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Persistence in foreign exchange rates

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This study examines the long-run behavior of seven daily nominal exchange rates using univariate and multivariate persistence measures. Our results indicate that for some currencies, the long-run behavior deviates from that of a pure random walk in certain periods. The multivariate estimates reflect the effect of both the EMS Exchange Rate Mechanism and increased post-Louvre Accord coordinated intervention. A large portion of the effect of a shock in one currency on another currency's long-run value can be attributed to contemporaneous effects on the other currencies. (JEL F31). Copyright © 1996 Elsevier Science Ltd

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. 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

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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 1970s. 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).

Pure random-walk behavior of exchange rates requires two properties: (1) the presence of a unit root, and (2) uncorrelated increments (Liu and He, 1991). Over the last few years, a lot of attention has been devoted to whether nominal exchange rates are non-stationary, i.e. whether they have a unit root. Meese and Singleton (1982) show that weekly spot rates in the late 1970s are non-stationary. For daily spot rates in the 1980s, Baillie and Bollerslev (1989a) and Copeland (1991) also report strong evidence of non-stationarity.[1] In terms of the (absence of) serial correlation in the increments, a number of recent studies have applied variance-ratio tests to examine the statistical significance of the low-order or short-run correlations (see e.g. Liu and He, 1991). It is, however, equally important to examine potential longer-run deviations from pure random-walk behavior, as has been emphasized in Booth, Kaen and Koveos (1982), Liu and He (1991) and Pan, Liu and Bastin (1992). To address this issue, we consider not only the *presence* of a unit root but also its *quantitative importance*, and compare the values obtained for several daily exchange rate series with the theoretical values expected when they would follow a pure random walk process.

The presence of a unit root implies, for example, 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. The magnitude of this retained portion is one for a random walk, and values different from one indicate long-run deviations from pure random-walk behavior. 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. Again, a random walk benchmark will be presented to assess deviations from pure random-walk behavior in an exchange rate's long-run fluctuations.

The quantitative importance of the permanent component in a time series is often called its *persistence*, and the aforementioned measures of the concept have recently been formalized in the economics literature. In this paper, we apply these measures to seven daily spot rates (relative to the US dollar) from 1974 to 1992, and assess their stability across different regimes of monetary policies and international agreements on currency coordination.

The contributions from our analyses are fivefold. First, we already indicated how our work adds a long-run dimension to previous empirical studies (see e.g. Giddy and Dufey, 1975; Cornell, 1977; Liu and He, 1991; Meese and Rogoff, 1983) which only considered short-term deviations between a pure and an approximate random-walk model.

Second, our analyses can provide additional insights into the appropriateness of alternative mathematical specifications (*i.e.* in the levels or in the differences) that have been used to test the efficiency of foreign exchange markets (see *e.g.* Baillie and Bollerslev, 1989a, for a review) or the validity of different models of exchange rate determination (*e.g.* Boothe and Glassman, 1987). 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 rather than a model in the differences should then be used.

Third, we assess the long-run effect 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 effect of an unexpected change in the own currency (Pesaran *et al.*, 1993). In addition, they also allow us to quantify the long-run effect on one currency of a shock to another currency. Our multivariate estimates therefore 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, and by Cornell (1977) and MacDonald and Taylor (1991) who used VAR models to capture lagged (temporary) effects of one currency on another.

Fourth, the persistence approach to study long-run interrelationships complements some of the recent cointegration analyses of the foreign exchange market (see e.g. Baillie and Bollerslev, 1989a; Hakkio and Rush, 1989; MacDonald and Taylor, 1991; Sephton and Larsen, 1991). These studies assessed the *existence* of a long-run equilibrium relationship between a number of currencies, but did not yet assess the *nature* of these relationships. As shown in Lütkepohl and Reimers (1992), it is often difficult to interpret cointegration relationships directly. In this paper, we apply the full-information maximum-likelihood approach of Johansen and Juselius (1990), and contrast our findings with previous studies. We will argue that, if cointegrating relationships are found, persistence estimates not only reflect the underlying long-run equilibria (and therefore provide an alternative testing procedure for many prevailing theories), but are also more readily interpretable.

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 (see e.g. Humpage, 1991; McFarland *et al.*, 1982). The sample period is divided into three subperiods, each characterized by a distinct shift in US 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.[2]

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 (1987) A(1) measure and Cochrane's (1988) V measure.

The A(1) measure describes what portion of an unexpected one-unit shock to a series affects its long-run level, and can be derived from the ARMA representation of the first-differenced series:

$$\langle 1\rangle \qquad \Phi(L)\Delta S_t = \Theta(L)u_t,$$

where $\Phi(L) = (1 - \phi_1 L - \ldots - \phi_p L^p)$ and $\Theta(L) = (1 - \theta_1 L - \ldots - \theta_q L^q)$ are polynomials in the lag operator L. Written in infinite-shock form, equation ⟨1⟩ becomes:

$$\langle 2\rangle \qquad \Delta S_t = \frac{\Theta(L)}{\Phi(L)}u_t = A(L)u_t = \left(1 + A_1 L + A_2 L^2 + A_3 L^3 + \ldots\right)u_t.$$

With this representation, the effect 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, and the total effect of the shock on the *level* of the exchange rate in $t + k$(*i.e.* on S_{t+k}) can be shown to equal $1 + A_1 + A_2 + \ldots + A_k$ (see *e.g.* Campbell and Mankiw, 1987). The long-run effect (*i.e.* for $k \to \infty$) $1 + A_1 + A_2 + \ldots$ is denoted as $A(1)$, and it is easily seen from equation ⟨2⟩ that it can be estimated as the ratio of the sum of the moving-average coefficients to the sum of the autoregressive coefficients in the ARMA model of the first-differenced series, *i.e.*

$$\langle 3\rangle \qquad \hat{A}(1) = \frac{\hat{\Theta}(1)}{\hat{\Phi}(1)} = \frac{1 - \hat{\theta}_1 - \ldots - \hat{\theta}_q}{1 - \hat{\phi}_1 - \ldots - \hat{\phi}_p}.$$

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 its long-run expectation. For a random walk, $A(1)$ equals one. This reflects the fact that the best long-run forecast of a random walk is its current value. For series which are neither stationary nor a pure random walk, $A(1)$ can take on any value greater than zero. If $A(1)$ is greater than one, an unexpected increase 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 between zero and one show a mean-reverting tendency in that they try to return to their pre-shock expected level. However, they never get there completely, even in the far future, because of the permanent component in their behavior. In sum, the extent to which an exchange rate's $A(1)$ estimate differs from one indicates how much its long-run forecast deviates from the value predicted by a pure random-walk process.[3]

A somewhat different but 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, where the random walk carries the permanent part of the series and the stationary process carries the temporary part. 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. This variance ratio, denoted as V, can be shown to equal the limiting value (*i.e.* for k going to infinity) of[4]

$$\langle 4\rangle \qquad V_k = \frac{1}{k+1}\frac{var(S_{t+k+1} - S_t)}{var(S_{t+1} - S_t)}.$$

V can be approximated by choosing a finite k value. The building blocks of equation $\langle 4\rangle$ 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, the random-walk benchmark for V_k equals one.[5] V_k can also be written as a function of the series' autocorrelation coefficients:

$$\langle 5\rangle \qquad V_k = 1 + 2\sum_{j=1}^{k}\left(1 - \frac{j}{k+1}\right)\rho_j,$$

where ρ_j is the jth autocorrelation coefficient of ΔS_t. An estimate for V_k is obtained by replacing the ρ_j by the sample autocorrelations. When implementing Cochrane's procedure, one must decide on the value of k, *i.e.* the number of autocorrelations in equation $\langle 5\rangle$.[6] k should be large enough to minimize the bias resulting from excluding higher-order autocorrelations, but should not approach the sample size T. Indeed, $\hat{V}_k$ is identically zero for $k = T - 1$. Studies typically report $\hat{V}_k$ for several values of k, and 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 one. Put differently, $\hat{V}_k$ is a consistent estimate of V, but is subject to small sample bias.

The long-run nature of the two persistence operationalizations is also evident from the fact that both correspond to a simple normalization of the spectral density at zero frequency. As such, the persistence concept is not just a convenient way of summarizing the autocorrelation structure in the data: it quantifies the relative importance of the very long-run movements which correspond to frequency zero (see Pesaran *et al.*, 1993 for a more elaborate discussion on this frequency domain interpretation).

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. However, useful information may also be contained in the present and past values of other currencies. When also incorporating this information, more efficient estimates of the long-run effect of a shock in the currency itself may be obtained. Moreover, multivariate calculations also provide an estimate of the effect of a shock in one currency on the long-run value of another currency.

I.B. Multivariate extension of A(1)

To allow for a wide range of interdependencies between the currencies, consider the following vector moving average (VMA) model, where for ease of exposition we have omitted all deterministic components:

$$\langle 6 \rangle \qquad \Delta S_t = \mathbf{A}(L)\boldsymbol{u}_t,$$

where ΔS_t denotes a $(M \times 1)$ vector of exchange rate changes $\Delta S_{i,t}$, and $\boldsymbol{u}_t$ is a $(M \times 1)$ vector of white noise innovations with mean zero and variance-covariance matrix Σ. The matrix polynomial

$$\langle 7 \rangle \qquad \mathbf{A}(L) = \sum_{k=0}^{\infty} \mathbf{A}_k L^k$$

is assumed to be absolutely summable, the $\mathbf{A}_k$ are $(M \times M)$ matrices of unknown coefficients, and $\mathbf{A}_0$ is an identity matrix. The (i, j)th element of $\mathbf{A}_k$ (a_{ij}^k) indicates the effect on $\Delta S_{i,t}$ of a one-unit unexpected change in the jth currency k periods ago, and instantaneous effects are reflected in Σ. As such, the proposed model is flexible enough to capture the different channels of influence through which an unexpected shock can affect the future values of the respective currencies: lagged effects in the own currency, lagged cross-country effects and instantaneous cross-country reactions. Contemporaneous effects, for example, could occur as a result of a closely coordinated central bank intervention in the foreign exchange market (*e.g.* among ERM countries or, in the post-Louvre Accord period, between Germany, Japan and the USA). Baillie and Humpage (1992) provide support for the inclusion of lagged effects:

they find that after the Louvre Accord, central banks often intervened for a number of successive periods (*i.e.* as long as the spot rate continued to move away from a specific target rate). In terms of the lagged cross-effect, a policy switch by one country may affect expectations about the monetary policy in another country. MacDonald and Taylor (1991), for example, found that the German money supply often leads the Italian and French money supply. This type of policy spillover suggests that a model used to quantify the total long-run effect of unexpected events should allow for both contemporaneous and lagged cross-effects, as our VMA model does.

Using a similar logic as in the univariate case, the long-run effect of a unit shock in $S_{j,t-k}$ on the *level* of $S_{i,t}$ is $(0 + a_{ij}^1 + a_{ij}^2 + a_{ij}^3 + ... + a_{ij}^k)$. This sum can be interpreted as a multivariate extension of Campbell and Mankiw's $A(1)$ measure, and the matrix which contains these sums for all (i,j) pairs will be labeled $\mathbf{A}(1)$. A deficiency of this method is that the contemporaneous impact on ΔS_i of a shock in S_j is not incorporated (since $a_{ij}^0 = 0$), which does not conform with earlier empirical findings. As such, this persistence operationalization fails to compute the *total* effect of the event which caused the initial shock.

One solution to this problem, implemented by Dekimpe and Hanssens (1995), 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 the German mark can have an immediate effect on the value of the French franc, but that a shock in the franc cannot have an immediate effect on the value of the mark. 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.C. 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 extension 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 ⟨4⟩), 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_i caused by a unit shock in S_j, the corresponding element in the variance–covariance matrix can be normalized by the unconditional variance of ΔS_j (the currency where the shock originated). As shown in Pesaran *et al.* (1993), the resulting multivariate persistence measure corresponds to a normalized estimate of the cross-spectrum at frequency zero, which again illustrates the long-run nature of the measure.

Pesaran *et al.* (1993) show that the variance–covariance matrix of the respective random walk components, denoted as $\mathbf{VC}(\tau_t)$, equals

$$\langle 8\rangle \qquad \mathbf{VC}(\tau_t) = \mathbf{A}(1)\boldsymbol{\Sigma}\mathbf{A}(1)',$$

where $\mathbf{A}(1)$ contains the sum of the moving-average coefficients in the VMA model (capturing the lagged effects), and where $\boldsymbol{\Sigma}$ is the residual variance–covariance matrix (reflecting the contemporaneous effects). Because of the inclusion of $\boldsymbol{\Sigma}$, no causal ordering is needed in this methodology. To obtain an estimate of $\mathbf{VC}(\tau_t)$, one needs consistent estimates of $\mathbf{A}(1)$ and $\boldsymbol{\Sigma}$. The procedures for estimating these components vary depending on whether or not the series of interest are cointegrated: the MA parameters are obtained directly from the AR representation in the absence of cointegration, and from the corresponding error correction model when cointegrating relationships exist (see Van de Gucht *et al.* (1995) for a discussion of the estimation issues involved). As with the univariate measures, we normalize the persistence estimates by dividing the elements in the jth column of $\mathbf{VC}(\tau_t)$ by the unconditional variance of $\Delta S_{j,t}$. Pesaran *et al.* (1993) focused on the diagonal elements of $\mathbf{VC}(\tau_t)$, and used the conditional variance of $\Delta S_{j,t}$ (*i.e.*, the jth diagonal element of $\boldsymbol{\Sigma}$) to scale the respective estimates. Both scaling factors are valid, but our normalization has the appealing properties that (i) it results in a straightforward generalization of the univariate measure proposed by Cochrane (1988), and (ii) the derivation of the associated standard errors (especially for the off-diagonal elements) is computationally more tractable.[7]

In any case, it is worth emphasizing that the existence of a cointegrating relationship between some of the currencies imposes certain restrictions on the associated persistence measures. Indeed, irrespective of the origin of a shock to the system, the underlying (cointegrating) equilibrium relationship should be preserved in the persistence measures. In terms of a bivariate model, if S_i and S_j are cointegrated, they should not diverge from each other after a shock to either spot rate. More generally, it can be shown that if r cointegrating relationships exist between M series, the persistence matrix $\mathbf{P} = [P_{i,j}]$, where $P_{i,j}$ is the long-run effect of a unit shock to j on the level of i $(i, j = 1, 2, \ldots, M)$, will be of rank $M - r$, reflecting the constraints imposed by the cointegrating relationship. Often (and especially when $r > 1$), the cointegrating vectors are difficult to interpret directly, in which case impulse-response functions and persistence estimates may provide additional insights (Lütkepohl and Reimers, 1992).[8] In Section IV, we indicate how these properties of multivariate persistence estimates can be used in assessing the long-run value of alternative models of exchange rate determination.

II. Data section

The data set consists of the natural log of daily spot rates for seven currencies expressed per US 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.[9]

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 *et al.*, 1982; Akgiray and Booth, 1988). There is also evidence that coordinated central bank intervention in the foreign exchange market may sometimes have an effect 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: (i) from 3 September 1974 to 9 October 1979, (ii) from 10 October 1979 to 19 September 1985, and (iii) 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 is therefore a common cut-off point for examining economic time series over several periods (*e.g.* Liu and He, 1991; Pan *et al.*, 1992). 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 (see *e.g.* Flynn and Boucher, 1993).

The evolution of the logarithm of the British pound and the French franc is plotted in Figure 1. Commonly used summary statistics for all currencies are reported in Table 1.[10] The figures in Table 1 indicate that the distribution of changes in the logarithm of daily spot rates is skewed and kurtotic, especially in the first subperiod. These findings are consistent with the results in Baillie and Bollerslev (1989b), who suggest that exchange rates follow a martingale process with heteroskedastic error terms. Even though Baillie and Bollerslev (1989b) have shown that the degree of skewness and kurtosis diminishes when working with more coarsely sampled data, we believe it is appealing to work with daily as opposed to weekly or monthly data when trying to link the notion of a shock to specific events such as the arrival of new information. As emphasized in Pesaran and Samiei (1991), persistence measures reflect the long-run effect of a shock sustained over the considered data interval. Since news about events such as money-supply announcements or central-bank interventions in the foreign exchange market arrives on a given day rather than piece-by-piece throughout a given week or month, and since foreign exchange markets have been found to react quickly to this information (see *e.g.* Ito and Roley, 1987; Loopesko, 1984), it is conceptually appealing to use daily data to trace their over-time effect.

TABLE 1. Summary statistics of the first difference of log spot rates*100

	CD	FF	DM	IL	JY	SF	BP
Panel A: Entire period (9 / 3 / 1974–5 / 27 / 1992; 4446 observations)							
mean	0.0044	0.0030	−0.0110	0.0139	−0.0190	−0.0159	0.0056
std. dev.	0.2468	0.6491	0.6594	0.6137	0.6140	0.7665	0.6503
skewness (0.0013)[a]	0.3358	0.2035	0.0524	0.0026	−0.4870	0.0383	0.0790
kurtosis (0.0054)[a]	4.6936	6.0043	4.0941	5.0133	4.4082	3.6923	3.8711
Panel B: First subperiod (9 / 3 / 1974–10 / 9 / 1979; 1278 observations)							
mean	0.0133	−0.01026	−0.0307	0.0181	−0.0229	−0.0484	0.0055
std. dev.	0.2169	0.4704	0.5047	0.4671	0.4800	0.7360	0.4826
skewness (0.0047)	0.8176	0.8826	1.4984	1.0061	−0.7917	0.7073	0.6478
kurtosis (0.0188)	7.4951	21.0333	20.6486	24.6344	17.6711	11.3547	12.2309
Panel C: Second subperiod (10 / 10 / 1979–9 / 19 / 1985; 1490 observations)							
mean	0.0110	0.0491	0.0320	0.0572	0.0046	0.0258	0.0319
std. dev.	0.2468	0.7256	0.7036	0.6544	0.6423	0.7741	0.7213
skewness (0.0040)	0.0973	0.2302	−0.3542	−0.2346	−0.5045	−0.2872	−0.2337
kurtosis (0.0161)	3.2948	5.7281	1.6438	3.1320	1.8737	1.1207	2.9371
Panel D: Third subperiod (2 / 19 / 1987–5 / 27 / 1992; 1326 observations)							
mean	−0.0077	−0.0078	−0.0085	−0.0043	−0.0127	−0.0031	−0.0122
std. dev.	0.2673	0.6644	0.6947	0.6389	0.6625	0.7532	0.6904
skewness (0.0045)	0.5056	0.0465	0.0868	0.0453	−0.2686	−0.0663	0.2812
kurtosis (0.0181)	4.6685	1.7891	1.8537	1.6830	2.8311	1.0790	1.7129

Notes: [a]Standard errors reported in parentheses. Under assumptions of normality, the skewness and kurtosis coefficients have asymptotic distributions of $N(0, 6/T)$ and $N(0, 24/T)$, respectively.
Canadian dollar (CD), French franc (FF), German mark (DM), Italian lire (IL), Japanese yen (JY), Swiss franc (SF) and British pound (BP) spot rates are obtained from the Federal Reserve International Finance tape.

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: (i) without drift or trend, (ii) with drift, (iii) with drift and trend, (iv) with day-of-the week dummy variables, and (v) 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.[11] 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

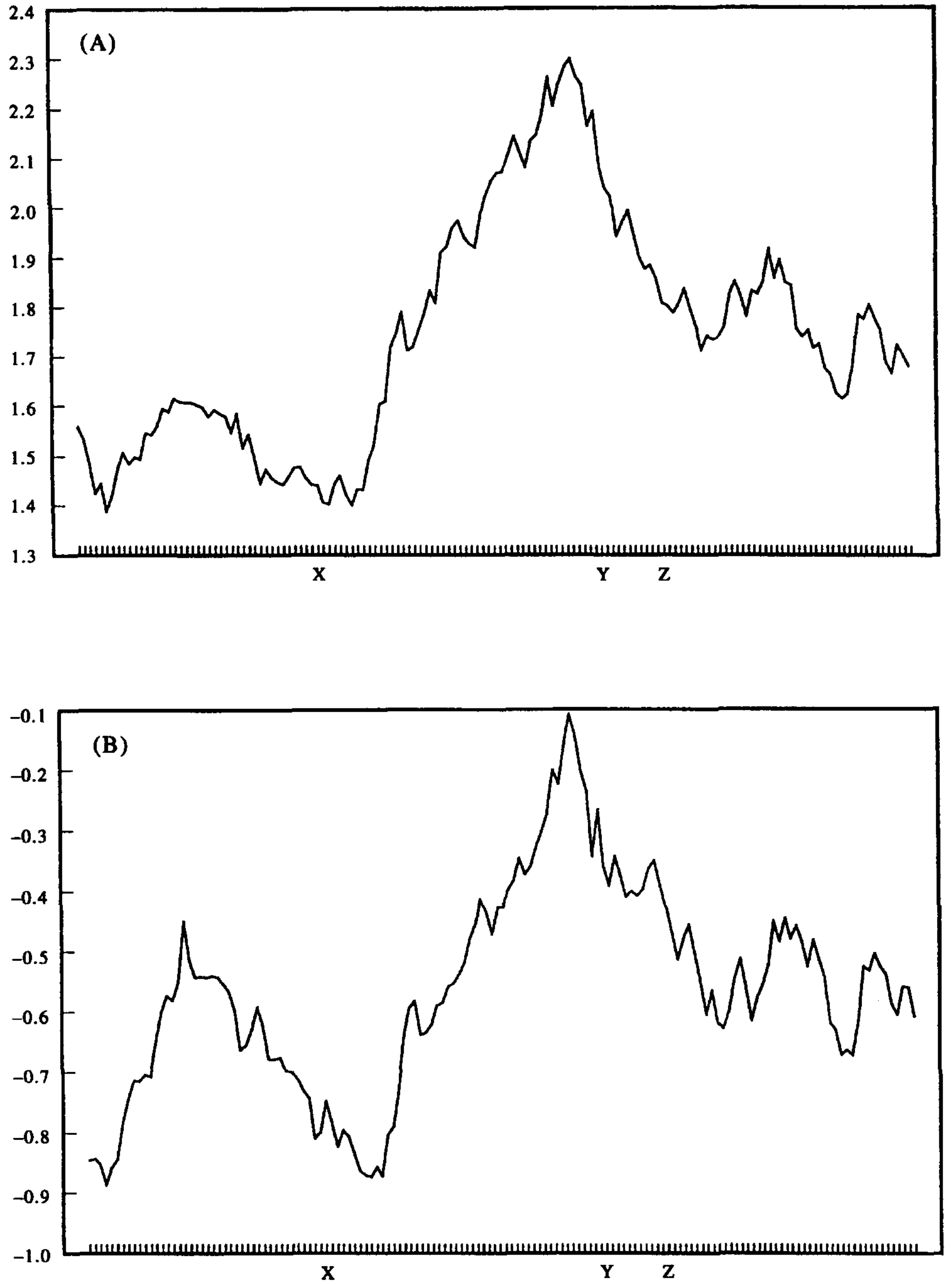

FIGURE 1. Plots of the logarithm of daily spot rates for the French franc and British pound (Spot rates are sampled every thirtieth weekday; x = 10/9/1979; y = 9/19/1985; z = 2/18/1987. (A) French Franc (entire period 9/3/1974–5/27/1992; 4446 observations). (B) British Pound (entire period 9/3/1974–5/27/1992; 4446 observations).

the need to augment an initially more restricted specification. A similar practice is found in Dekimpe and Hanssens (1995).

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 we therefore relied on the results of the first test. Allowing for deterministic components did not change our conclusions, however. Also in the first and second subperiods, we could never reject the unit root null hypothesis. In none of the analyzed cases did we find evidence of significant day-of-the-week effects, suggesting that seasonality is not a problem with this particular data set.

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 US dollar declined sharply and the trend was significant (visual support for this finding is found in Figure 1). When this time interval was excluded from the third subperiod, all spot rates were again found to have a unit root. In what follows, the third subperiod is restricted to start on 19 February 1987.

To assess whether an explicit allowance for structural breaks would alter our conclusion that all exchange rates have a unit root, we applied Perron's (1989) trend breaking test: *i.e.* we tested the unit root null hypothesis against the alternative hypothesis that the process is stationary around a deterministic trend with a shift in intercept and slope at a single, known point in time.[12] We considered three different break-points: 10 October 1979 (the end of our first subperiod), 19 September 1985 (the end of the second subperiod), and 18 February 1987 (the start of our new third subperiod). In none of the cases was the unit root null hypothesis rejected. To further investigate the behavior of the exchange rates in the third subperiod, we subsequently limited the sample to the period between 20 September 1985 (the Plaza Accord) and 27 May 1992, and allowed for a single break on 17 February 1987 (the Louvre Accord). Again, the unit root null hypothesis could not be rejected for any of the currencies. It should be noted, however, that the Perron tests consider the null of a unit root process against the alternative of stationarity around a 'broken' deterministic trend (*i.e.* stationarity before and after the break-point), and that they do not explicitly deal with the situation we encountered in the third subperiod, *i.e.* a transition from stationarity to unit-root behavior. In this respect, the recent work by Banerjee, Lumsdaine and Stock (1992) who use recursive and rolling tests based on changing subsamples of the data may be of interest. The application of moving window unit root tests and persistence calculations is beyond the scope of the current paper, however, and we leave an in-depth investigation on the causes of the stationarity in many exchange rate series between the Plaza and the Louvre Accord as an important area for future research.

Finally, we tested for the presence of a second unit root, which was rejected in all periods.[13] Overall, our results concur with those reported in Baillie and Bollerslev (1989a) and Copeland (1991).[14]

III.B. Univariate V_k estimates

Figure 2 presents the plots of the $\hat{V}_k$ values for the British 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 most interest are for k ranging from $T/3$ to $T/2$, where T is the sample size (see *e.g.* Campbell and Mankiw, 1987). The random walk benchmark, as discussed in Section I, is $(T-k)/T$.

One recurring pattern in the $\hat{V}_k$ plots is that they are hump shaped: they first rise above one, and then fall.[15] Huizinga (1987), who found a similar pattern for real exchange rates, shows 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 ⟨5⟩). Similarly, the point where $\hat{V}_k$ first falls below $(T-k)/T$ is not a reliable indicator of the minimum time horizon in which the spot rate exhibits mean-reverting behavior.

While all plots are hump-shaped, not all fall below the random walk benchmark. For the entire period, the $\hat{V}_k$ approach the random walk benchmark. In the first subperiod, on the other hand, the $\hat{V}_k$ are always substantially larger than one. Hence, unexpected changes were typically reinforced by other positive changes and the exchange rate continued to diverge from its pre-shock expectation. During the second subperiod, the permanent component becomes less pronounced, and the $\hat{V}_k$ are very close to the random walk benchmark. Finally, in the third subperiod, $\hat{V}_k$ falls below $(T-k)/T$, indicating the presence of a mean-reverting component.

As can be seen in Figure 2, the $\hat{V}_k$ vary with the selected value of k, and the estimated standard errors become quite large.[16] As such, many of the univariate V_k estimates are not different from one in conventional significance tests. When interpreting these graphs, attention is therefore focused *not* on an individual V_k estimate and its associated significance level, but on (i) the overall pattern of successive V_k estimates, as well as on (ii) 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/differences of these strings across real exchange rates. Liu and He (1991) also indicate the relevance of *long* successive strings of (even small) autocorrelations *of the same sign* in their discussion on the behavior of weekly nominal exchange rates.

The graphs of the $\hat{V}_k$ for the other ERM currencies, *i.e.* the French franc, the mark and the lire all exhibit a similar pattern as the pound (which joined the ERM in 1990), *i.e.* they indicate that their mean-reverting component becomes stronger in the second subperiod and even stronger in the third subperiod.[17]

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 and the lire, 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 the 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. Because of space limitations, we do not give the graphs for all currencies in all subperiods. Instead, we present in Table 2 the maximum of

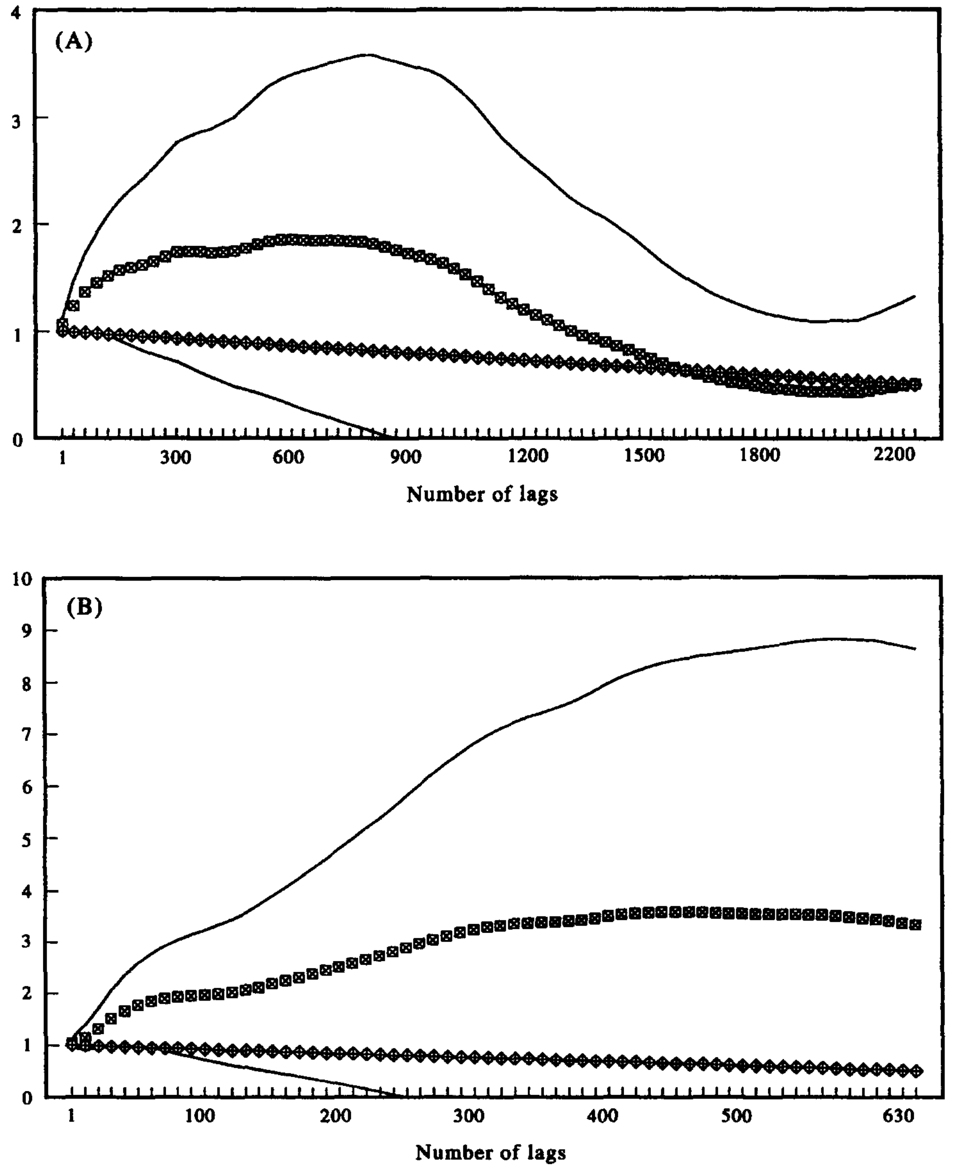

FIGURE 2—*legend opposite.*

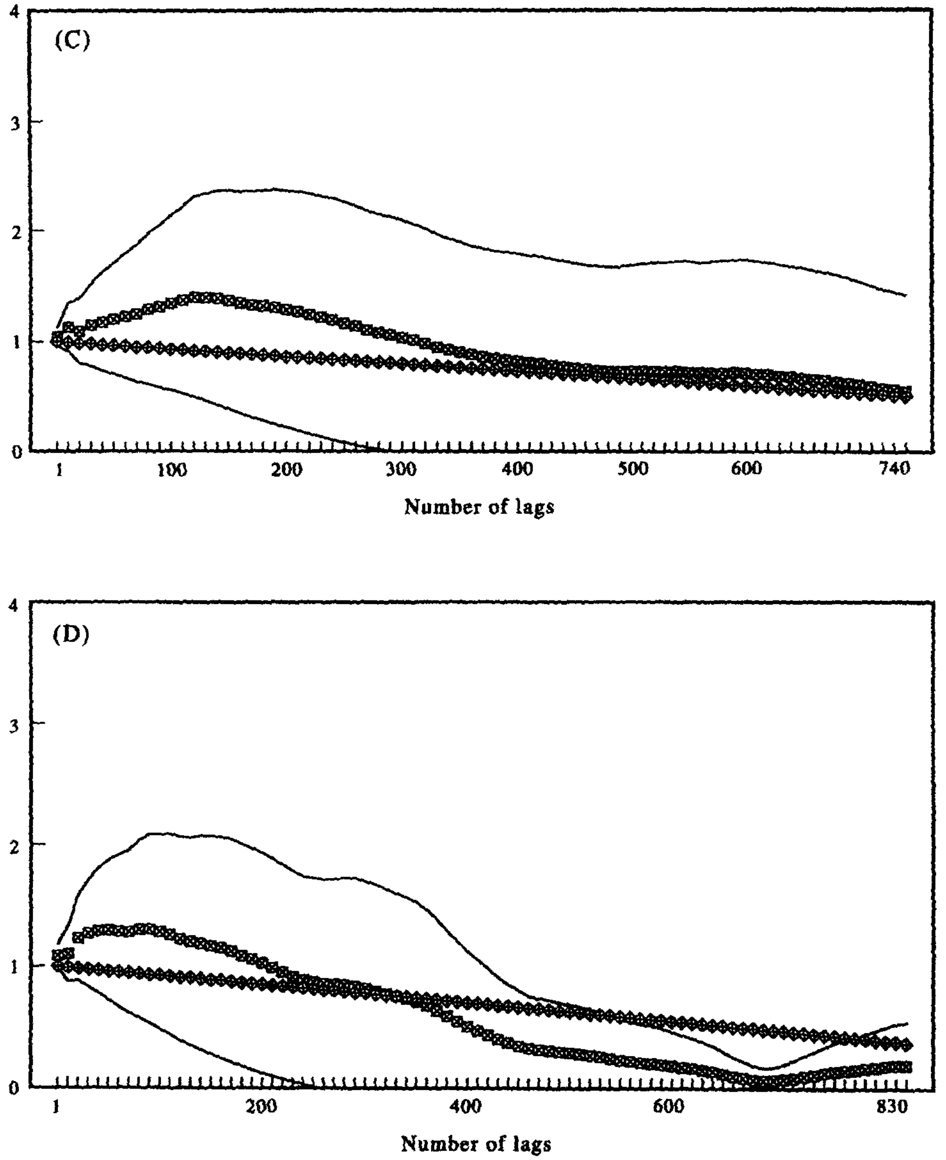

FIGURE 2. Univariate V_k estimates for the British Pound spot rate (V_k estimates are sampled every thirtieth observation for the entire period and every tenth observation for the subperiods) (A) Entire period (9/3/1974–5/27/1992; 4446 observations). (B) First subperiod (9/3/1974–10/9/1979; 1728 observations). (C) Second subperiod (10/10/1979–9/19/1985; 1490 observations). (D) Third subperiod (2/19/1987–5/27/1992; 1326 observations). ⊠, Vestimates; ⊕, random walk benchmark; —, 95% confidence bound.

the $\hat{V}_k$ over the range $k = 0$ to $T/2$ (which gives an idea of the hump shape in all currencies), as well as $\hat{V}_k$ for $k = 0.4T$ (*i.e.* for a value of k located near the midpoint of the aforementioned interval $[T/3, T/2]$. An inspection of the $\hat{V}_k$ at

$k = 0.4T$ allows the reader to verify our previous discussion on the similarities/differences among the currencies, and also indicates that in a number of instances significant differences from unity are observed.

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 (1) across spot rates and (2) across time. There is evidence in Figure 1 that, especially in certain periods and for certain

TABLE 2. V_k estimates

	V_{max} (std error; lag)	V_k at $k = 0.4^*$T (std error)	Random walk benchmark $(T-k)/T$ at $k = 0.4^*T$
	Panel A: Entire period (9 / 3 / 1974–5 / 27 / 1992; 4446 observations)		
CD	1.952 (1.075; 1010)	1.623 (1.185)	0.600
FF	2.523 (1.153; 696)	1.436 (1.049)	0.600
DM	1.951 (0.809; 573)	0.655 (0.479)	0.600
IL	2.689 (1.177; 638)	2.120 (1.549)	0.600
JY	2.047 (0.659; 345)	0.337 (0.246)	0.600
SF	1.663 (0.568; 388)	0.487 (0.356)	0.600
BP	1.854 (0.785; 597)	0.504 (0.368)	0.600
	Panel B: First subperiod (9 / 3 / 1974–10 / 9 / 1979; 1278 observations)		
CD	1.784 (0.794; 189)	1.081 (0.790)	0.600
FF	1.580 (0.871; 290)	1.120 (0.819)	0.600
DM	1.216 (0.408; 107)	0.600 (0.439)	0.600
IL	2.323 (1.253; 278)	1.950 (1.425)	0.600
JY	2.892 (1.535; 269)	1.965 (1.436)	0.600
SF	1.235 (0.343; 73)	0.530 (0.387)	0.600
BP	3.588 (2.455; 448)	3.548 (2.593)	0.600
	Panel C: Second subperiod (10 / 10 / 1979–9 / 19 / 1985; 1490 observations)		
CD	1.318 (0.163; 16)	0.332 (0.243)	0.600
FF	1.249 (0.190; 25)	0.606 (0.443)	0.600
DM	1.254 (0.195; 26)	0.211 (0.154)	0.600
IL	1.102 (0.165; 24)	0.312 (0.228)	0.600
JY	1.392 (0.250; 35)	0.029 (0.021)	0.600
SF	1.296 (0.216; 30)	0.141 (0.103)	0.600
BP	1.406 (0.466; 122)	0.714 (0.522)	0.600
	Panel D: Third subperiod (2 / 19 / 1987–5 / 27 / 1992; 1326 observations)		
CD	1.036 (0.099; 8)	0.932 (0.681)	0.600
FF	1.328 (0.415; 96)	0.157 (0.115)	0.600
DM	1.310 (0.392; 88)	0.124 (0.091)	0.600
IL	1.254 (0.377; 89)	0.173 (0.126)	0.600
JY	1.341 (0.288; 45)	0.638 (0.467)	0.600
SF	1.338 (0.420; 97)	0.222 (0.162)	0.600
BP	1.309 (0.387; 86)	0.255 (0.186)	0.600

See Note to Table 1.

currencies, the *long-run* behavior of exchange rates deviates from that of a pure 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 for the mark 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 US 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 exhibits a strong mean-reverting component in the third subperiod, there still is a permanent component in its behavior; *i.e.* there is no complete mean reversion. For the yen, on the other hand, we do not find evidence of an increased mean reversion in the third subperiod. Overall, the similarity among the ERM currencies is quite striking. The importance of the ERM will become even more obvious in our multivariate analyses.

III.C. Univariate A(1) estimates

For each currency and for each subperiod, a total of 240 ARMA (p, q) models (*i.e.* for $p = 0, 1, \ldots, 15$ and for $q = 1, \ldots, 15$) were estimated.[18] Maximum-likelihood estimation procedures which explicitly allow for roots on the unit circle were used. In Table 3, we present the $A(1)$ estimates for the ARMA models which minimized Akaike's Information Criterion.

The general picture that arises from Table 3 is an $\hat{A}(1)$ which is slightly greater than one. In a number of instances, the persistence estimates are even significantly higher than the random walk benchmark of one, and a significant decrease from period 1 to period 3 is observed for a number of currencies (Canadian dollar, lire, yen). At first sight, one could therefore conclude that the $\hat{A}(1)$ estimates corroborate our earlier conclusion that the long-run behavior of certain currencies tends to deviate from that of a pure random walk. However, the underlying causes for these deviations seem to be somewhat different. Indeed, the $\hat{A}(1)$ do not reflect an important mean-reverting component in the second and third subperiods. This discrepancy between the different measures is not uncommon.[19] A possible explanation is that low-order ARMA 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 ARMA models, where the maximum p and q equal 15, focus on capturing the short-term dynamics, and thereby (slightly) misspecify the higher-order autocorrelations by fitting a smaller-order model. This is illustrated in Cochrane (1988) and Watson (1986). The former shows that a simple ARIMA (0, 1, 1)

TABLE 3. Univariate $A(1)$ estimates

	CD	FF	DM	IL	JY	SF	BP
	Panel A: Entire period (9/3/1974–5/27/1992; 4446 observations)						
$A(1)$	1.156	1.170*	1.124	1.715	1.319*	1.325	1.305
std. error	(0.157)	(0.064)	(0.127)	(0.436)	(0.080)	(0.187)	(0.294)
$(p,q)^a$	(5,7)	(8,7)	(6,7)	(11,7)	(8,15)	(6,5)	(14,9)
	Panel B: First subperiod (9/3/1974–10/9/1979; 1278 observations)						
$A(1)$	1.477*	0.982	1.083	1.771*	1.465	1.984	1.626
std. error	(0.211)	(0.013)	(0.090)	(0.331)	(0.136)	(1.223)	(0.532)
$(p,q)^a$	(1,3)	(5,5)	(12,6)	(15,15)	(8,15)	(13,7)	(2,2)
	Panel C: Second subperiod (10/10/1979–9/19/1985; 1490 observations)						
$A(1)$	1.259*	1.242	1.092	1.252	1.342	1.081*	1.061
std. error	(0.098)	(0.168)	(0.064)	(0.221)	(0.550)	(0.037)	(0.126)
$(p,q)^a$	(10,11)	(6,7)	(6,4)	(9,9)	(11,11)	(0,2)	(14,12)
	Panel D: Third subperiod (2/19/1987–5/27/1992; 1326 observations)						
$A(1)$	1.013	1.012	1.155	1.116	1.160*	1.050	1.034
std. error	(0.040)	(0.042)	(0.105)	(0.098)	(0.065)	(0.062)	(0.022)
$(p,q)^i$	(4,2)	(5,7)	(6,11)	(6,11)	(11,11)	(10,10)	(2,2)

Notes: $^a(p,q)$ were selected on the basis of the Akaike Information Criterion.
*indicates significantly ($\alpha < 0.05$) different from the random-walk benchmark.
See Note to Table 1.

model *cannot* capture simultaneously a positive autocorrelation at short lags and a small random walk component at long lags. Since the maximum likelihood procedure focuses on matching the series' short-term properties, the accuracy of the long-run representation is sacrificed. Similarly, Watson (1986, Figure 1b) illustrates graphically that low-order ARMA models cannot represent a large number of small high-order correlations. 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 $\hat{V}_k$ graph for the full period suggests an initial sequence of small positive autocorrelations followed by a long string of small negative autocorrelations. The ARMA(14, 9) model results in an $\hat{A}(1)$ estimate of 1.305, which suggests that it captures the initial increase, but does not reflect the subsequent string of negative autocorrelations.[20]

In summary, somewhat different results are obtained for the $A(1)$ and V_k measures. These differences are basically due to their different treatment of the (insignificant) higher order autocorrelation coefficients. This issue is discussed further in the following section which compares multivariate persistence estimates based on, respectively, a restricted and unrestricted VAR model.

III.D. Multivariate persistence estimates

As indicated before, the estimation procedure for the multivariate persistence measures is affected by the presence/absence of cointegration among the currencies. We therefore applied the Johansen and Juselius (1990) FIML approach to cointegration testing. Both the test with and without drift were applied, and the obtained test statistics were compared with the critical values listed in Osterwald-Lenum (1992). Detailed test results are available from the authors upon request, but in none of the cases (entire period, three subperiods) did we find any evidence of cointegration. Baillie and Bollerslev (1989a), on the other hand, report one cointegrating vector between the seven currencies considered in our study for the period between 1 March 1980 and 28 January 1985. However, Sephton and Larsen (1991) re-analyzed the data of Baillie and Bollerslev for 16 different subsamples, and found no cointegration in a vast majority of the cases, which suggests that the outcome of the cointegration tests depends upon the time period sampled.

Based on our cointegration tests, we did not include any error-correction terms into our model, and used an inverted VAR model to derive an estimate of $\mathbf{A}(1)$ and Σ, *i.e.* exchange rate changes were explained only in terms of lagged changes in the own currency and all other currencies. 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 $\mathbf{VC}(\tau_t)$ were calculated according to equation $\langle 8 \rangle$, after which the (i, j) elements were normalized by $\mathrm{Var}(\Delta S_{j,t})$. The associated standard errors were computed using the delta method (see Van de Gucht *et al.*, 1995 for details). 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 4, are similar to those obtained from the unrestricted VAR(2) specification.[21] 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 effect of an unexpected change in the own currency, taking into account all possible channels of influence, including lagged effects and contemporaneous shocks in other currencies induced by the shock of interest. The off-diagonal element (i, j) is an estimate of the effect of a shock in currency j on the long-run value of currency i.

With respect to the diagonal elements in Table 4, the univariate pattern of increased mean reversion over time is not always observed in the multivariate estimates. For most currencies (*e.g.* French franc, mark, Swiss franc), the 95 percent 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.[22] However, the associated standard errors for

TABLE 4. Multivariate persistence estimates using a restricted VAR(2) model[a].

	Source of the shock						
Effect on	CD	FF	DM	IL	JY	SF	BP
	Panel A: Entire period (9 / 3 / 1974–5 / 27 / 1992; 4446 observations)						
CD	1.124	0.092	0.091	0.096	0.054	0.076	0.102
	(0.034)	(0.006)	(0.005)	(0.007)	(0.004)	(0.004)	(0.007)
FF	0.639	1.017	0.924	0.960	0.652	0.726	0.745
	(0.036)	(0.018)	(0.012)	(0.017)	(0.015)	(0.012)	(0.015)
DM	0.652	0.953	1.015	0.960	0.684	0.785	0.768
	(0.036)	(0.012)	(0.010)	(0.014)	(0.015)	(0.010)	(0.014)
IL	0.595	0.858	0.832	1.081	0.582	0.647	0.677
	(0.040)	(0.016)	(0.012)	(0.020)	(0.015)	(0.012)	(0.017)
JY	0.337	0.584	0.593	0.582	1.055	0.511	0.497
	(0.024)	(0.014)	(0.014)	(0.015)	(0.026)	(0.014)	(0.012)
SF	0.729	1.012	1.061	1.009	0.796	1.035	0.811
	(0.039)	(0.016)	(0.014)	(0.020)	(0.021)	(0.018)	(0.018)
BP	0.707	0.747	0.747	0.760	0.557	0.584	1.086
	(0.045)	(0.015)	(0.014)	(0.018)	(0.013)	(0.013)	(0.027)
	Panel B: First subperiod (9 / 3 / 1974–10 / 9 / 1979; 1278 observations)						
CD	1.216	0.005	0.019	−0.017	−0.097	0.005	0.018
	(0.077)	(0.014)	(0.015)	(0.027)	(0.021)	(0.010)	(0.015)
FF	0.026	0.964	0.681	0.545	0.450	0.410	0.437
	(0.062)	(0.046)	(0.035)	(0.029)	(0.028)	(0.025)	(0.033)
DM	0.103	0.784	0.992	0.508	0.506	0.547	0.455
	(0.080)	(0.040)	(0.042)	(0.035)	(0.025)	(0.028)	(0.039)
IL	−0.080	0.537	0.435	1.454	0.320	0.299	0.353
	(0.125)	(0.030)	(0.030)	(0.087)	(0.017)	(0.022)	(0.034)
JY	−0.477	0.469	0.458	0.338	1.022	0.300	0.256
	(0.101)	(0.028)	(0.022)	(0.016)	(0.011)	(0.016)	(0.027)
SF	0.059	1.003	1.164	0.742	0.705	1.010	0.558
	(0.119)	(0.061)	(0.060)	(0.056)	(0.039)	(0.057)	(0.059)
BP	0.091	0.460	0.416	0.377	0.259	0.240	0.979
	(0.069)	(0.035)	(0.036)	(0.035)	(0.029)	(0.025)	(0.030)

— *Continued*

Table 4. (*Continued*)

Effect on	Source of the shock CD	FF	DM	IL	JY	SF	BP
	Panel C: Second subperiod (10 / 10 / 1979–9 / 19 / 1985; 1490 observations)						
CD	1.136	0.176	0.187	0.191	0.147	0.171	0.166
	(0.063)	(0.016)	(0.016)	(0.017)	(0.019)	(0.015)	(0.013)
FF	1.520	1.006	0.973	1.007	0.613	0.830	0.752
	(0.137)	(0.054)	(0.052)	(0.055)	(0.048)	(0.047)	(0.046)
DM	1.518	0.915	1.039	1.001	0.640	0.879	0.760
	(0.130)	(0.048)	(0.051)	(0.052)	(0.048)	(0.047)	(0.044)
IL	1.342	0.818	0.866	0.962	0.543	0.734	0.639
	(0.116)	(0.044)	(0.045)	(0.049)	(0.041)	(0.041)	(0.040)
JY	0.999	0.480	0.534	0.524	1.032	0.531	0.415
	(0.128)	(0.037)	(0.040)	(0.040)	(0.055)	(0.038)	(0.035)
SF	1.685	0.945	1.066	1.029	0.771	1.082	0.814
	(0.142)	(0.054)	(0.057)	(0.057)	(0.055)	(0.056)	(0.051)
BP	1.419	0.743	0.800	0.778	0.523	0.707	1.057
	(0.105)	(0.046)	(0.047)	(0.049)	(0.044)	(0.044)	(0.048)
	Panel D: Third subperiod (2 / 19 / 1987–5 / 27 / 1992; 1326 observations)						
CD	0.980	0.033	0.029	0.037	0.010	0.021	0.045
	(0.006)	(0.016)	(0.016)	(0.017)	(0.011)	(0.013)	(0.014)
FF	0.205	1.068	1.010	1.090	0.732	0.846	0.905
	(0.103)	(0.067)	(0.064)	(0.069)	(0.057)	(0.057)	(0.058)
DM	0.198	1.104	1.058	1.131	0.761	0.883	0.939
	(0.107)	(0.070)	(0.067)	(0.071)	(0.059)	(0.059)	(0.060)
IL	0.209	1.008	0.956	1.069	0.700	0.805	0.867
	(0.097)	(0.063)	(0.061)	(0.066)	(0.055)	(0.054)	(0.056)
JY	0.062	0.728	0.692	0.753	1.100	0.631	0.697
	(0.070)	(0.056)	(0.053)	(0.058)	(0.064)	(0.047)	(0.050)
SF	0.165	1.088	1.038	1.118	0.816	1.007	0.965
	(0.108)	(0.073)	(0.069)	(0.075)	(0.061)	(0.065)	(0.064)
BP	0.297	0.977	0.927	1.012	0.757	0.810	1.138
	(0.094)	(0.062)	(0.059)	(0.064)	(0.055)	(0.054)	(0.069)

Note: [a]Standard errors are reported in parentheses.
See Note to Table 1.

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.[23]

The following observations can be made with respect to the off-diagonal elements. First, they tend to be smaller than the diagonal elements in the same column, and their 95 percent confidence intervals usually lie below one (some exceptions can nevertheless be observed, such as the estimates describing the long-run effect of the Swiss franc, which are sometimes larger than one). This implies that, on average, a shock has a larger effect 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 effect 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 effect 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 effect 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 effect 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, which was also observed in Liu and He (1991).

Third, the off-diagonal elements are not symmetric. For example, for the entire period, the long-run effect of a shock to the Swiss franc on the French franc is 0.726, while the long-run effect 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 effect on the European currencies of a shock in the yen ranges between 0.700 and 0.816, while the effect 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 UK 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 percent 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 effect 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 effect of such a shock on the mark (0.992) was quite different from the long-run effect 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 percent above and below and central rate, and this greater latitude is reflected also in the persistence estimates: row 4 in panel C (which measures the effect of different shocks on the lire) differs more from row 3 (the effect on the mark) than the latter does from row 2 (the effect on the French franc).

Fifth, 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.

Finally, in order to assess the relative contribution to the multivariate persistence estimates of, respectively, the instantaneous cross-reactions and the lagged effects, we derived persistence estimates (i) from a model where Σ is estimated freely (*i.e.* where it is allowed to have non-zero off-diagonal elements), and (ii) from a model where Σ is restricted to be a diagonal matrix (*i.e.* where instantaneous cross-effects are excluded).[24] The need to incorporate the instantaneous cross-country effects was evident from the Lagrange multiplier test suggested by Breusch and Pagan (1980): the null hypothesis of zero off-diagonal elements was firmly rejected even at the $p = 0.0001$ level. In terms of their impact on the persistence matrix, we found that especially the off-diagonal elements of the persistence matrix changed considerably.[25] For the mark and the lire, for example, restricting Σ changed the long-run effect of a shock to the French franc (for the entire period) from, respectively, 1.005 (0.045)[26] and 0.886 (0.041) to 0.107 (0.078) and 0.056 (0.050). These results imply that a large fraction of the long-term effect on another currency is due to the fact that a certain event does not affect one currency in isolation, but also initiates a whole set of contemporaneous reactions (shocks) in the other currencies.

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 US 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 effect of an unexpected change in the current value of the exchange rate are discussed: Campbell and Mankiw's (1987) $A(1)$ 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 or when testing models of exchange rate determination. We do find, however, that the mean-reverting component of several ERM 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 allow us to quantify the long-run effect 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 effect of shocks originating in another currency is similar for the ERM currencies. Furthermore, the cross-effects reflect certain changes in US monetary policy, and the increase in coordinated intervention after the Louvre Accord. Our results also suggest that a large portion of the long-run effect of a shock in one currency on another currency can be attributed to the fact that shocks do not occur in isolation, but also initiate contemporaneous shocks in the other currencies. Finally, our persistence approach to studying long-run interrelationships between currencies complements 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. Also, more work is needed on the distinct behavior of most currencies between the Plaza Accord and the Louvre Accord. As a first step, a moving-window test could be used to assess whether this is the only one-year period where the unit root null hypothesis is rejected or whether this occurs more frequently, which could indicate that a one-year time span may be too small to disentangle long-run movements in exchange rate series (see also Hakkio and Rush, 1991). 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 effect of any shock should be the same on a currency's spot rate as on its forward rate. Baillie and Bollerslev (1989a) 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.

Finally, multivariate persistence measures may serve as a useful tool to study the interrelationships between the long-run properties of the exchange rate series and the long-run properties of market fundamentals, and can therefore be used to assess the long-run validity of the restrictions imposed by different models of exchange rate determinations, such as the pure flexible price monetary model of Frenkel (1976), the sticky price monetary model of Dornbush (1976) or the real interest rate differential model of Frankel (1979). Recently, several authors have treated the reduced form of these exchange rate models as a *long-run* equilibrium condition which can be tested empirically in a cointegration framework (see *e.g.* MacDonald and Taylor, 1994; Gardeazabal and Regulez, 1992). Unfortunately, this cointegration approach suffers from interpretational difficulties when multiple equilibrium relationships are found. MacDonald and Taylor, for example, find three cointegrating vectors between the sterling–dollar exchange rate and some market fundamentals, while Gardeazabal and Regulez find up to six cointegrating vectors for the yen–dollar exchange rate. Moreover, the FIML approach used in their analyses only allows one to test coefficient restrictions on the *full* set of cointegrating vectors, thereby prohibiting statistical tests on whether one specific vector satisfies the restrictions imposed by the exchange rate model under consideration. The multivariate persistence approach can complement the aforementioned studies in that: (i) the cointegrating relationships are automatically reflected in the rank of the persistence matrix (see Section I.C.); (ii) it does not suffer from the interpretational difficulties which arise when more than one cointegrating vector is found (Lüthkepohl and Reimers, 1992); and (iii) it allows for standard statistical tests on the equality and magnitude of the respective estimates. Specifically, to assess the long-run value of the monetary model discussed in MacDonald and Taylor (1994), one could first estimate an error-correlation model between the exchange rate and the respective market fundamentals, invert the model to derive the corresponding persistence estimates, and test *e.g.* whether the long-run effect on the exchange rate of a shock to the domestic money supply has the same absolute magnitude (but opposite sign) as the long-run effect of a shock to the foreign money supply. Similarly, in testing the Dornbush sticky-price model, one can add the short-term interest rates to the model, and test whether their long-run effect on the exchange rate is significantly different from zero.

Notes

1. An exception is the study by Whitt (1992), which is discussed in somewhat more detail

in a following section.

2. Sephton and Larsen (1991) do not explicitly consider changes in monetary policy, but provide some indirect evidence 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 depends on the time frame used in the cointegration tests.
3. We point out that a persistence of one does not guarantee that the series follows a pure random walk (consider *e.g.* the case where the ARMA model for ΔS_t has $\theta_1 = 0.2$ and $\phi_7 = 0.2$). However, a persistence estimate different from one guarantees a long-run behavior different from that of a pure random walk.
4. The variance ratio given in equation ⟨4⟩ was also used by Liu and He (1991) to determine the significance of the 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.
5. Although $A(1)$ and V_k have the same value for the random-walk benchmark (*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. As k approaches infinity, $A(1) = [V/(1 - R^2)]^{1/2}$, where V is the limiting value of V_k, and R^2 is the proportion of the variance of ΔS_t explained by equation ⟨1⟩. 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.
6. 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).
7. Mathematical details are given in Van de Gucht *et al.* (1995), and are available from the authors upon request. We thank Kevin Lee and Richard Pierse for their suggestions with respect to the derivations of the standard error estimates.
8. Cointegration studies of foreign exchange markets have only examined the existence of a long-run equilibrium relationship without further assessment of the nature of that relationship (see *e.g.* Baillie and Bollerslev, 1989a; Hakkio and Rush, 1989; MacDonald and Taylor, 1991).
9. Following Copeland (1991), banking holidays (182 days) were 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.
10. The starting date for the third subperiod in Table 1 differs from the date mentioned before. We refer to Section III.A. for a formal motivation of the new starting date.
11. 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 (1989b).
12. We included six lagged differences in the implementation of the Perron tests.
13. 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.
14. 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 since the results of Bayesian unit root tests have been shown to be very sensitive to the choice of the prior distribution, while the test used by Whitt is biased towards finding stationarity (Phillips, 1991).
15. 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.

16. The standard error $\hat{V}_k$ is derived as $\hat{V}_k/[0.75^*T/(k+1)]^{1/2}$. Mathematically, it is easy to show that, for a given value of $\hat{V}_k$, the standard errors increase with k. Intuitively, this increase can be explained by the fact that larger values of k result in a smaller number of *nonoverlapping* runs that can be used in the derivation of the variance of the long-run component. This number of runs gives a rough indication for the diminishing number of degrees of freedom available in the estimation (Cochrane, 1988).
17. Given that all spot rates are expressed per US dollar, their behavior may be determined primarily by unexpected US events, which may explain why the pattern of persistence estimates is similar for most currencies.
18. To ensure that a global optimum (within the limits set by the maximum values of p and q) would be found, 240 ARMA(p,q) estimates were calculated. In practice, the number of ARMA models to be estimated could be reduced by assessing the robustness of the findings across a smaller subset, or by using special subroutines to select which models to estimate. The order of the moving average component was always at least one to be able to detect overdifferencing. No such evidence was found, giving further support to the presence of a unit root in all currencies. 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 3.
19. Cochrane (1988), for example, found no evidence of a large random walk component in GNP, while Campbell and Mankiw (1987) reported $A(1)$ values larger than or equal to one.
20. One could, of course, attempt to estimate a very high-order ARIMA model. We did not pursue this strategy since (1) this would result in computational difficulties (convergence problems and software limitations on the maximum values for p and q), (2) these highly over-parameterized models were unlikely to be selected with the AIC criterion, and (3) this would not be in line with the underlying philosophy of the $A(1)$ methodology.
21. To give the reader some insights into the relative importance of the lagged effects in the restricted VAR specification, we computed for the entire period the number of withheld terms. For the VAR(2) model, 6 own effects (maximum = 14) and 16 (maximum = 84) cross effects were still included in the restricted specification. For the VAR(10) model, the corresponding figures are 26 (maximum = 70) and 112 (maximum = 420). In what follows, we also assess the relative importance of lagged and contemporaneous effects on the resulting persistence estimates.
22. 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 after the first 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 9,800 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.
23. As with the univariate analyses, we will assess the temporal stability of the multivariate persistence estimates across different subperiods. However, to also get some insights into the stability of the model's *short-run* properties, we tested for parameter equality across the different subperiods in the unrestricted VAR(2) models. In spite of the over-parameterization of these models (which makes the detection of significant differences more difficult), we did find for a number of currencies that the lagged own- and/or cross-effects had changed significantly. For the lire, for example, the coefficients of the second subperiod were significantly different from the ones in the other periods ($p < 0.01$). This was also reflected in the restricted VAR specifications, where

(in a number of instances) different explanatory variables were withheld for the respective subperiods.

24. We implemented this comparison in terms of an *un*restricted VAR(2) model (unrestricted in terms of the lagged effects), because in that case the lagged effects are estimated in the same way irrespective of whether Σ is a diagonal matrix. Indeed, if all equations have the same right-hand side variables, there is no efficiency gain from a systems approach, and OLS could be used in both cases. As such, the autoregressive coefficients used to derive $A(1) = [\Pi(1)]^{-1}$ were not affected by whether all elements of Σ were subsequently estimated or only the diagonal elements.
25. Overall, the changes in the diagonal elements were much smaller, and in many cases marginal. For example, for the entire period, the long-run impact on the mark (lire) of a shock to the mark (lire) changed from 1.073 (1.091) to, respectively, 1.133 and 0.972. The largest change (again for the entire period) was observed for the French franc, where the long-run impact of its own shock changed from 1.058 to 0.763.
26. The values between parentheses are the associated asymptotic standard errors.

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