (0.028)	0.410 (0.025)	0.437 (0.033)
GM	0.103 (0.080)	0.784 (0.040)	0.992 (0.042)	0.508 (0.035)	0.506 (0.025)	0.547 (0.028)	0.455 (0.039)
IL	-0.080 (0.125)	0.537 (0.030)	0.435 (0.030)	1.454 (0.087)	0.320 (0.017)	0.299 (0.022)	0.353 (0.034)
JY	-0.477 (0.101)	0.469 (0.028)	0.458 (0.022)	0.338 (0.016)	1.022 (0.011)	0.300 (0.016)	0.256 (0.027)
SF	0.059 (0.119)	1.003 (0.061)	1.164 (0.060)	0.742 (0.056)	0.705 (0.039)	1.010 (0.057)	0.558 (0.059)
BP	0.091 (0.069)	0.460 (0.035)	0.416 (0.036)	0.377 (0.035)	0.259 (0.029)	0.240 (0.025)	0.979 (0.030)

Table 2 continued

Panel C: Second subperiod (10 October 1979 - 19 September 1985; 1490 observations).

Impact on	Source of the shock						
	CD	FF	GM	IL	JY	SF	BP
CD	1.136 (0.063)	0.176 (0.016)	0.187 (0.016)	0.191 (0.017)	0.147 (0.019)	0.171 (0.015)	0.166 (0.013)
FF	1.520 (0.137)	1.006 (0.054)	0.973 (0.052)	1.007 (0.055)	0.613 (0.048)	0.830 (0.047)	0.752 (0.046)
GM	1.518 (0.130)	0.915 (0.048)	1.039 (0.051)	1.001 (0.052)	0.640 (0.048)	0.879 (0.047)	0.760 (0.044)
IL	1.342 (0.116)	0.818 (0.044)	0.866 (0.045)	0.962 (0.049)	0.543 (0.041)	0.734 (0.041)	0.639 (0.040)
JY	0.999 (0.128)	0.480 (0.037)	0.534 (0.040)	0.524 (0.040)	1.032 (0.055)	0.531 (0.038)	0.415 (0.035)
SF	1.685 (0.142)	0.945 (0.054)	1.066 (0.057)	1.029 (0.057)	0.771 (0.055)	1.082 (0.056)	0.814 (0.051)
BP	1.419 (0.105)	0.743 (0.046)	0.800 (0.047)	0.778 (0.049)	0.523 (0.044)	0.707 (0.044)	1.057 (0.048)

Panel D: Third subperiod (19 February 1987 - 27 May 1992; 1768 observations).

Impact on	Source of the shock						
	CD	FF	GM	IL	JY	SF	BP
CD	0.980 (0.006)	0.033 (0.016)	0.029 (0.016)	0.037 (0.017)	0.010 (0.011)	0.021 (0.013)	0.045 (0.014)
FF	0.205 (0.103)	1.068 (0.067)	1.010 (0.064)	1.090 (0.069)	0.732 (0.057)	0.846 (0.057)	0.905 (0.058)
GM	0.198 (0.107)	1.104 (0.070)	1.058 (0.067)	1.131 (0.071)	0.761 (0.059)	0.883 (0.059)	0.939 (0.060)
IL	0.209 (0.097)	1.008 (0.063)	0.956 (0.061)	1.069 (0.066)	0.700 (0.055)	0.805 (0.054)	0.867 (0.056)
JY	0.062 (0.070)	0.728 (0.056)	0.692 (0.053)	0.753 (0.058)	1.100 (0.064)	0.631 (0.047)	0.697 (0.050)
SF	0.165 (0.108)	1.088 (0.073)	1.038 (0.069)	1.118 (0.075)	0.816 (0.061)	1.007 (0.065)	0.965 (0.064)
BP	0.297 (0.094)	0.977 (0.062)	0.927 (0.059)	1.012 (0.064)	0.757 (0.055)	0.810 (0.054)	1.138 (0.069)

Note: Standard errors are reported in parentheses. Spot rates for the Canadian dollar (CD), French franc (FF), German mark (GM), Italian lire (IL), Japanese yen (JY), Swiss franc (SF) and British pound (BP) are taken from the International Finance data tape of the Federal Reserve Bank.

Figure 1: Univariate V_k Estimates for the British Pound Spot Rate

Panel A: Entire period (3 September 1974 - 27 May 1992; 4446 observations).

Panel B: First subperiod (3 September 1974 - 9 October 1979; 1278 observations).

Figure 1 continued

Panel C: Second subperiod (10 October 1979 - 19 September 1985; 1490 observations).

Panel D: Third subperiod (19 February 1987 - 27 May 1992; 1678 observations).

Note: V_k estimates are sampled every 30th observation for the entire period (panel A) and every 10th observation for the subperiods (panels B, C, D). The random walk benchmark is $(T-k)/T$. Daily spot rates are taken from the International Finance data tape of the Federal Reserve Bank.