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ABSTRACT

Theoretical and empirical studies have treated excess returns as processes with time-varying but temporary disturbances. By contrast, empirical evidence indicates that the behavior of asset price levels can be well-approximated by processes with some permanent disturbances. These two observations restrict the relationship between the levels of asset prices and the excess returns they generate. In this paper, we begin by testing these restrictions for foreign exchange and bond returns. Surprisingly, we reject these restrictions for some returns, implying that excess returns contain some permanent shocks. We then evaluate the possible reasons for these results. This behavior appears inconsistent with conventional models of the risk premia. On the other hand, this behavior could arise from the presence of some traders in the market who have "regressive expectations" or from anticipated shifts in the distribution of asset prices inducing a "peso problem". We test and reject a simple model implied by a steady-state presence of traders with regressive expectations. However, we cannot distinguish between a model where the effects of these traders vary over time or where a peso problem exists or both.

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Many excess returns in financial markets may be written as the return from holding a forward contract on an asset relative to buying the asset in the future spot market. Examples include holding long term bonds and uncovered positions on holding foreign currency denominated bonds. Excess returns of this type can be decomposed into two components: a risk premium and an *ex post* error in forecasting the risky variable. In theoretical and empirical applications, researchers have treated the risk premium component as a time varying but mean reverting process with temporary disturbances. Similarly, non-overlapping forecast errors have been treated as transitory white noise processes, an implication of standard rational expectations assumptions. Therefore, as the sum of these two terms, excess returns have been assumed to contain only short-lived transitory disturbances.

By contrast, the empirical evidence is mounting that asset prices themselves are well-approximated by processes with some permanent disturbances.[1] Although in practice researchers may be unable to distinguish between literally permanent disturbances and very slow mean-reverting components, the overall empirical evidence indicates that the degree of persistence in shocks to asset price disturbances is very high.

Since researchers treat asset prices as containing permanent shocks but excess returns as containing only temporary shocks, this treatment suggests a relationship between the disturbances to the levels of spot asset prices and their excess returns. In particular, the long run relationship between realized spot rates and their corresponding forward rates must be such that the excess returns they generate do not contain any permanent disturbances. That is, the permanent shocks to spot and forward rates must cancel out when they are combined to form excess returns. Perhaps surprisingly, these restrictions have not yet been directly tested in the literature.

In this paper, we begin by testing these restrictions using data from the foreign exchange market and the interest rate term structure for the U.S., the U.K., Germany, and Japan. A rejection of the restrictions implies that some component of excess returns contains permanent disturbances, in contrast to standard practice. Surprisingly, we reject the restrictions for many of the excess returns, suggesting

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[1]For example, so-called unit roots have been found in nominal interest rates by Campbell and Shiller (1987) and Mishkin (1991), and in exchange rates by Baillie and Bollerslev (1989) and Meese and Singleton (1982). Although not the focus of the empirical work in this paper, researchers have found similar behavior in equity and commodity prices.

that some component of excess returns indeed contains the same degree of persistence as asset prices.[2]

Given this striking result, we then evaluate the nature of the evidence and its implications for the behavior of asset returns. We generate estimates of the component of permanent disturbances in excess returns and examine its behavior over time in the context of possible explanations.

Under conventional wisdom, the only predictable component of excess returns is the risk premium. Hence, we begin by evaluating whether risk premia under standard models can be responsible for these permanent shocks. Based upon studies in the literature, we find that shocks to asset return variances do not appear sufficiently persistent to explain the permanent shock component in excess returns.

We next turn to the other component of excess returns: the forecast error. We show that excess returns may appear to contain permanent shocks under alternative assumptions about market forecasts. The first case occurs if some traders in the market are not rational and their forecasts of future asset prices are incorporated into the determination of forward rates.[3] The second case arises if the market is fully rational but occasional shifts in the distribution of asset prices induce a "peso problem" in the market's forecasts.[4] As we demonstrate, excess returns in either case may appear to contain permanent disturbances similar to our findings.

Under these two alternative explanations, our estimates of the permanent component in excess

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[2]These results may seem surprising in light of standard econometric practice where excess returns are typically treated as stationary processes. The reader may therefore wonder if empirical studies of excess returns should instead treat these variables as containing permanent disturbances. However, as we will demonstrate, the permanent shock component is empirically very small so that the excess returns are very close to stationary. Since Monte Carlo studies suggest that better econometric properties hold when treating near-stationary processes as stationary and near-unit root processes as non-stationary, the same principle likely holds here here as well. See, for example, Schwert (1989) and Campbell and Perron (1991).

[3]Delong, Schleifer, Summers, and Waldman (1990b) show that irrational "noise" traders need not be driven out of the market by rational traders, even in the long run. See also the discussion in Schleifer and Summers (1990). In the text, we will focus upon an example in the spirit of "trend chasing" chartists and fundamentalists described in Frankel and Froot (1986).

[4]The "peso problem" arises when the market rationally anticipates a future event that materializes infrequently in the sample. This effect was first pointed out by Rogoff (1980). Recent evidence of "peso problem" behavior in interest rates and in exchange rates includes Lewis (1991a) and Lewis (1989), respectively.

returns have a natural interpretation. The sign of the permanent shock reflects persistent forecast errors over some periods within the sample. According to this interpretation, the evidence implies that market participants placed positive weight on the event that the the dollar would weaken during its steady appreciation during the early 1980's. Similarly, the market believed that U.S. interest rates would be lower than they turned out to be *ex post* during their sharp ascent following the 1979 change in Fed operating procedures.

We then consider this explanation in light of the two expectational explanations. First, if these expectations arose from a "peso problem", market participants during the early 1980's persistently believed that the dollar exchange rate would shift to a depreciating regime and that U.S. interest rates would be lowered. Of course, these expectations could have been rational since both events indeed occured by the late 1980's. Second, these expectations may have come from the influence of heterogeneous traders, including some who were basing their expectations on a perceived long run equilibrium level. In this case, our estimates would suggest that these traders were responsible for the apparent bias in expectations.

We then evaluate which expectational explanation is more likely. First, we consider the case where heterogeneous traders are in the market, including some traders who condition their forecasts on a long-run mean reverting component. These expectations are consistent with the "regressive expectations" in the market considered by Froot and Frankel (1987). When an equilibrium level of these traders are in the market, then spot prices will systematically diverge from forward prices and excess returns will appear to contain permanent disturbances. Furthermore, this relationship will be constant over time.

By contrast, if peso problems are responsible, this relationship will depend upon the market's assessed probability of a shift in the distribution and should be time-varying. We test and reject the hypothesis that the relationship is constant over time. This result suggests that peso problems and not a steady state presence of heterogeneous traders are responsible. Of course, the evidence could also imply that the influence of heterogeneous trading upon forward rates were time-varying.

The paper is organized in the following way. Section 1 tests whether excess returns contain permanent disturbances in common with forward rates. For some foreign exchange and bond returns, we find evidence that they do. Section 2 evaluates whether these results are more likely to come from permanent disturbances in risk premia or forecast errors. Section 3 examines whether the effect of these disturbances is constant over time. Concluding remarks follow.

Section 1: Are Excess Returns Affected by Permanent Shocks?

We begin by considering the empirical behavior of excess returns implied by conventional assumptions in the literature. Define S_t as the (logarithm of the) spot rate on an asset at time *t* and F_t^k as the (logarithm of the) time *t* forward rate on a contract to buy or sell the asset *k* periods in the future. Then, the speculative return on a forward contract to sell the asset in the future period is,

$$(1)\quad S_{t+k} - F_t^k = rp_t + \epsilon_{t+k},$$

where rp_t is the risk premium on this speculative position and ϵ_{t+k} is the market's error in forecasting the spot rate; i.e., $\epsilon_{t+k} \equiv (S_{t+k} - E_t^m S_{t+k})$, where $E_t^m(x)$ is the market's (*m*) expectation of x given information available at time *t*.

When the speculative strategy is to hold a contract to sell dollars in the future, then S_{t+k} and F_t^k are the spot dollar price of foreign currency and its forward rate, respectively. In this case, rp_t is the "foreign exchange risk premium" while ϵ_{t+k} is the error in forecasting the exchange rate. When the speculative position is to hold a long term bond and resell it as a shorter maturity bond in the future, then S_{t+k} is the spot rate yield on the bond at $t+k$ and F_t^k is the forward rate yield implied by holding the long term bond and selling it at $t+k$. In this case, rp_t is the "term risk premium" on holding long term relative to shorter term bonds, and ϵ_{t+k} is the error in forecasting shorter term bond returns. Similar analogies can be made for other speculative positions using the general representation of excess returns in equation (1).

1.1 The Long Run Relationship Between Spot and Forward Rates

Empirical studies have treated excess returns as processes with temporary disturbances considered to be covariance stationary. In the literature, these processes have been denoted "I(0)", and we will follow this notation below. Equation (1) illustrates why we should indeed expect these returns to have only temporary disturbances under conventional assumptions. The excess returns are comprised of two components: a risk premium, rp_t, and a forecast error, ϵ_{t+k}. Risk premia have been considered stationary on theoretical grounds.[5] And under rational expectations, non-overlapping forecast errors follow white noise, a stationary process.[6] Since the sum of two stationary variables must be stationary, the sum of the risk premium and the forecast error must also be stationary under the assumptions above.

By contrast, the levels of asset prices have been found to contain very persistent shocks, well-approximated as permanent disturbances with "unit roots". These processes are covariance stationary after first differencing. Processes with these types of shocks have been denoted "I(1)" in the literature. Using this terminology empirical studies have found that the spot and forward rates on the left-hand side of equation (1) are I(1) variables.[7]

Engle and Granger (1987) have defined the relationship between variables with permanent disturbances in terms of cointegration. If a linear combination of I(1) variables is I(0) stationary, then these variables are said to be "cointegrated".[8] By definition, if X_t is an n-dimensional vector containing cointegrated I(1) variables, then a "cointegrating vector" α must exist such that $\alpha' X_t$ is I(0) stationary. Therefore, the requirement that both sides of equation (1) be stationary places restrictions upon the relationship between spot and forward rates. Specifically, if excess returns are I(0) stationary but spot

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[5]For example, standard models of time-varying risk premia imply that risk premia are stationary since they depend upon the time-series properties of the change in consumption. See Grossman and Shiller (1981) and Backus, Gregory, and Zin (1989) for some applications.

[6]Forecast errors for k periods ahead contain overlapping forecasts and therefore follow a moving average process of order k-1, also a covariance stationary process. See Hansen and Hodrick (1980).

[7]Meese and Rogoff (1983) and Meese and Singleton (1982) found that exchange rates follow a random walk. More recently, Baillie and Bollerslev (1989) test directly for unit roots in exchange rates and find that exchange rates and forward rates are cointegrated. Campbell and Shiller (1987) and Mishkin (1991) among others find that U.S. interest rates over the post-War period are non-stationary.

[8]More specifically, Engle and Granger (1987) define a vector of time series variable to be cointegrated of orders d and b if it satisfies: (i) each element of X_t is integrated of order zero after differencing d times, and (ii) there exists at least one cointegrating vertor, α, such the the linear combinations $\alpha' X_t$ is integrated of order zero after differencing d - b time. In this paper, we will only consider the case where $d = b = 1$.

and forward rates are I(1), then for the vector $X_t \equiv (S_{t+k}, F_t^k)'$ the cointegrating vector must be $\alpha = (1, -1)$ since premultiplying by this vector, $\alpha' X_t$, gives the excess returns.

We may therefore consider the validity of this assumption by estimating the cointegrating regression of spot rates on forward rates and testing whether the coefficient on forward rates equals one. In other words,

(2) $S_{t+k} = a_0 + a_1 F_t + v_t$, where Ho: $a_1 = 1$,

under the null hypothesis of stationary excess returns, so that $a_0 + v_t = rp_t + \epsilon_{t+k}$. We can test this hypothesis since a cointegrating regression of equation (2) will provide estimates of the vector, $\alpha \equiv (1, -a_1)$. By construction, this estimate will generate a vector such that $\alpha' X_t$ is a covariance stationary process equal to v_t.

Of course, this cointegrating regression will be well-defined only if S_t and F_t are cointegrated and share a unique common trend. Therefore, as a first step in our estimates reported below, we tested for cointegration and found evidence for cointegration, consistent with other studies.[9]

Since the implicit assumption of stationarity is prevalent in many studies of financial returns, testing this hypothesis is interesting regardless of the outcome. If we do not reject, then researchers may feel reassured that spot and forward rates are cointegrated one-for-one as they should be under conventional assumptions. On the other hand, if we reject, the alternative hypothesis is at least as interesting. Rejecting the hypothesis that $a_1 = 1$ in (2) implies that a component of excess returns contains a permanent shock component. Excess returns share this permanent shock with the forward rate. This result would imply that excess returns are much more persistent than previously thought.

To see why rejecting $a_1 = 1$ yields this implication, note that under the null hypothesis the residual in the cointegrating regression, v_t, equals the sum of the time-varying component of the risk premium and the forecast error. If both are stationary, then the sum must be asymptotically independent of all

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[9]See the references in footnote 1.

permanent shock components.[10] Hence, if stationary, rp_t and ϵ_{t+k} cannot contaminate the estimate of a_1. Therefore, a_1 must equal one if excess returns are stationary.

To see how departures of a_1 from one affect excess returns, subtract the forward rate from both sides of the cointegrating regression (2) and note that the result equals the definition of excess returns in (1), or:

$$(3) \qquad S_{t+k} - F_t^k = rp_t + \epsilon_{t+k} = a_0 + (a_1 - 1) F_t^k + v_t.$$

Since v_t is stationary by construction, the effect of permanent shocks in the forward rate upon excess returns will depend upon the coefficient $(a_1 - 1)$. Alternatively, if a_1 equals one, excess returns depend only upon the stationary term, v_t.

To see more directly how the permanent component in spot rates relates to forward rates, we will focus upon the one month ahead forward rate where $k = 1$, although the $k = 3$ case is a straightforward extension. For this horizon, note that equation (1) can be rewritten:

$$(1') \quad F_t^1 = S_t + \Delta S_{t+1} - rp_t + \epsilon_{t+1}$$

where Δ is the backward difference operator; i.e., $\Delta x_{t+1} \equiv x_{t+1} - x_t$. Since S_t contains a permanent shock component, S_t is I(1) and is stationary after first-differencing. That is,

$$(4) \quad S_{t+1} \equiv S_t + \Delta S_{t+1} = S_t + \text{I(0) terms.}$$

Therefore, under the null hypothesis that the excess return, $\epsilon_{t+1} + rp_t$, is I(0), all of the variables except S_t on the right-hand-side of (1') are I(0) terms. By contrast, both the forward and spot rates are I(1) so that we can write (1') in terms of its I(1) and I(0) components:

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[10]See Stock (1987), West (1988), or the surveys in Pagan and Wickens (1989), Campbell and Perron (1991), or Diebold and Nerlove (1990).

(5) $F_t^i = S_t + I(0)$ terms.

Assuming that the forward and spot rates share the same permanent shock, the I(1) component of F_t must move one-for-one with the I(1) component in the spot rate. Thus, for the pair of spot and forward rates considered in the cointegrating equation (2) given by $X_t \equiv [S_{t+1}, F_t]$, the only cointegrating vector α, that will imply $\alpha' X_t \sim I(0)$ is $\alpha \equiv [1, -a_1] = [1, -1]$. Thus, under the null, $a_1 = 1$.

1.2 Data Description

We begin our empirical investigation by estimating equation (2) for foreign exchange and interest rate term structure returns for various countries. The spot and one month forward exchange rates are sampled at the end of the month from Citicorp Database Services for the period 1975 to 1989.[11] In particular, we examine the exchange rate for the U.S. dollar against the Deutschemark, the British pound, and the Japanese yen.

The interest rate data are Eurocurrency deposits from Harris Bank. From these data, we sample one month spot rates on deposits and the forward rate on a one month deposit for one month ahead.[12] In keeping with the countries above, we chose interest rates denominated in the currencies for Germany, the U.K., Japan, and the U.S. Some of these deposits were available for earlier periods than for exchange rates.

1.3 Estimation Method and Results on Excess Returns

With the data on spot and forward rates in foreign exchange and bond markets, we estimated equation (2) and tested the restriction that $a_1 = 1$. Various methods have been introduced for the purpose

[11]These data were kindly provided by Geert Bekaert and Robert Hodrick. The data correctly sample the correct future spot rate to correspond to the current forward rate. See Bekaert and Hodrick (1991).

[12]Using the linearized term structure relationship from Campbell and Shiller (1991) for the case of pure discount bonds as we have here, the forward rate on an *k* period bond contracted for trade in *n* periods is: $[(k+n) R^{k+n}_t - n R^n_t]/k$, where R^j_t is the rate on a j period deposit at time *t*. In this paper, we only consider the case where k = n for one month and three month deposits, so that $F_t^k \equiv [2k R^{2k}_t - k R^k_t]/k$ for $k = 1, 3$.

of estimating cointegrating regressions such as these and providing statistical inference. Several econometric issues arise. First, we must obtain the correct standard errors on the coefficients since the asymptotic distribution of a_1 is non-normal, invalidating OLS standard errors. Another empirical problem is that OLS estimates of the parameters in cointegrating regressions are known to be biased in finite samples. Therefore, an adjustment for this bias is necessary to estimate our equations.

For robustness, we estimated these equations individually using the methods described in Stock and Watson (1989) and that in Hansen and Phillips (1990). In order to improve efficiency of estimation, we also estimated groups of equations where the residuals were most likely to be correlated. Specifically, using the Stock-Watson method, we estimated jointly the three exchange rates, and groups of three equations for the U.S. interest rate, a dollar-foreign currency exchange rate, and the foreign interest rate. Details of this estimation method are provided in the appendix.

Table 1 provides the results for the joint equation estimates for various groups of returns.[13] Each panel provides the bias-adjusted point estimates of a_1 for each return along with the p-values in parentheses for the hypothesis that the coefficient is significantly different than one.

The entries in parentheses are the p-values for the hypothesis that the bias-adjusted a_1 equals one. The second and third rows are constructed assuming the residual in the Stock-Watson regressions follow an MA(3) and MA(6), respectively.[14] Each panel also reports the joint test that all the a_1 coefficients in the system are jointly equal to zero.

As these results demonstrate, the joint hypothesis is rejected for 3 month returns at the 5% confidence level for all four systems. At the one month horizon, the evidence is more mixed. For the foreign exchange system and the combined U.S. and Japanese interest rate returns system, the hypothesis is also rejected at the 5% confidence level. On the other hand, the p-values are larger at near 10% for

[13]We found similar implications for rejections of the null hypothesis based upon single equation estimates for the interest rates and the dollar-yen rate. However, the dollar-pound and dollar-DM equations were imprecisely estimated and did not reject the hypothesis in some cases.

[14]The Stock-Watson method requires regressing the level of the dependent variable on the level of the regressor as well as leads and lags of the first-difference of the regressor. All p-values for the one-month and three-month estimates assume the number of leads and lags are three and six, respectively. The overall results were robust to changes in these particular specifications.

the systems including British interest rates in Panel B and German interest rates in Panel C.

Many of the coefficients on individual returns are significantly different than one as well. Generally speaking, the foreign exchange returns on the yen, the 1-month British pound, and term structure returns on British, German, and especially Japanese deposits differ significantly from one. Furthermore, the hypothesis on the interest rate coefficients tend to be rejected for longer horizons.[15]

Two related features of these results are particularly noteworthy. First, the evidence that excess returns contain permanent disturbances may seem surprising since researchers have not seen evidence of these disturbances before. In Evans and Lewis (1990), we question why permanent components in excess returns from the U.S. bond market have not been previously detected. We conducted Monte Carlo experiments where the true cointegrating coefficient on forward rates were close to but not equal to one. Generating excess returns that contained these unit roots by construction and replicating the series 1000 times, we found that standard Dickey-Fuller tests would always reject a unit root. Furthermore, the first-order autocorrelation coefficients were typically less than .10 and hence far away from an autocorrelation of 1.

In general, the closer the cointegrating coefficients are to unity, the smaller is the variance of the permanent shock component in excess returns and the less detectable is this component.[16] To see this relationship, consider the sample variance of excess returns based upon the decomposition from the cointegrating regression given in equation (3). Taking this sample variance implies:

$$\mathrm{Var}(S_{t+k} - F_t^k \mid \underline{X}_T) = (a_1 - 1)^2\, \mathrm{Var}(F_t^k \mid \underline{X}_T) + \mathrm{Var}(v_t \mid \underline{X}_T),$$

where $\mathrm{Var}(\ \mid \underline{X}_T)$ is the variance conditional upon the sample of T spot and forward rates observations, $\underline{X}_T$. The sample variance shows that, even though the variance of the forward rate F_t is growing, its

[15]Using the entire post-War sample period, Evans and Lewis (1990, 1991) find strong rejections of the null hypothesis for U.S. interest rates. Using the shorter sample period in the present paper, we do not reject this null for U.S. interest rates.

[16]Thus, this evidence comprises an example of near "observational equivalence" between a process without a unit root and one with a very small unit root component. This "observational equivalence" is discussed in Campbell and Perron (1991).

contribution to the total variance in the sample depends upon $(a_1 - 1)^2$. Since this value is a small number when a_1 is close to one, the impact of the permanent shock upon the total variance of excess returns may be very small.

Therefore, permanent shocks in excess returns would be consistent with standard treatment of excess returns as stationary only if a_1 were very close to one. In this regard, it comforting to note that the point estimates in Table 1 are all close to one. Even though we find evidence of non-stationary components, these components are empirically small and would be difficult to detect using standard techniques.

The second noteworthy feature of the results is that the point estimates are very close to one. Although we would expect this finding in light of standard treatment of excess returns, the proximity of these estimates to their values under the null hypothesis raises the question: how important are these small deviations from one for explaining excess returns? We will address this question in section (1.5) below. Before examining this issue, however, we will first evaluate the robustness of the results in Table 1.

1.4 Estimation Results Using Contemporaneous Spot Rates

Table 1 demonstrates the striking result that not all excess returns are stationary. Surprisingly, the returns on most interest rates and the dollar-yen exchange rate appear to contain small permanent shocks. One may be concerned, however, that we have spuriously related transitory "high frequency" components to permanent "low frequency" components. If so, then the results would differ substantially if we used other dependent variables with the same permanent shocks but different short run dynamics.

Therefore, to check robustness we also considered another measure of the permanent shock, "low frequency" component, of these asset prices. We substituted the contemporaneous spot rate for the future spot rate as the dependent variable in equation (2) to produce a similar cointegrating regression. To see why, note that by (3), $S_{t+k} = S_t + I(0)$ terms for both $k = 1$ and 3 months. For the case of $k = 1$, we can therefore rewrite equation (2) as:

(6) $S_t = b_0 + b_1 F_t^1 + w_t$, where Ho: $b_1 = 1$,

for the hypothesis of no stochastic trends in excess returns, such that $b_0 + w_t = rp_t + \epsilon_{t+k} - \Delta S_{t+1}$.

The null hypothesis in (6) parallels that of (2). Comparing the constant and residual in the two equations, $b_0 + w_t$ with $a_0 + v_t$, reveals that they are equivalent except for ΔS_{t+1}. Since the change in the exchange rate is stationary by definition, the intuition for the null hypothesis in the cointegrating regression (6) is identical to that in the regression (2). Specifically, if risk premia and forecast errors are stationary, then they must be asymptotically independent of the I(1) right-hand side variable, F_t.

Table 2 reports the results from a cointegrating regression of equation (6) for the same systems as in Table 1. The estimates for the 3-month ahead forecast continue to be generally rejected at about the 5% marginal significance level. The marginal significance levels for the 1-month ahead forecast horizon systems increase slightly in some cases. However, the exchange rate system and the Japanese return system still provide strong rejections of the null. Overall, the results appear consistent with the evidence in Table 1, suggesting that the estimates indeed capture the long run dynamics between spot and forward rates.

1.5 How Important Are the Permanent Components in Excess Returns?

The results in Tables 1 and 2 reject the hypothesis that spot and forward rates vary in a one-to-one proportion in the long run. Instead, they move together according to a proportionality coefficient very close to one. This result leads naturally to the question: how important is this finding for the behavior of excess returns?

To address this question, we estimated a time series model for the excess returns that decomposes their movements into a permanent and transitory component. We parameterized the permanent I(1) component to be consistent with the estimates of b_1 in Table 2. The parameters of the transitory I(0) components were estimated with maximum likelihood.

Specifically, we let the excess returns follow the process,

$$S_{t+k} - F_t^k = z_o + z_{1,t+k} + z_{2,t+k},$$

where

$$z_{1,t} = z_{1,t-1} + e_t, \qquad z_{2,t} \sim \text{ARMA}(p,q), \qquad \text{and } z_o \text{ is a constant.}$$

We obtain an estimate of the variance of e_t from $(b_1 - 1)^2$ times the sample variance of the forward rate.[17] The only parameters to be estimated are the constant, z_o, and the coefficients in the ARMA process, characterizing the transitory component of excess returns. In order to identify the ARMA component for each excess return, we estimated all of the combinations of the autoregressive and moving average orders up to an ARMA(2,2). We then selected the best model on the basis of Akaike's information criterion. Using the parameter estimates from this model we then backed out the permanent and transitory components, $z_{1,t}$ and $z_{2,t}$. In this way, we identified the permanent shocks to excess returns that contain the same variance as our cointegrating regression estimates imply.

Figure 1 depicts the estimates of the I(1) component of the 3-month foreign exchange excess returns together with the realization of the spot exchange rate. Note that this component is predictable from time t information. As the figures show, the permanent shock component in excess returns exhibits considerable swings, particularly in the Yen/dollar excess returns. Furthermore, for all three cases, the contribution of the small random walk component in excess returns are quite sizeable over some periods. The I(1) component of the £/\$ excess return ranges between about 2% and -4%, while the range for the Yen/\$ return is 6 to -6%. Therefore, although the coefficient estimates on the exchange rate forward rates in Tables 1 and 2 are close to one, they imply sizeable *ex ante* predictable components of excess returns.

Another interesting feature of these returns is that they follow the basic pattern of the spot exchange rate. Figure 1 depicts the exchange rate for comparison. As the figure shows, the I(1) component in returns tend to be positive and increasing as the dollar increased in value during the early 1980's. Interestingly, the returns peak near the height of the dollar's strength in 1985, and then fall significantly.

We calculated similar estimates for the interest rate excess returns examined in Tables 1 and 2. Figure 2 illustrates the three month excess returns for the Eurodollar deposits.[18] The estimates of the I(1) component in returns displays considerable serial correlation. The figure also shows the 3-month

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[17]Technically, we set $Var(e_t) = (b_1 - 1)^2 \; Var_o(F_t \mid \underline{X}_T)$, where $Var_o(\mid \underline{X}_T)$ is the estimate of the variance of the forward rate at frequency zero based upon the sample $\underline{X}_T$.

[18]Estimates for the other currency deposits are not reported in order to conserve space.

Eurodollar interest rate over the period. As the estimates show, the I(1) component in excess returns tended to be higher during the high U.S. interest rate period in the early 1980's, and then dropped dramatically with the fall in these rates during 1982-83.

Overall, the estimates described in Figures 1 and 2 show that the I(1) component in excess returns of the size implied by the cointegrating regression estimates are relatively large. Although the point estimates of the cointegrating coefficients in Tables 1 and 2 are very close to one, the implications for the excess returns can be sizeable. Later, in the next section, we will evaluate these estimates in light of different explanations of the permanent shock components to excess returns.

Section 2: Risk Premia, Noise Traders, or Peso Problems that Trend with Rates?

Above we found surprising evidence that at least some excess returns on foreign exchange and long deposit holdings are affected by shocks that appear permanent. In this section, we consider the possible sources of these findings. A rejection of the hypothesis that the coefficient on the forward rate equals one implies that some component of excess returns contains a unit root permanent shock. Since excess returns are comprised of risk premia and forecast errors, these rejections mean that either risk premia or forecast errors are affected by permanent shocks.

We begin this section by addressing whether risk premia are the source of this rejection. We then consider the possibility that forecast errors are responsible for rejecting the null. We examine the effects upon forecast errors when, first, different agents including irrational traders are present in the market, and, second, infrequent discrete events induce peso problems in rational forecasts.

2.1 Permanent Shifts to Risk Premia?

Since conventional assumptions require forecast errors to be stationary and uncorrelated with all current information, one would naturally think that any permanent shocks must necessarily come from the risk premium. In this vein, we first ask whether standard risk premium models could explain these results.

Models of the risk premium in asset returns generally depend upon the second moments of returns. Therefore, risk premia could theoretically contain unit root components if the variability of second moments were sufficiently persistent. For example, suppose that the risk premium on excess returns depended upon the variance of these returns.[19]

$$(7)\quad S_{t+1} - F_t = c_0 + c_1\,\sigma_t^2 + u_{t+1}, \qquad \text{where } \sigma_t^2 = E_t u_{t+1}^2.$$

Researchers have examined risk premium models like (7) conditioned upon a particular process for the variances such as ARCH-in-mean. If risk premia behave in this way, then returns may contain a unit root because shocks to conditional variances are very persistent.

Alternatively, general equilibrium asset pricing models relate the risk premium to the covariance between consumption and returns.[20] In these models, the persistence in risk premium disturbances depends upon the persistence in disturbances to consumption covariances and the risk-free rate. In addition, these covariances may be related to the variances of returns. For example, Lewis (1991b) cannot reject the hypothesis that the conditional covariance between consumption and excess returns moves in proportion to movements in variances. Therefore, the persistence of shocks to risk premia will depend upon the persistence of shocks in conditional variances of asset returns.

Thus, according to either type of model, the risk premium will appear to contain a unit root if conditional variances are highly persistent. Recent studies have examined the behavior of conditional variances of exchange rates and interest rates in some detail. For example, Baillie and Bollerslev (1991) use daily and intra-daily exchange rate movements to examine the behavior of conditional heteroscedasticity. They find no evidence that the variances of exchange rates are affected by permanent disturbances, processes that are so-called "integrated in variance". Similarly, studies of the conditional variances of bond returns find that these variances contain mean-reverting disturbances.

Evidence in the literature thus suggests that the variances of term structure and foreign exchange returns are not persistent enough to generate unit root components in the risk premia under standard models of risk premium. Hence, below we will assume that the risk premium, rp_t, is an I(0) covariance

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[19] Risk premium models such as these have been examined by French, Schwert, and Stambaugh (1987) among others.

[20] See Hansen and Hodrick (1983) and Backus, Gregory, and Zin (1989), for example.

stationary process and will examine the other component of excess returns.

2.2 Permanent Shocks to Forecast Errors?

We now turn to explanations based upon expectations. Two versions of systematic expectational errors have recently been discussed in the literature and will be considered below. The first version argues that the market includes heterogeneous types of traders and some may make systematically incorrect forecasts. For example, some traders may try to follow trends in asset prices, traders that are called "feedback traders" or "chartists" in the literature. Others trade on the anticipation that asset prices will move back towards the level implied by some fundamentals. These traders are therefore called "fundamentalists". Cutler, Poterba, and Summers (1990) and Frankel and Froot (1986) describe the interactions of these types of traders. For convenience below, we will refer to all traders who do not form forecasts according to rational expectations as "noise traders".[21] The presence of both these types of traders will tend to bias the market-determined measure of the expected future spot rate inherent in the forward rate.

A second version of market expectations presumes that these forecasts are formed rationally, but that traders believe the determinants of asset prices may shift discretely at infrequent intervals. For example, infrequent changes in tax policy may affect the implicit discount rate faced by traders, thus affecting the way they evaluate the future. Alternatively, the process of fundamentals that affect the asset prices may itself shift discretely. Examples include shifts in monetary operating procedures or structural shifts in income that affect the expected future values of asset prices and, hence, the current asset price. In this case, even fully rational agents may make systematic forecast errors within a given sample due to the so-called "peso problem."[22]

In either case, market forecasts of asset prices may be systematically biased within the sample. Furthermore, as we will demonstrate below, these forecast errors can appear to contain the same

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[21]For discussions of "noise trading" in particular, see Black (1986), Delong, Schliefer, Summers, and Waldman (1990b) or Schleifer and Summers (1990).

[22]Rogoff (1980) first pointed out the "peso problem", the market's rational belief of a discrete event that does not materialize in the sample. Similar effects arise when the market is rationally learning as described in Lewis (1989).

permanent disturbances as the forward rate. Forecast errors that behave in this way could generate the empirical results in Tables 1 and 2.

Below, we will show how biased forecasts arising from either noise traders or peso problems would affect the relationship between spot and forward rates. Furthermore, we will provide examples of the form that these expectations must take in order to generate our empirical findings above. We will continue to focus upon the one month ahead ($k = 1$) case and will subsume the superscript on the forward rate.

Noise Traders: Suppose that some noise traders in the market are irrational. Their forecasts of the future spot rate systematically deviate from the *ex post* realized rates, possibly by only a small amount. Furthermore, suppose that the market contains an equilibrium share of these traders, denoted by n.[23] For now, we will follow De Long, et al (1990b) and assume that the market has a constant steady state share of these traders, although we will later relax this assumption. In this case, an aggregate market measure of the expected future spot rate is:

$$(8)\quad E_tS_{t+1} = (1 - n)\, E_t\, S_{t+1}{}^r + n\, E_tS_{t+1}{}^*$$

where S^r and S^* is the process that rational traders and "noise" traders, respectively, believe asset prices are following.

An example may serve to clarify this relationship. Suppose that some agents in the market believe that asset prices will revert to a long-run equilibrium value but also base their expectations upon the current exchange rate. Frankel and Froot (1987) refer to these expectations as "regressive expectations", given by:[24]

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[23]For a discussion of how such an equilibrium share may arise with both noise traders and rational traders earning returns, see De Long, Schleifer, Summers, and Waldman (1990b).

[24]Frankel and Froot (1987) use survey expectations as a measure of the total market expectations and define this relationship as in equation (9) below. None of our essential results below would be changed if we defined equation (9) as the market expectations rather than the "noise trader" expectations. We follow the notation as in the text for expositional simplicity only.

(9) $E_tS_{t+1}^n = (1 - g) S_t + g S_o$,

where S_o is a long run asset price level and g is the weight assigned to this level. Using survey data for foreign exchange, they find that these expectations are biased and that g is significantly different than zero. They consider alternative cases where S_o is (a) a constant, and (b) the ratio of price levels. In either case, the long run value is a stationary process.

Thus, we could write these expectations more generally as:

(9') $E_tS_{t+1}^n = (1 - g) S_t + I(0)$ "Noise Traders".

All expectations that weight the current trend in the asset price in a proportion different than one will have a similar form.

On the other hand, rational agents' forecasts will be equivalent to the mathematical expectation of the future spot rate conditioned upon the observed process. Thus, rational agents' forecasts can be written as the mathematical expectation of the true process as $E_t S_{t+1}^r = E_t S_{t+1} = S_t + I(0)$.

Substituting these rational and noise trader expectations into (8), we obtain the market forecast of the future spot. Further substituting this forecast into the definition of the forward rate in (1) and rearranging gives the forward rate in terms of the current spot rate.[25]

(10) $F_t = E_t^m S_{t+1} - rp_t = (1 - n) E_t S_{t+1}^r + n E_tS_{t+1}^n - rp_t$

$$= (1 - ng)S_t + (1 - ng)E_t\Delta S_{t+1} - rp_t = (1 - ng)S_t + I(0) \text{ terms.}$$

[25]While under conventional models, the forward rate contains the market's expected future spot rate, it is not clear without a specific model how the forward rate would be determined in the presence of noise traders. For example, if rational agents realized that noise traders were incorrectly forecasting the spot rate, they may want to trade on only one side of the forward contract transaction. Without finite horizons, rational agents may try to take infinite positions against the noise traders so that no equilibrium would exist. On the other hand, assuming an equilibrium exists and that both types of agents' expectations are incorporated into the forward rate, a relationship as defined in (10) would obtain where n measured the equilibrium effect of noise traders in the determination of the forward rate. We will use this interpretation below.

Given this relationship between forward rates and expectations, we now consider the implied cointegrating expression in (6) between the current spot rate and the forward rate. As we showed above, the coefficient on the forward rate, b_1, provides an estimate for the vector $X_t \equiv [S_t, F_t]'$ of the normalized cointegrating vector, $\alpha \equiv [1, -b_1]'$ such that $\alpha' X_t$ is a stationary I(0) process. Equation (10) reveals that these types of traders would induce the forward rate to trend with the spot rate in proportion to (1 - ng). Thus, the cointegrating vector should yield, $b_1 = (1 - ng)$.

Table 3 illustrates the difference between b_1 under standard rational expectations and in the presence of some traders who do not form expectations rationally. While standard rational expectations implies that b_1 must equal one, the presence of noise traders will make the coefficient differ from one by the term *ng*. Thus, the size of the deviation from one will depend upon first, the importance of noise traders in the market-determined forward rate through *n*, and second, the amount of their weight on regressive expectations through g. As this makes clear, the presence of noise traders who mispredict the impact of the stochastic trend upon the future asset price would cause the cointegrating coefficient to deviate from one.

While we follow the literature above and assume that the share of noise traders in the market is constant, in principle this share could vary over time. Therefore, in Table 3, we also report this possibility in parentheses, allowing n_t to vary with time.

Peso Problems: An alternative interpretation with similar empirical implications arises when the underlying determinants of asset prices experience infrequent, discrete shifts. When these shifts occur, they also alter the behavior of asset prices. Therefore, rational agents incorporate the effects of alternative processes into their expectations about future asset prices.

To be more concrete, suppose that the market is rational but believes that the process followed by asset prices may switch to an alternative process. We may write expectations in this case as:

$$(11)\quad E_t S_{t+1} = (1 - \lambda_t)\, E_t\, S_{t+1}^{C} + \lambda_t\, E_t S_{t+1}^{A}$$

where S^C is the current process that asset prices are following, S^A is the alternative process that traders believe prices may follow in the future, and λ_t is the market's assessed probability that the process will switch. Thus, equation (11) is a probability-weighted average of the expected future spot rates conditional upon each process.

Expectations in the form of equation (11) could arise in two ways. First, traders may anticipate that the process of fundamentals may discretely change in the future. For example, Lewis (1991a) found that bond traders anticipated a shift in Federal Reserve operating procedures during the non-borrowed reserves targeting period from 1979 to 1982. In this case, S^C is the spot rate realized from the current process while S^A is the rate that would be generated by the alternative expected process. Second, traders may be learning about a change in past policy. For example, if the nature of monetary policy changed in the past but traders were unsure whether the change was permanent, they would observe the new process for money over time. In this case, the exchange rate implied by observations from the new monetary policy regime would be the current S^C process while the exchange rate based upon the past would be S^A. In this case, λ_t is the probability the market associates with the old monetary regime, while $(1 - \lambda_t)$ is their assessed probability of the new monetary regime. Based upon observing the new regime, market participants would update their probability of the old regime generating the money supply. Eventually, rational agents would learn the spot rates are generated by the new process and λ_t would disappear.

Therefore, either from anticipated future shifts or learning about past shifts in the process of prices, expectations by rational market traders would take the form in equation (11). Suppose further that the alternative expected process trends with the current process. Then, we can relate the market's expectations conditional upon the current process and the alternative process with the mathematical expectation of the process based upon realizations from the current process. For descriptive purposes, we will assume that the spot rates are realized from the current process throughout the sample, although the substantive results below do not depend upon this assumption.

Expectations conditional upon each process can then be related to the actual process. Forecasts conditioned upon the current process as observed in the sample equal the mathematical expectation of the realized process. Thus, $E_t S_{t+1}^c = E_t S_{t+1} = S_t + I(0)$ terms. Rational agents also condition their expectations on the alternative process they believe asset price determinants may follow in the future.

To suggest the types of alternative expected processes that would explain the results in Tables 1 and 2, consider the following example. Suppose that asset prices conditional upon each regime depend upon the discounted present value effect of future fundamentals defined as Y_t.[26] Thus, asset prices conditional upon the current regime depend upon the expected future path of fundamentals according to:

(12a) $S_t^C = \theta^C Y_t$,

where θ measures the effect of future fundamentals upon the asset price. In this case, the permanent shock component in asset prices arises from the permanent shock component in fundamentals, Y_t. Suppose also that people expect a change in regime that will affect the structural relationship between fundamentals and asset prices. In other words, market participants believe θ may change, for instance from a change in tax policy that alters the perceived discount rate of future fundamentals. Alternatively, traders may believe that the expected future fundamentals process may shift. For example, market-determined long term bond rates may incorporate an anticipated shift down in short term interest rates through a switch in monetary policy. In either case, the asset price conditional upon the alternative regime will be given by:

(12b) $S_t^A = \theta^A Y_t$.

Therefore, the expected future asset prices under each regime will be given by:

(13) $E_t S_{t+1}^i = \theta^i E_t Y_{t+1}$.

Substituting the expectations conditioned on each regime from (13) into the total forecast given in (11), we obtain the market forecast of the future spot. Further substituting this forecast into the definition of the forward rate in (1) and rearranging gives the forward rate in terms of the current spot

[26]For example, a standard model for exchange rates is: $S_t = \phi E_t \Sigma_{j=0}^{\infty} (1 - \phi)^j \beta y_{t+j}$, where y_τ is the combined effects of fundamental determinants of exchange rates at time τ, and β measures the effect of fundamentals upon the exchange rate. Using the definition in the text,
$Y_t \equiv \Sigma_{j=0}^{\infty} (1 - \phi)^j y_{t+j}$, so that $S_t \equiv \phi\beta E_t Y_t$. A similar relationship can be derived for bond returns.

rate.

$$(14)\quad F_t = E_t^m S_{t+1} - rp_t = (1 - \lambda_t\gamma)\, S_t + (1 - \lambda_t\gamma)\, E_t\Delta S_{t+1} - rp_t$$
$$= (1 - \lambda_t\gamma)S_t + I(0) \text{ terms},$$

where $\gamma = (\theta^C - \theta^A)/\theta^C$.

Given this relationship between forward rates and expectations with peso problems, we may now reconsider the implied cointegrating expression between the current spot rate and the forward rate given in equation (5). As we have shown repeatedly, the coefficient on the forward rate, b_1, provides an estimate such that $S_t - b_1 F_t$ is I(0) stationary. Inspecting (14) reveals that such a coefficient is: $b_1 = (1 - \lambda_t \gamma)$.

In Table 3, we provide a comparison of this case with the standard rational expectations case and the noise trader case. This comparison shows that the deviation of b_1 away from one arises from the presence of the peso problem; i.e., $\lambda_t \gamma$. In this case, the size of the deviation will depend upon first, the probability of realizing the alternative distribution through λ, and second, the difference between the conditional mean based upon the current relative to the alternative distribution through γ. As this makes clear, the presence of a peso problem could make the cointegrating coefficient deviate from one.

2.3 Interpretting the Permanent Shocks in Excess Returns as Forecast Errors

Given the description of how forecast errors can generate the results in Tables 1 and 2, we can now evaluate the permanent shock components depicted in Figures 1 and 2 in light of these explanations. In this case, the permanent shock component will arise from persistent mistaken expectations in forecasts. Under both the noise trader and peso problem expectations, the nature of the forecast can be inferred from the permanent shock component.

Noise Traders: Consider first the market expectations in the presence of noise traders defined in equation (8). To transform these expectations into forecast errors, subtract the realized future rate from both sides of the equation, yielding:

(15) $(S_{t+1} - E_t^m S_{t+1}) = (S_{t+1} - E_t S_{t+1}^r) + n(E_t S_{t+1}^r - E_t S_{t+1}^n).$

Since the expectations by rational traders equals the true mathematical expectation of the future spot rate, $E_t S_{t+1}^r = E_t S_{t+1}$ and therefore, the first component on the right-hand side of equation (15) is a white noise forecast error. This term is clearly I(0) stationary. However, the second term captures the difference between the rational forecast and the noise traders' forecast of the future spot rate. As described above, since noise traders are forecasting a variable with permanent disturbances, a bias in their forecasts will appear in excess returns as a small component with permanent shocks. This component is simply $n(E_t S_{t+1}^r - E_t S_{t+1}^n)$. Clearly, this permanent shock component will be positive when $E_t S_{t+1}^r > E_t S_{t+1}^n$ and vice versa. Furthermore, the magnitude of this effect depends upon how important noise traders are in the market, measured by n, and the bias in their forecasts, measured by the difference in their forecasts from the rational forecast.

In light of this interpretation, consider again the permanent shock component in excess returns illustrated in Figures 1 and 2. For the exchange rate excess returns, this component is mostly positive during the run up in the value of the dollar through 1985. This observation together with equation (15) implies that noise traders systematically under-predicted the strength of the dollar. Some of these traders persistently conditioned their forecasts on a long-run value of the exchange rate. As the figure shows, this relationship was reversed after the fall of the dollar in 1985 with the permanent component of excess returns becoming negative, implying "noise traders" expected a weaker dollar than was rational.

A similar pattern can be found in the excess returns on U.S. interest rates as described in Figure 2. During the dramatic increase in U.S. interest rates beginning in the fall of 1979, the permanent component in excess returns became positive, suggesting that the rationally expected interest rate was higher than the noise traders' forecasts. This bias becomes quite negative, however, with the fall in U.S. interest rates in 1982 consistent with noise trader forecasts of higher interest rates than are rational.

Peso Problem: Alternatively, the permanent component in excess returns could be capturing the market's rational beliefs about changes in policy regimes that affect exchange rates and interest rates. To see the relationship between these rational forecast errors and the permanent component in excess returns, subtract the market expectations in equation (11) from the realized future spot rate. Consider first the case where the future spot rate is a realization from the current "C" process, defined as S_{t+1}^C.

(16) $(S_{t+1}^C - E_t^m S_{t+1}) = (S_{t+1}^C - E_t S_{t+1}^C) + \lambda_t(E_t S_{t+1}^C - E_t S_{t+1}^A)$.

Clearly, the first component on the right-hand side of equation (16) is white-noise and therefore I(0) stationary. However, the second component depends upon the difference between the market's expectations conditional upon the current regime, C, and the alternative regime, A. As discussed above, this component can appear to contain permanent disturbances when, for example, the expected shift in policy will change the market's discount rate or will produce a shift in fundamentals. In this case, the permanent component incorporates $\lambda_t(E_t S_{t+1}^C - E_t S_{t+1}^A)$. This depends upon both the market's assessed probability of a switch in the regime, λ_t, and the difference between the spot rate conditional upon the two regimes. Clearly, this component will be positive when the expectations conditional upon the current process are greater than the expectations conditional upon the alternative process and vice versa.

Since shifts in the regime may have occured in the sample, consider the forecasts when the alternative process is realized but traders believe that the process may switch back to the "C" regime. Defining the probability of a transition from the "A" regime to the "C" regime as λ_t, the analogue to (16) may be written:

(16') $(S_{t+1}^A - E_t^m S_{t+1}) = (S_{t+1}^A - E_t S_{t+1}^A) + \lambda_t(E_t S_{t+1}^A - E_t S_{t+1}^C)$.

As in (16), the forecast error will be stationary except for the deviation between the expected spot rate conditional upon two regimes. In both cases, the permanent component will be positive if the process truly generating the realizations has larger conditional expectations than the other regime. Therefore, it is not restrictive to discuss the results in Figures 1 and 2 in light of equation (16) alone.

Consider now the permanent shock components to excess returns to foreign exchange depicted in Figure 1. As these figures show, the permanent shock component increased through the early part of the sample and peaked with the height of the dollar in early 1985. Equation (16') suggests that this behavior arose from expectations of a switch to a weaker dollar than was realized *ex post* during the sample. Since the magnitude of this component varied over time, either the probability of this switch or the difference in expected spot rates conditional upon the two regimes fluctuated considerably over time. For most currencies, this peso problem effect peaked in early 1985 and then plummetted with the subsequent fall in the dollar. The large negative permanent components in 1986-87 would suggest that the regime switched and that traders conditioned their forecasts upon beliefs that the dollar might switch to an appreciating regime.

The permanent component in U.S. interest rate excess returns provide similarly plausible interpretations. The positive values of the permanent component following October 1979 implies that the market was conditioning the expected future interest rate on a lower interest rate process than was realized *ex post*. This behavior could arise rationally either because the market did not believe that the Federal Reserve could continue to allow such high interest rates or because the market required time to learn about the new interest rate regime. Both explanations of the term structure of U.S. interest rates over this period has been suggested elsewhere. Therefore, the results implied by our cointegrating regression results provide independent supporting evidence of these plausible stories.

Moreover, whether the permanent shock components in excess returns depicted in the figures arises from peso problems or noise traders or both, these estimates generate behavior that is consistent with popular accounts of the market's expectations of foreign exchange and interest rates over these periods.

Section 3: Are the Effects from Permanent Shocks Constant Over Time?

A comparison of the peso problem to the noise trader case in Table 3 shows that they are similar. Indeed, if the probabilities, λ_t, were constant, then we would find $b_1 = (1 - \lambda \gamma)$. Thus, the probability λ would be analogous to the noise trader effect n, while the difference in expectations γ would be analogous to the bias in noise traders' expectations, g. If agents are rational, however, these probabilities would likely vary over time, implying that the b_1 coefficient would also change over time.

If the peso problem arises from learning about changes in the process or from future changes that occur infrequently in the sample, then we would also expect shifts in these effects over time. As agents learn, λ_t will tend towards zero and the significance of the effects from learning will dissipate. If agents acquire new information indicating that a new policy process is more likely in the future, λ_t may jump up in response to the new information.

Therefore, the parameter b_1 in the cointegrating regressions should vary over time if peso problem effects explain the trend movements in excess returns. By contrast, if an equilibrium share of noise traders influence the forward rate, then the coefficients would be constant over time. A straight-forward test of these two possibilities, then, is to test the parameter stability of the cointegrating regressions above. Hansen (1991) provides a test of time-variation in the parameters in cointegrating regressions. In effect, the test considers whether the coefficient in the cointegrating regression such as (6) should be considered a time-varying coefficient so that:

(6') $S_t = b_0 + b_{1,t} F_t + w_t$,

where the coefficient $b_{1,t}$ has been subscripted with t, denoting its dependence on time.

Table 4 reports the results of these estimates. In the interest of space, we report only the results for the contemporaneous spot rates, comparable to Table 2. The table reports the t-statistics for the hypothesis that b_1 is constant before and after January 1983, the midpoint of the sample. However, similar results obtained when we tested the constancy before and after other dates, including January of 1980 and 1985. These statistics are based upon individual equation estimates for each of the exchange rates and interest rates. As the results show, the constant b_1 hypothesis is strongly rejected for the German mark, the Japanese yen, and the U.S. and German term structure at both one month and three month horizons. The hypothesis is also rejected for the 3 month Japanese deposit rate.

Thus, if the permanent shocks in excess returns arise from expectations conditioned upon a process that trends with the current process, the relationship varies over time. Peso problem effects are inherently time-varying and therefore seem to be natural candidate explanations. By contrast, a simple story where noise traders' impact upon the equilibrium forward rate is constant can be rejected. Of course, if the share of noise traders' impact upon the market-determined price varied over time, the effect upon b_1 would be time-varying as well as given in parentheses in Table 3. Therefore, the evidence could also be consistent with a time-varying distribution of noise traders.

Concluding Remarks

In this paper, we tested and generally rejected the hypothesis that excess returns in exchange rates and interest rates for four countries are not affected by permanent shocks. We evaluated whether risk premia in conventional models could explain these results, but found that shocks to variances in asset prices were insufficiently persistent. We then showed how the presence of heterogeneous traders or peso problems could generate rejections of the restrictions. A simple heterogeneous trading story states that the relationship between trends in forward rates and spot rates is constant over time, while a peso problem indicates the relationship should be time-varying. Therefore, we tested whether this relationship was constant. The results indicated that the relationship shifted over time, eliminating the simple heterogeneous trading story as a possible explanation for our rejections. On the other hand, a time-varying distribution of trading effects remains a possibility.

An interesting feature of our analysis is that the permanent shock in excess returns may be interpreted as either the "noise trader" or the "peso problem" effect. Estimates of the permanent shock component produced estimates of the "peso problem" in the dollar exchange rate and U.S. interest rates consistent with independent studies based upon a structural model for these effects. Similarly, the implications for "noise traders" expectations matches popular accounts of the market's expectations during the early 1980's. This evidence suggests that the permanent shock component implied by our tests may arise from time-varying effects from heterogeneous traders or from "peso problems" or both.

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Appendix

Estimation Methods for the Cointegrating Regressions

In this appendix, we describe the methods to adjust for simultaneous equation bias in the cointegrating regressions (2) and (9). For more detailed and thorough discussions, see Hansen and Phillips (1990) and Stock and Watson (1989). We will focus upon equation (9) although the same analysis applies to (2) as well.

For notational simplicity, note that equation (9) may be written as,

$$y_t = \gamma x_t + u_{1t} \tag{A1}$$

$$\Delta x_t = u_{2t}, \tag{A2}$$

where u_{1t} and u_{2t} are stationary, $y_t \equiv S_t$, $x_t \equiv F^k_t$, and the constant term is omitted for simplicity. We are interested in testing $\gamma = 1$. Since the forward rate level, x_t, contains a stochastic trend component that is the cumulation of a component of u_{1t}, $\mathrm{Cov}(x_t, u_{1t}) \neq 0$. Therefore, γ will be biased in any finite sample. Hence, test statistics on γ must take account of this bias. Note that this bias arises even though the estimate of γ remains consistent with this simultaneity. We use two methods to adjust for the bias. Below, we give the steps for estimating (A1) and (A2) using each of these methods.

A. Hansen and Phillips method

(1) Estimate (A1) and (A2) to get the estimates of u_{1t} and u_{2t} and the OLS estimate of γ. For future reference, define the vector of residual estimates as: $u_t \equiv [u_{1t}, u_{2t}]'$.

(2) Calculate

$$y_t^+ = Y_t - \Omega_{12}\, \Omega_{22}^{-1}\, \Delta x_t$$

where $\Omega \equiv [\Omega_{ij}] = T^{-1}\Sigma_1^T u_t u_t' + T^{-1}\Sigma_{\tau=1}^{\ell} \omega_{\tau\ell} \Sigma_{t=\tau+1}^{T} (u_t u_{t-\tau}' + u_{t-\tau} u_t')$,

i.e., Ω is the Newey-West estimator of the variance-covariance matrix with ℓ lags of autocovariances and weights ω.

(3) Calculate the "bias-adjusted" estimate of γ as:

$$\gamma^+ = [\, \Sigma_1^T y_t^+ x_t - T\, [I, -\Omega_{12}\Omega_{22}^{-1}][\Lambda_{21}, \Lambda_{22}]'\,][\Sigma_1^T x_t^2]^{-1}$$

where $\Lambda \equiv [\Lambda_{ij}] = T^{-1}[\Sigma_1^T u_t u_t' + \Sigma_{\tau=1}^{\ell} \omega_{\tau\ell} \Sigma_{t=\tau+1}^{T} (u_t\, u_{t-\tau}')]$.

(4) Calculate a modified Wald statistic, known as the G-statistic, to test $\gamma^+ = 1$. This G-statistic is:

$$G_\ell = (\gamma^+ - 1)^2\, [\, \Omega_{11.2} \otimes (\Sigma\, x_t^2)^{-1}]^{-1} \sim \chi^2_{(1)}$$

where $\Omega_{11.2} = \Omega_{11} - \Omega_{12}\, \Omega_{22}^{-1}\, \Omega_{21}$.

Note that subscript ℓ refers to the number of lags included in the estimators Ω and Λ.

B. Stock and Watson Procedure

Rewrite equation (A1) according to,

$$y_t = \gamma_0 + \gamma\, x_t + \beta(L)\, \Delta x_t + u_{1t} \qquad \text{(A3)}$$

where $\beta(L)$ is a polynomial in the lag operator, L, i.e., $\beta(L) = (L^n + L^{n-1} + \cdots + L + 1 + L^{-1} + \cdots + L^{-n+1} + L^{-n})$. The idea to rewriting (A1) in this form is to include as many leads and lags of Δx on the right-hand side of the equation to make u_{1t} independent of x. This implies that the asymptotic distribution of the OLS estimator of γ is normal. Intuitively, including the leads and lags of Δx_t on the right-hand side "soaks up" the simultaneous equation bias. Note that since u_{1t} will be serially correlated in general, the Wald test of $\gamma = 1$ from (A3) should also use the Newey-West estimator for the covariance matrix. The results reported in the text are not sensitive to these choices, however.

C. Joint Equation Estimation Using Stock-Watson

For the joint equation estimation, we rewrite equations (A1) and (A2) in their stacked equation form:

$$\underline{y}_t = \gamma \underline{x}_t + u_{1t} \qquad \text{(A1')}$$

$$\Delta \underline{x}_t = u_{2t}, \qquad \text{(A2')}$$

where $\underline{y}_t$ is r x 1, $\underline{x}_t$ is (n-r) x 1, and γ is an r x (n-r) matrix. Writing the errors in terms of the primary innovations, we have the representation:

$$u_{1t} = C_{12}(L)\, \epsilon_{2t} + C_{11}(L)\, \epsilon_{1t}$$

$$u_{2t} = C_{22}(L)\, \epsilon_{2t},$$

where ϵ_{1t}, ϵ_{2t} are independent, $C_{11}(L)$ and $C_{22}(L)$ are one sided and $C_{12}(l)$ is a two-sided polynomial matrix in the lag operator. For more details, see Stock and Watson (1988).

Using this representation, we can write (A1') and (A2') as:

$$\underline{y}_t = \gamma \underline{x}_t + d(L)\, \Delta \underline{x}_t + C_{11}(L)\, \epsilon_{1t} \qquad \text{(A4)}$$

$$\Delta \underline{x}_t = C_{22}(L)\, \epsilon_{2t} \qquad \text{(A5)}$$

where $d(L) \equiv C_{12}(L)\, C_{22}(L)^{-1}$.

To estimate (A4), write (A4) as the set of r individual equations, stacked as:

$$\begin{bmatrix} y_1 \\ y_2 \\ \\ y_r \end{bmatrix} = \begin{bmatrix} x_1 & 0 & & 0 \\ 0 & x_2 & & 0 \\ & & & \\ 0 & 0 & & x_r \end{bmatrix} \begin{bmatrix} \delta_1 \\ \delta_2 \\ \\ \delta_r \end{bmatrix} + \begin{bmatrix} e_1 \\ e_2 \\ \\ e_r \end{bmatrix}$$

$$Y = X\delta + e$$

where $y_i = (y_{i1}, \ldots, y_{iT})'$, $x_i = (x_{i1}, \ldots, x_{iT})'$, and $e_i = (e_{i1}, \ldots, e_{iT})'$. Then, defining Ω as a consistent estimate of the long run covariance matrix of e, the estimator of δ is given by:

$$\hat{\delta} = (X'(\Omega \otimes I_T)X)^{-1} X'(\Omega \otimes I_T)Y.$$

The asymptotic variance of this estimator is given by:

$$Var(\hat{\delta}) = (Z'X)^{-1} Z'VZ (X'Z)^{-1},$$

where $Z \equiv (\Omega^{-1} \otimes I_T) X$ and $V \equiv E(e e')$. Thus, Z'VZ can be estimated with the Newey-West estimator allowing for both serial correlation and heteroskedasticity. Notice that our estimator is essentially an instrument variable estimator using Z as instruments. The properties of this estimator are discussed in Hansen and Phillips (1990).

Table 1

Cointegrating Regressions of Future Spot Rates on Current Forward Rates

$$S^i_{t+k} = a^i_0 + a^i_1 F^{i,k}_t + u^i_{t+k}$$

Asset i	$k = 1$ Mo. $\hat{a}_1$ $(H_0 : a_1 = 1)$	$k = 3$ Mo. $\hat{a}_1$ $(H_0 : a_1 = 1)$
A. Exchange Rates		
X^{UK}	1.002	1.003
	(.022)	(.405)
	(.056)	(.313)
X^{Germany}	0.999	0.998
	(.182)	(.281)
	(.180)	(.314)
X^{Japan}	0.998	0.992
	(.009)	(.000)
	(.047)	(.002)
$H_0 : a^i_1 = 1, \quad \forall i$	(.003)	(.001)
	(.012)	(.008)
B. U.S. Bond, U.K. Bond, and Dollar-Pound Rates		
R^{US}	0.995	1.034
	(.706)	(.207)
	(.756)	(.233)
R^{UK}	1.040	1.132
	(.016)	(.001)
	(.018)	(.005)
X^{UK}	1.001	0.997
	(.752)	(.458)
	(.706)	(.491)
$H_0 : a^i_1 = 1, \quad \forall i$	(.102)	(.005)
	(.078)	(.010)

(continued)

Table 1 (Continued)

Asset i	$k = 1$ Mo. $\hat{a}_1$ ($H_0 : a_1 = 1$)	$k = 3$ Mo. $\hat{a}_1$ ($H_0 : a_1 = 1$)
C. U.S. Bond, German Bond, and Dollar-DM Rates		
R^{US}	1.001	1.047
	(.909)	(.100)
	(.496)	(.144)
R^{Germany}	1.009	1.072
	(.370)	(.011)
	(.369)	(.013)
X^{Germany}	0.998	0.995
	(.017)	(.085)
	(.018)	(.081)
$H_0 : a_1^i = 1, \quad \forall i$	(.105)	(.058)
	(.112)	(.029)
D. U.S. Bond, Japanese Bond, and Dollar-Yen Rates		
R^{US}	0.998	1.031
	(.869)	(.234)
	(.930)	(.139)
R^{Japan}	1.039	1.160
	(.029)	(>.000)
	(.022)	(>.000)
X^{Japan}	0.997	0.991
	(.012)	(.015)
	(.021)	(.013)
$H_0 : a_1^i = 1, \quad \forall i$	(.034)	(.001)
	(.050)	(.002)

Notes: S_{t+k}^i are spot rates for asset i at $t+k$ and $F_t^{i,k}$ are forward rates at t for contracts on asset i at $t+k$.

Assets i are R^{US} $\equiv$ U.S. Eurodollar Deposit Rate for k month maturity

R^* $\equiv$ * Country's Euro Deposit Rate for k month maturity

X^* $\equiv$ * Country's Currency Price in terms of U.S. Dollars,

where * = UK, Germany, Japan.

The entries in the top rows in the table are the OLS estimates of a_1. The entries in parentheses are the p-values for the hypothesis that $a_1 = 1$.

The first row assumes the residual after estimating the Stock-Watson (1989) cointegrating regressions has MA(3) while the second row assumes the residual has MA(6).

The p-values for 1 Month and 3 Month horizons assume the number of leads and lags in the Stock-Watson correction are 3 and 6, respectively.

Table 2

Cointegrating Regressions of Current Spot Rates on Current Forward Rates

$$S_t^i = b_0^i + b_1^i F_t^{i,k} + e_t^i$$

Asset i	$k = 1$ Mo. $\hat{b}_1$ $(H_0 : b_1 = 1)$	$k = 3$ Mo. $\hat{b}_1$ $(H_0 : b_1 = 1)$
A. Exchange Rates		
X^{UK}	1.002	1.004
	(.016)	(.156)
	(.039)	(.201)
X^{Germany}	0.999	0.998
	(.217)	(.291)
	(.215)	(.334)
X^{Japan}	0.998	0.994
	(.016)	(.003)
	(.073)	(.020)
$H_0 : a_1^i = 1, \quad \forall i$	(.003)	(.007)
	(.012)	(.032)
B. U.S. Bond, U.K. Bond, and Dollar-Pound Rates		
R^{US}	0.993	1.034
	(.594)	(.160)
	(.545)	(.212)
R^{UK}	1.040	1.136
	(.039)	(.008)
	(.061)	(.025)
X^{UK}	1.001	1.001
	(.736)	(.839)
	(.744)	(.878)
$H_0 : a_1^i = 1, \quad \forall i$	(.160)	(.047)
	(.154)	(.154)

(continued)

Table 2 (Continued)

Asset i	$k = 1$ Mo. $\hat{b}_1$ $(H_0 : b_1 = 1)$	$k = 3$ Mo. $\hat{b}_1$ $(H_0 : b_1 = 1)$
C. U.S. Bond, German Bond, and Dollar-DM Rates		
R^{US}	1.000	1.055
	(.997)	(.040)
	(.669)	(.037)
R^{Germany}	1.008	1.092
	(.423)	(.001)
	(.589)	(.002)
X^{Germany}	0.998	0.996
	(.017)	(.216)
	(.019)	(.230)
$H_0 : a_1^i = 1, \quad \forall i$	(.100)	(.002)
	(.141)	(.010)
D. U.S. Bond, Japanese Bond, and Dollar-Yen Rates		
R^{US}	0.996	1.030
	(.764)	(.210)
	(.957)	(.277)
R^{Japan}	1.034	1.131
	(.066)	(.005)
	(.078)	(.018)
X^{Japan}	0.997	0.992
	(.010)	(.105)
	(.023)	(.090)
$H_0 : a_1^i = 1, \quad \forall i$	(.044)	(.028)
	(.094)	(.075)

Notes: S_t^i and F_t^i are as defined in Table 1.
The entries are the same as described in Table 1.

Table 3
Implied Parameters Under Different Assumptions

Cointegrating Coefficient[a]	Standard Rational Expectations	Noise Traders	Peso Problems
b_1	1	$1 - ng$	$1 - \lambda_t\gamma$
		$(1 - n_t g)$	

Notes: [a]Cointegrating regressions are given by: (5) $S_t = b_o + b_1 F_t + u_t$.

The parameters relate to the expectations models in the text:
$n \equiv$ share of noise traders,
$g \equiv$ noise traders' bias in expectations,
$\lambda_t \equiv$ rational probability at *t* of shift in asset price determinants,
$\gamma \equiv$ difference in expectations of the alternative regime from the current regime.

Parentheses under "Noise Traders" represents coefficient when noise trader effects are time-varying.

Table 4

Subsample Stability Tests for Cointegrating Regressions

$$S_t^i = b_0^i + b_1^i F_t^{i,k} + e_t^i$$

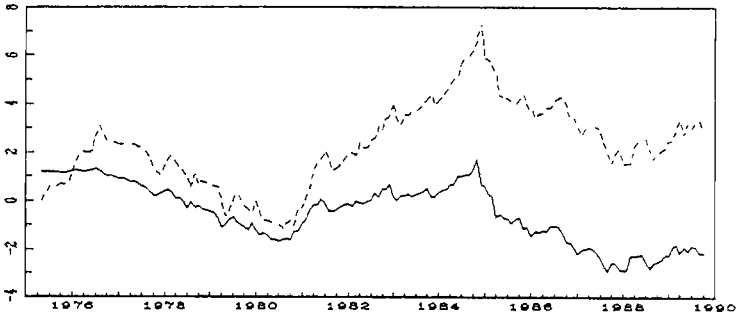
Asset i	Horizon k	$H_0 : b_1$ constant[a] t-statistic
X^{UK}	1	1.18
X^{UK}	3	1.60
X^{Germany}	1	2.65[b]
X^{Germany}	3	2.93[b]
X^{Japan}	1	7.77[b]
X^{Japan}	3	8.18[b]
R^{US}	1	3.57[b]
R^{US}	3	5.21[b]
R^{UK}	1	1.55
R^{UK}	3	1.22
R^{Germany}	1	2.63[b]
R^{Germany}	3	2.61[b]
R^{Japan}	1	1.36
R^{Japan}	3	5.46[b]

[a] Tests the hypothesis that b_1 is the same before and and after January 1983. Reported figures give t-statistic that coefficient on dummy variable equals zero.

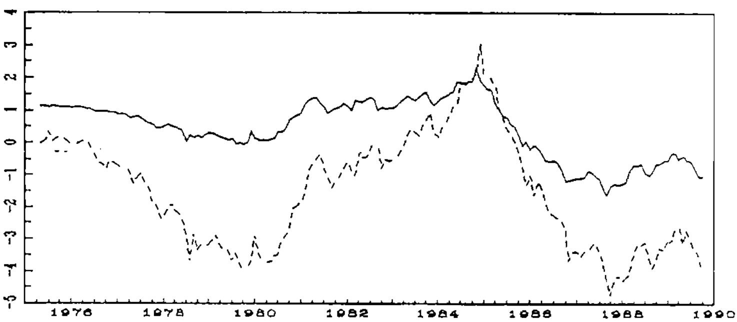
[b] Significant at the 99% confidence level.

Note: All estimates assume six leads and lags of first-differenced regressors in Stock-Watson regression and MA(6) in their residuals. Results are robust to this specification.

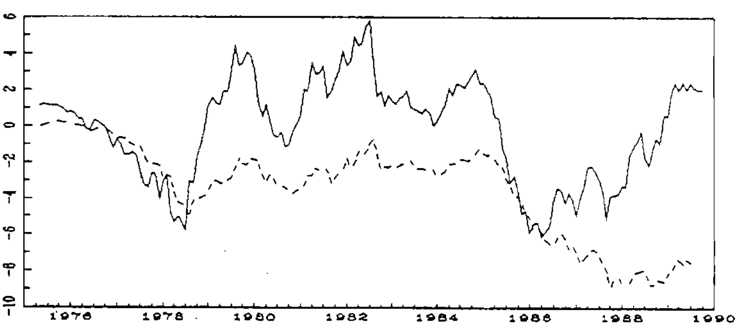
Figure 1



key: ———— I(1) component in £/$ premium (%), - - - log £/$ exchange rate (x10)

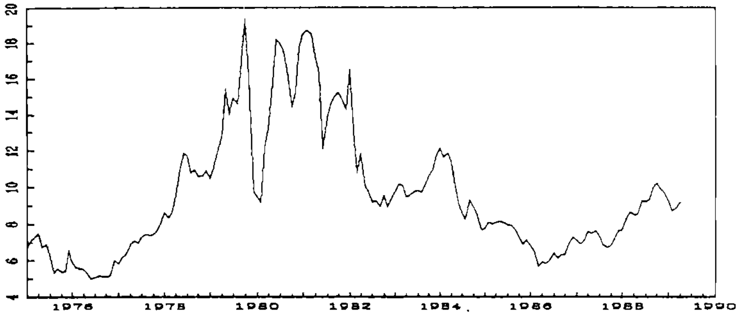


key: ———— I(1) component in Dm/$ premium (%), — — — log DM/$ exchange rate (x10)

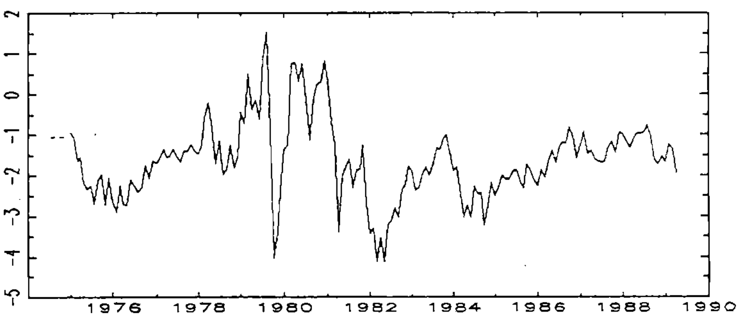


key: ———— I(1) component in ¥/$ premium (%), — — — log ¥/$ exchange rate (x10)

Figure 2



key: 3 month $ Euro Deposit rate (%)



key: I(1) component of the 3 month forward premium on $ Euro Deposit rates (%)