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Imperfect Knowledge, Temporal Instability and an Uncertainty Premium: Towards a Resolution of the Excess-Returns Puzzle in the Foreign Exchange Market*

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Abstract

This paper offers a refinement and explores a resolution of the excess-returns puzzle in the foreign exchange market. We find that the predictions of the forward premium are not negatively biased throughout the three decades of floating, as commonly believed, but rather are sometimes *positively biased,* negatively biased, unbiased or possess no predictive content depending on the subperiod examined. To explain this modified puzzle, the paper makes use of a recently developed model of the risk premium, which we have called an *aggregate uncertainty premium.* Our model employs an alternative approach to modeling exchange rate expectations, dubbed *Imperfect Knowledge Expectations (IKE),* which recognizes that *rational* agents *do* form expectations based on imperfect knowledge. Our model also makes use of a dynamic extension of the assumption of myopic loss aversion. We find that our IKE-based approach can account for the pattern of positive and negative biases estimated over three decades of floating rates.

1 Introduction

The ubiquitous use of the *Rational Expectations Hypothesis* (REH) testifies to the widespread belief among economists that the REH provides *the* solution to the perennial problem of modeling the expectations of *rational* agents. Yet it is precisely in those settings in which agents' expectations play a central role, such as asset markets, that REH-based models encounter their greatest difficulty. In the foreign exchange market, Dornbusch pointed out as early as 1983 that his seminal overshooting model does not explain the observed long swings in exchange rates (Dornbusch (1983)).[1] Ever since, macroeconomists using the REH have grappled with this and other anomalous features of exchange rate movements over the modern floating-rate period.

This paper examines one of the core puzzles in the foreign exchange market, the excess-returns puzzle (also referred to as the forward-premium anomaly). The puzzle stems from a widely reported finding that the estimate of the slope coefficient in a regression of the future change in the spot rate on the forward premium (to be referred to as the Bilson (1981) and Fama (1984) (BF) regression) is significantly negative. Froot and Thaler (1990) report that the average estimate of this slope coefficient is -0.88 in over 75 published articles, indicating that the forward premium is a biased predictor of the future change in the spot rate. As Obstfeld and Rogoff (1989, p.589) put it "if taken literally, the finding that [the slope] is negative, and often significantly so, is startling. It suggests that one can make predictable profits by betting against the forward rate."

The literature has noted two explanations for this negative bias: either the forward premium is correlated with agents' forecast errors and/or it is correlated with a risk premium. Because the former explanation suggests the "irrationality" of expectations (see Froot and Frankel (1989)), it has been largely resisted by the literature. However, despite volumes of research, empirical attempts to explain the negative bias of forward rates as the equilibrium compensation for risk under the *Rational Expectations Hypothesis*

[1] Dornbusch and Frankel surmised that the source of the problem may be the use of the REH. As they put it "The chief problem with the overshooting theory, and indeed with the more general rational expectations approach, is that it does not explain well the shorter-term [long-swings] dynamics (Dornbusch and Frankel, 1988, p. 16)." Excessive fluctuations in other asset markets have also been documented. For an early seminal treatment of such stock price dynamics see Shiller (1981).

(REH) have not succeeded on the whole.[2] Given the universal acceptance of the REH as *the* model of *rational* expectations formation, the literature has not looked for explanations of a negative bias in models of risk premia that do not employ the REH.

This paper offers a refinement and explores a resolution of the excess-returns puzzle. Our refinement is based on an examination of the temporal stability of the BF regression over the modern floating-rate period, using US dollar rates with respect to the British pound (BP), German mark (DM) and Japanese yen (JY). We find this regression to be considerably more unstable than hitherto believed.[3] Our structural change findings include: 1) numerous structural breaks throughout the sample; 2) a significantly negative estimate of the slope coefficient for all three currencies when estimated over the 1980's (despite numerous break points in that decade); and 3) substantial subperiods within the 1970's and 1990's during which the estimate of the slope coefficient is either significantly *greater than one*, significantly negative, insignificantly different from *one* (and zero) and significantly different from one but not zero. Hence, we find that the excess-returns puzzle is not as simple as suggested by the extant literature. What needs to be explained is not a negative bias throughout the three decades of floating, but rather why the predictions of the forward premium are sometimes negatively biased, positively biased, unbiased or possess no predictive content depending on the subperiod examined. If the explanation of this *modified* excess-returns puzzle involves a time-varying risk premium, then any model of such a phenomenon that purports to explain the negative bias in the 1980's must also be consistent with the observed temporal instability and the positive biases found in other subperiods.

We explore a resolution of the excess-returns puzzle, as modified in this paper, that is partly based on a new model of the aggregate risk premium due to Frydman and Goldberg (2001,2002,2003a). Our model of the risk premium employs an alternative approach to modeling exchange rate expectations, dubbed *Imperfect Knowledge Expectations (IKE)*.[4] It also makes use

[2]See Lewis (1995) and Engel (1996) for review articles. One of the most troubling aspects of the empirical record for risk-premium models under the REH are the frequent sign reversals observed in excess returns. See Mark and Wu (1998).

[3]Bekaert and Hodrick (1993), Lewis (1995) and Mark and Wu (1998) report subsample estimates for the 1970's and 1980's and find that the negative slope coefficient is mainly due to behavior from 1980's.

[4]The IKE approach is developed in Frydman and Goldberg (2002). It builds on the

of a dynamic extension of myopic loss aversion, developed in Frydman and Goldberg (2001,2002) and based on Kahneman and Tversky (1979) and Benartzi and Thaler (1995). Our attempt to model the risk premium under IKE, which we call an *uncertainty premium*, has been motivated not only by the inability of risk-premium models under the REH to account for the facts, but also by the diverse set of arguments that have been developed for over two decades suggesting that the REH is very unlikely to characterize, even in a highly abstract and stylized manner, expectations formed by *rational* agents.[5] As Sargent has reminded us quite recently, rather than being a model of individual expectations,

> rational expectations is an equilibrium concept that *at best* describes how the system might eventually behave if the system will ever settle down to a situation in which all of the agents have solved their "scientific problems" (emphasis added, Sargent (1993), p. 23).

The IKE framework recognizes the importance of imperfect knowledge and yet preserves the postulate of individual rationality. As with conventional economic analysis, it supposes that economic agents *do not* ignore exploitable profit opportunities. We show in this paper that in a world of imperfect knowledge this basic rationality assumption leads to an explanation of one of the primary aspects of the modified excess-returns puzzle: the temporal instability of the BF regression. In particular, the *rationality* of agents implies temporal instability of aggregate forecast functions, which in turn, leads to instability in the relationship between the aggregate uncertainty premium and the forward premium and, therefore, to instability of the slope coefficient in the BF regression.[6]

idea of *Theories Consistent Expectations* (TCE), proposed in Frydman and Phelps [1990] and developed in Goldberg [1991] and Goldberg and Frydman [1993,1996a], according to which the extant stock of economic models provides agents with *qualitative* knowledge that can be used in forming individual expectations.

[5]For recent critical analyses of the epistemological and behavioral foundations of the REH and misconceptions concerning the connection between the REH and individual rationality, see Sargent [1993], Kurz and Motolese [1999], Evans and Honkapohja [2001], Frydman and Goldberg [2002] and references therein. For an early discussion of the main problems with the REH, see Frydman [1982], Frydman and Phelps [1983] and Frydman [1983].

[6]As we show in Frydman and Goldberg (2002), our approach to updating of expectations is very different from the so-called "rational learning" models, which we have argued is in general inconsistent with individual rationality in a world of imperfect knowledge.

The IKE-based model of the uncertainty premium also serves as a key ingredient in our explanation of the other important feature of the modified excess-returns puzzle: the pattern of positive and negative estimates of the slope coefficient over three decades of floating rates. This explanation involves three components. First, under IKE and some fairly general assumptions concerning the updating of individual forecast functions, otherwise standard exchange rate models generate long swings – time periods during which the exchange rate moves persistently away from long-run benchmark levels (such as purchasing power parity (PPP)), followed by time periods involving persistent countermovements back.[7] Second, our dynamic model of myopic loss aversion implies that the absolute value of the uncertainty premium on the overvalued currency will rise (fall) persistently during subperiods characterized by a persistent swing in the exchange rate away from (back to) the long-run benchmark level. Third, with IKE, models of the exchange rate imply no permanently stable relationship between swings in the exchange rate and movements of the forward premium, and thus of the differential in nominal interest rates, (see Goldberg and Frydman (1996a) and Frydman and Goldberg (2002)).

To see how this works, consider the most pronounced swing in the DM/$ rate away from PPP during the 1980's, which occurred from the middle of 1981 until the beginning of 1985. Our IKE-based model of the uncertainty premium predicts that the excess return on the U.S. dollar relative to the German mark should be positive and trending up during this exchange rate swing. Since the forward premium happened to be trending down from the end of 1982 until the middle of 1984, we would expect a negative correlation between the uncertainty premium and forward premium during this period. But this implies that the estimate of the slope coefficient in the BF regression should be significantly less than one from the end of 1982 until the middle of 1984. We find that this prediction is consistent with the empirical record: the *ex post* excess return on the dollar was indeed trending up during the 1981-85 period and, despite a small sample, the estimate of the slope coefficient during the 1982-84 subperiod is significantly negative (-19.8!). We also examine the most pronounced exchange rate swing away for the other currencies in the 1980's, as well as the major swing away for each currency during the 1970's

[7]In Frydman and Goldberg [2002] we use IKE to close the monetary model of the exchange rate and show that, contrary to the usual presumption, long swings can occur even if individual forecast functions are *solely* a function of macroeconomic fundamentals.

and 1990's. We find that the three components of our IKE-based explanation can account for the negative biases observed during the swings away in the 1980's for the BP and JY and the positive biases observed for each of the swings in the 1970's and 1990's for all three currencies.[8]

There have been attempts in the literature to explain the excess-returns puzzle (confined to explaining a negative bias) as the consequence of a peso problem arising from the long-swings nature of exchange rate movements. These studies have modeled the structural change and long swings using the Hamilton (1988,1989) Markov switching model.[9] Although we find that swings in the exchange rate and structural change play an important role in explaining the modified excess returns puzzle, we present evidence in this paper that the role of such swings and structural change is quite different than what is implied by these REH-based studies.

A key question, which we leave open in this paper, is the contribution of a time-varying uncertainty premium relative to systematic forecast errors in explaining the modified excess-returns puzzle. As we emphasize in Frydman and Goldberg (2003b), with imperfect knowledge, correlations between aggregate forecast errors and publicly available information, such as the forward premium, will appear from time to time. However, such correlations should *not* be interpreted, necessarily, as evidence that agents' forecasts are irrational. This is because as *rational* agents cope with *imperfect knowledge* and attempt to exploit information in their forecast errors, such correlations experience temporal instability and disappear at *unpredictable* moments of time.[10] Using survey data on exchange rate expectations, we examine the stability of the correlation between forecast errors and the forward premium in Frydman and Goldberg (2003b) and explore the relative contributions of the uncertainty premium and forecast errors in explaining the excess-returns puzzle.

The remainder of this paper is organized as follows. In section 2 we show

[8] Although our IKE model implies that there will also be an uncertainty premium during subperiods involving no apparent swings in the exchange rate, the absence of such swings will be, in general, associated with frequent structural shifts in the relationship between the uncertainty premium and the forward premium. This suggests that such subperiods of no swings will result in an insignificant estimate of the slope coefficient in the BF regression, even if the forward premium is trending.

[9] See, Engel and Hamilton (1990), Baekert and Hodrick (1993), Kaminsky (1993), and Evans and Lewis (1995).

[10] For a formal analysis of this point see Frydman (1982).

that with rational agents and imperfect knowledge, the relationship between the aggregate risk premium and the forward premium, and therefore the slope coefficient in the BF regression, will be temporally unstable. This motivates the structural change analysis presented in section 3, which leads to the refinement of the excess-returns puzzle. In section 4 we explore the potential of the REH approach to model structural change and to explain the pattern of positive and negative slope coefficients in the BF regression documented in section 3. Section 5 outlines our IKE-based model of the uncertainty premium and shows how it can be used in explaining the excess-returns puzzle, as modified in section 3. Finally, section 6 contains concluding remarks.

2 Imperfect Knowledge, Structural Change and the BF Regression

In this section we first reproduce the standard finding of a negative estimate of the slope coefficient in the BF regression when the entire sample period is used. As a way to motivate the structural change analysis of the next section, we then examine the channels through which structural change can arise in the BF regression. We show that temporal instability in the aggregate expectation function, which is implied by our IKE framework, directly implies temporal instability in the relationship between the aggregate risk premium and the forward premium, leading to temporal instability of the slope coefficient in the BF regression.

The BF regression is based on the following equation:

$$\Delta s_{t+1} = \alpha + \beta fp_t + \upsilon_t \tag{1}$$

where Δs_{t+1} denote the one-period ahead change in the exchange rate, defined as the number of units of foreign currency per US dollar and fp_t is the forward premium, defined as the difference between the one-period foreign and US interest rates. If market agents are risk neutral and their forecast errors are uncorrelated with the forward premium, then equilibrium in the foreign exchange market implies that $\alpha = 0$ and $\beta = 1$. Since the objective of studies based on (1) is to examine the unbiasedness of forward-premium predictions and since the estimate of the slope coefficient may be influenced by the presence of an intercept term in the regression, we estimate (1) with and without α. To save space we report our results in the following manner: if the estimate of the slope coefficient based on (1) is significantly different than one, then we report estimates of both α and β; if the estimate of the slope based on (1) is insignificantly different from one, but an estimate of the slope based on the regression without the intercept is significantly different from one, we report this estimate of β; otherwise, we report estimates based on the BF regression both with and without the intercept.

Table 1 reports our estimates of the BF regression, which are based on data from DRI on U.S. dollar spot and one-month forward rates with respect to the British pound (BP), German mark (DM) and Japanese yen (JY). Our sample runs from May 1973 (1973:5) through December 1996 (1996:12).[11] The results in Table 1 are consistent with those found in the literature.[12] When the entire sample is used, the estimate of the slope coefficient ($\hat{\beta}$) is negative for both the BP and DM and significantly less than one. As for the JY, $\hat{\beta}$ is positive and insignificantly different from one or zero over the full sample. But if the regression begins in 1975:01, which is common (see Fama (1984) and others), then $\hat{\beta}$ becomes significantly negative.

Although the literature has noted that the departure of β from the null of one can be attributed to a correlation of the forward premium with either the market (aggregate) forecast error and/or the risk premium, the main sources of the temporal instability of β have been largely ignored. There are two main channels through which temporal instability can arise in the BF regression: either the aggregate expectation function is temporally unstable or the correlation between the forward premium and the market forecast error is temporally unstable. The instability of the aggregate expectation function, in turn, implies that the correlation between the forward premium and the risk premium will also be temporally unstable.[13]

To present this more formally, define the risk premium, pr, and the market forecast error, η, at time t as follows:

$$pr_t = \Delta s^e_{t+1} - fp_t \tag{2}$$

and

[11] Note that the ending observation in 1996:12 includes the one-period change in the exchange rate from 1996:12 to 1997:01.

[12] For extensive surveys of the literature see Lewis (1995) and Engel (1996).

[13] Temporal instability of β can also arise from non-stationarity of the fp process. Mark, Wu and Hai (1993) and Baillie and Bollerslev (1994) find that fp is very persistent, but probably stationary.

$$\eta_t = \Delta s_{t+1} - \Delta s^e_{t+1} \tag{3}$$

where Δs^e_{t+1} denotes the one-period ahead aggregate (market) forecast of the change in the exchange rate. Using (2) and (3), we can write the ordinary least squares (OLS) estimate of β in (1) as follows:

$$\hat{\beta}_{t,T_j} = \frac{\widehat{Cov}_{t,T_j}(\Delta s, fp)}{\widehat{Var}_{t,T_j}(fp)} = 1 + \frac{\widehat{Cov}_{t,T_j}(pr, fp)}{\widehat{Var}_{t,T_j}(fp)} + \frac{\widehat{Cov}_{t,T_j}(\eta, fp)}{\widehat{Var}_t(fp)} \tag{4}$$

where $\widehat{Cov}_{t,T_j}(.,.)$ and $\widehat{Var}_{t,T_j}(.,.)$ denote sample covariances and variances, respectively, computed for a specific sample of size T_j spanning the period of time from t through $t + T_j - 1$.[14] It is typical in the literature to assume that the OLS estimator, $\hat{\beta}_{t,T_j}$,is consistent, so that the sample estimates in equation (4) can be replaced with time-invariant probability limits. Thus we define:

$$\beta = p\lim\,(\hat{\beta}_{t,T_j}) \tag{5}$$

As we discuss in Frydman and Goldberg (2002), with imperfect knowledge, rational agents recognize that their models and rules are not strictly correct (i.e., that they "have not solved their scientific problems") and consequently behave like scientists, namely, they test their models and rules against available data and in the face of contradictory evidence revise their trading strategies. This notion of rationality applies no matter what objective functions market players possess, whether these are standard utility functions or those based on behavioral considerations: whatever the objective function, rational agents will be testing and updating their expectations functions because they recognize they do not possess the true model. To the extent that agents' forecast functions include fp or variables correlated with fp, this updating implies that the $plim$ $(\widehat{Cov}_{t,T_j}(\Delta s^e, fp))$ is temporally unstable.[15]

[14] If the BF regression does not include the constant term, the covariance and variance terms in (4) are replaced by appropriate moments.

[15] In Goldberg and Frydman (1996a) we present evidence based on survey data that the relationship between exchange rate expectations and macroeconomic fundamentals is indeed temporally unstable in the 1980's and 1990's. Unsurprsingly, we find in Goldberg and Frydman (1996b,2001) that empirical exchange rate models are also temporally unstable over the modern period of floating.

This in turn implies that the $p\lim(\widehat{Cov}_{t,T_j}(pr, fp))$ is also temporally unstable. To see this, note from (2) that

$$\frac{\widehat{Cov}_{t,T_j}(pr, fp)}{\widehat{Var}_{t,T_j}(fp)} = \frac{\widehat{Cov}_{t,T_j}(\Delta s^e, fp)}{\widehat{Var}_t(fp)} - 1 \tag{6}$$

Furthermore, with IKE, relatively stable correlations between aggregate forecast errors and publicly available information (e.g., fp) may appear from time to time. But rational agents will attempt to exploit such correlations and, in the process, update and revise their trading strategies. Individual rationality implies, therefore, that even if forecast errors were to to be correlated with the forward premium during select subperiods, such correlations will arise sporadically, experience temporal instability and disappear at *unpredictable* moments of time. Thus, individual rationality and imperfect knowledge imply that the *plim* $(\widehat{Cov}_{t,T_j}(\eta, fp))$ may also be temporally unstable, further exacerbating the instability of β in (4).

Beyond clarifying the sources of the temporal instability of β, (4) also makes clear (as noted by Froot and Frankel (1989) and others) the two sources of departure of β from 1: nonzero correlations between the forward premium and the risk premium and/or forecast errors. We close this section with two observations. First, beyond the difficulties with modeling *pr* under the REH that have been pointed out in the literature (see footnote 2), the finding of temporal instability of β in the BF regression implies that any REH-based model of *pr* would have to be consistent with temporal instability of the correlation between the risk premium and the forward premium. Second, we note that under IKE, models of the exchange rate imply no permanently stable relationship between exchange rate expectations and the forward premium (see Frydman and Goldberg (2002) and section 5.2 below). Moreover, with IKE, the correlation between exchange rate expectations (and therefore the risk premium) and the forward premium can switch between positive and negative values from one subperiod to the next. In the next section we show that instability of such a striking form is indeed an evident feature of the modern period of floating rates. We explore in section 4 the implication of this finding for the modeling of structural change under the REH, while in section 5 we explore the potential of our IKE-based model of the uncertainty premium to explain the observed switches in the value and sign of the slope coefficient in the BF regression. We leave for future research an examination of the role of transient correlations between η and fp.

3 Temporal Instability and a Refinement of the Excess-Returns Puzzle

There are a number of studies that report subsample estimates of the BF regression and find that the negative β appears to be the result of behavior during the 1980's (see footnote 3). Bekaert and Hodrick (1993) formally tests for parameter constancy using a forecast Chow test and find a break at the beginning of 1980 at the one-percent significance level. In this section we find that the temporal instability of the BF regression is much more frequent and striking than suggested by this study. Our structural change findings not only lend support to our IKE framework, but they lead to a refinement of the excess-returns puzzle.

We begin by looking at subsample results using a rough division of the sample into three subperiods of floating, the 1970's, 1980's and 1990's. These results are reported in Table 2. Table 2 shows that the results of the BF regression are sensitive to the sample period chosen. For all three currencies, $\hat{\beta}$ is positive and insignificantly different from one and zero for the 1970's and, with the exception of the JY, for the 1990's as well. During the 1980's, however, $\hat{\beta}$ is negative and significantly so for all three currencies. These results confirm those of Bekaert and Hodrick (1993) for the 1970's and 1980's and, moreover, show that the negative β obtained when using the entire sample also appears to have little to do with the behavior in the 1990's. They are also suggestive of structural change.

To test for structural change, we use the forecast χ^2 test due to Hendry (1979), which is analogous to the structural change test used in Bekaert and Hodrick (1993). These results are reported in the last column of Table 2. The results show that the null of parameter constancy can be rejected at high significance levels when the 1970's are used for the estimation period and the 1980's used for the forecast period for all three currencies (the forecast period is shown in parentheses in column 2 of the table). The results also show that while the null of no break at 1990:01 can be rejected at a high significance level for the DM, this is not the case for the BP and JY. However, as the table reports, when the first and second half of the 1980's are examined, the null of parameter constancy can be rejected for all three currencies. Finally, using the second half of the 1980's as the estimation period, the null of no break at 1990:01 is rejected at less than the five percent level for all three currencies.

Table 2 is based on crude divisions of the sample. As noted in the preceding section, with IKE we would expect the temporal instability of the BF regression to be more widespread than suggested by Table 2. To test this conjecture we use the recursive, one-step ahead Chow test provided in Hendry and Doornik (1996). The test statistic is calculated as follows: $Chow_t = \frac{(RSS_t - RSS_{t-1})(t-k-1)}{RSS_t/(t-k)}$, where RSS_t denotes the residual sum of squares for the model fitted up to and including period t, and k denotes the number of variables ($k = 1$ in (1)). Under the null of no structural change between periods $t-1$ and t, this statistic has an F distribution with 1 and $t-k$ degrees of freedom. Since the statistic and the critical values are functions of t, it is useful to divide the $Chow_t$ value by its $x\%$ critical value from the F table to yield a scaled recursive Chow test for each recursion., i.e., $\frac{Chow_t}{F_x(1,t-k)}$, which under the null should be less than one. Thus, values of this statistic greater than one imply that the null of no structural change between $t-1$ and t can be rejected at the $x\%$ level of significance.

Figures 1 through 3 report the results of this test using the five percent significance level. The figures show numerous points of structural change for all the three currencies throughout the 1970's, 1980's and 1990's. These results indicate that the temporal instability of the Fama regression is pervasive over the modern period of floating rates.[16]

As we mentioned in the introduction and outline in section 5.1 below, our IKE-based model of the uncertainty premium also implies that there will be a significant bias, either positive or negative, during subperiods characterized by two features: 1) a persistent swing in the exchange rate away from or towards the long-run benchmark level (which according to our model leads to a corresponding swing in the uncertainty premium); and 2) a trending forward premium. This suggests that the places to look for significant biases in the forward premium, both positive and negative, as well as points of structural change, are during subperiods characterized by exchange rate swings. A complete empirical analysis of this conjecture would require some definition of a swing in the exchange rate and a measure of the long-run benchmark level used by market agents in assessing potential losses. Since

[16]To gauge the extent ot the temporal instability of the BF regression we also estimated equation (1) for successive three-year subperiods, beginning in 1973:05 and rolling the three-year window through the entire sample. The results of this analysis, which are available from the authors upon request, show a similar picture to that presented in figures 1 through 3, namely the temporal instability of the BF regression for all three currencies is pervasive over the modern period of floating.

the aim of this section is to provide evidence of instability involving both positive and negative biases during swings (leading to a modification of the excess returns puzzle), we do not provide such a complete analysis of the evolution of β over the entire floating-rate period. We leave this for future research. In this paper we focus on the behavior of β during the most pronounced swing away from PPP in each of the three decades of floating. In this way we sidestep the difficulties of defining a swing and rigorously justifying our measure of the long-run benchmark level, since by most accounts the subperiods containing the most pronounced swings are widely viewed to have involved large departures from long-run bench mark levels.

Consider first the DM/$ exchange rate, which is plotted in figure 4 along with its PPP value.[17] The figure shows that major swings in this rate away from PPP occurred during the 1970's (between 1975:11 and 1978:10), 1980's (between 1982:06 and 1985:02) and 1990's (between 1993:09 and 1995:04). Given our model of the uncertainty premium, these subperiods of floating should involve a positive and growing excess return on the overvalued currency. If the forward premium also happens to be trending during these subperiods then we would expect β to deviate from one. Figure 5 plots the forward premium in each of the three subperiods during the 1970's, 1980's and 1990's. The figure shows that while the forward premium trends down throughout the exchange rate swing in the 1990's, this is not the case for the large swings in the 1970's and 1980's. For the subperiods during the 1970's and 1980's, a downward trend in fp arises only after the beginning of each of the exchange rate swings (1976:12 and 1982:11, respectively), and in the 1980's it abruptly changes direction before the exchange rate swing ends (1984:06). Consequently, the second panel of table 3 provides estimates of the BF regression based on subperiods in which the exchange rate swings are associated with clear trends in the forward premium. Despite the small sample sizes, the estimates of β are large in magnitude, significantly different from one in the 1970's and 1980's, and marginally so in the 1990's. *Strikingly, while the swing away in the 1980's is characterized by a negative bias, the bias is positive during the swings away in the 1970's and 1990's.*

A similar picture emerges when the subperiods of the most pronounced exchange rate swings away are examined for the BP and JY. Figures 6 and 7

[17] We benchmark our PPP series using the April 1989 BigMac PPP exchange rates as reported in the *Economist* magazine. We then use CPI inflation rates to derive the PPP exchange rates for all other months.

plot the BP and JY rates along with their PPP levels, respectively, whereas figures 8 and 9 plot the forward premium during the subperiods involving the most pronounced exchange rate swings away from PPP. The first and third panels of table 3 provide estimates of the BF regression, again based on those subperiods during which the exchange rate swings away from PPP are associated with clear trends in the forward premium.[18] The table does not report BP estimates for the 1990's because as figure 6 reveals, there were no major swings away from PPP during this decade. The results show that the pattern of biases for the BP and JY is similar to the pattern found for the DM. For the BP, the first panel shows that estimates of β are large in magnitude, significantly different from one during both swings, and as strikingly as with the DM, the bias during the 1970's swing is positive, while the bias during the 1980's swing is negative. Finally, for the JY (the third panel) we find that the estimates of β for the 1970's and 1980's are significantly different from one, with the former being positive and the latter negative, and the estimate for the 1990's, while significantly positive is insignificantly different from one.

Although some studies (e.g. Engel and Hamilton (1990), Bekaert and Hodrick (1993), Kaminsky (1993) and Evans and Lewis (1995)) have attempted to explain a negative β as the consequence of long swings in the exchange rate, the results in table 3 for all three currencies lead to the conclusion that such long swings away can result in *both* positive (and greater than one) and negative values of the slope coefficient in the BF regression. In section 5.2 below we offer an explanation of these results using our IKE-based model of the uncertainty premium.

The last column of table 3 also presents additional evidence that the temporal instability of the BF regression is pervasive and that it is connected with the long-swings behavior of the exchange rate. The table shows that there is a structural break at the end of *every* major swing at p-values of less *than* .001. It should be noted that these results, together with those in table 2 and in figures 1 through 3, are consistent with our earlier findings that empirical exchange rate models experience numerous break points over the modern period of floating (Goldberg and Frydman, 1996b,2001).

[18] The subperiods of exchange rate swings are 1975:02-1976:10 and 1982:06-1985:02 for the BP and 1977:05-1978:10, 1981:05-1985:02 and 1990:04-1995:04 for the JY. The subperiods involving both swings in the exchange rate and forward premium are 1975:06-1976:10 and 1983:03-1984:06 for the BP and 1977:05-1978:09, 1983:01-1984:06 and 1991:04-1995:04 for the JY.

In addition to evidence of positive and negative β's, the floating-rate experience contains many episodes, particularly in the 1970's and the 1990's, during which the estimate of the slope coefficient in the BF regression is not significantly different from zero or one. Table 4 reports the results for the other subperiods in the 1970's and the 1990's not examined in table 3. The table shows that except for four cases, the estimates are not significantly different from either one or zero. Since the standard errors of the estimates are relatively large, this evidence does not allow us to distinguish between the case in which the forward premium is an unbiased predictor of the future change in exchange rate and the other three cases examined in this paper, i.e. a positive bias, a negative bias, and the case in which the forward premium does not appear to be a predictor. Although our aim here is not to explain the behavior in these subperiods, we note that most of them do not involve persistent trends in both the exchange rate and the forward premium, and thus our theory suggests that the uncertainty premium (and therefore excess returns) would not be correlated with the forward premium. In contrast, the four subperiods characterized by an estimate of the slope coefficient that is significantly different from one in table 4 (1973:05-1975:05 and 1976:10-1979-12 for the BP and 1978:11-1979:12 and 1995:05-1996:12 for the JY) do involve trends in both the exchange rate and the forward premium for substantial parts of these subperiods. We leave an examination of these subperiods for future research.

Taken together, the results of this section provide evidence that the excess-returns puzzle is more complicated than hitherto believed. What needs to be explained is not a negative bias throughout the three decades of floating, but rather why the predictions of the forward premium are sometimes negatively biased, positively biased, unbiased or possess no predictive content, depending on the subperiod examined. Before we offer our IKE-based explanation of the excess-returns puzzle as modified in this paper, we explore the potential of the REH approach to explain the structural change documented in this section.

4 Structural Change, the REH and the Modified Excess-Returns Puzzle

The foregoing structural-change findings suggest that models relating the future change in the exchange rate to the forward premium are subperiod specific, that is, the signs and values of the intercept and slope in the BF regression seem to take on many different values over the three decades of floating. In this section we explore the question of whether such findings can be reconciled with the REH.

According to the REH, agents form their expectations according to the economist's model. Having tied expectations rigidly to the structure of the economist's model, and presuming that the parameters of this structure are functions of "deep", largely invariant, parameters of "preferences" and "technology," the RE approach implies that changes in government policy are *the primary* source of instability of the parameters of macroeconometric models (Lucas (1976)). By pointing out that changes in government policy influence agents' expectations and lead, therefore, to changes in the parameters of aggregate macroeconomic models, the Lucas critique has profoundly altered our view of the effects of government policy on the macroeconomy. Although the Lucas critique focussed on the usefulness of the older Keynesian models of the neoclassical synthesis, what seems to have remained largely unnoticed is that the structural instability implied by the Lucas critique also raises some difficulties for the *REH-based* models. Under the REH, agents are aware that government policy *will change in the future* (e.g., new Fed chairmen will be appointed), implying that a full REH solution requires specification of *all* future shifts in policy rules.[19] One way to accomplish this is to assume that the impact of all policy regimes (*past and future)* on the model structure can be characterized by a finite number of *quantitative* sub-models and to

[19] This may help explain why, except for some particular and simple shifts in policy variables (Wilson 1979), temporal instability arising from shifts in future policy rules has not received much attention in the analysis of REH-based macroeconometric models. As noted by Sargent in his study of the US inflation:

> [Rational expectations econometrics] left the drift in coefficients unexplained...Yet coefficients continue to drift for macroeconometric models...Thus, the forecasting literature has taken coefficient drift increasingly seriously, but with little help from the rational expectations tradition (Sargent (1999), p.16).

suppose that a *stationary* process governs the switches of structure between these sub-models.[20]

Such a model of structural change that can be used in REH-based models was developed by Hamilton (1988, 1989) and applied to a wide range of problems, including the excess-returns puzzle in the foreign exchange market (Engel and Hamilton (1990), Bekaert and Hodrick (1993), Kaminsky (1993), and Evans and Lewis (1995)). In particular, Engel and Hamilton assume that the change in the exchange rate at time t is a realization from one of the two "regimes", specified as random walks with different values of the drift coefficient. The specific regime that the exchange rate is in at any given time depends on the realization of an unobserved random variable characterized by a stationary Markov chain. Under the REH, this model of structural change presumes that the impact of past and future policy rules can be specified as one overarching model of exchange rate movements, composed of two specific sub-models and some stationary switching process between these sub-models. This presumption is tantamount to the assumption that structural change can be modeled as a stationary process, thereby allowing the REH to be used to model agents' expectations. As in standard settings, the REH assumes that agents use the economist's model, which now provides for a specific form of structural change, that is agents form expectations on the basis of the true values of the parameters of the two sub-models and the Markov transition probabilities.[21]

Engel and Hamilton (1990) estimate the parameters of the two random-walk models and transition probabilities for the DM/$, French franc and

[20] The so-called "rational" learning, which assumes that all agents use some learning algorithm to learn about changes in policy, has also been used to model structural change resulting from changes in government policy. We show in Frydman (1982) and extensively discuss in Frydman and Goldberg (2002) that the assumption that agents adhere to such rules during learning is, in general, incompatible with the postulate of *individual* rationality.

[21] The assumption that agents' knowledge of the parameters of the economist's model is superior to the economist himself is basic to the RE approach. As Sargent noted

> Rational expectations equilibrium...typically imputes to the people more knowledge about the system they are operating in than is available to the economists using the model to try to understand their behavior (Sargent (1993, p.21)).

Moreover, we note that the REH is even more problematic here than in the simpler settings, since agents are assumed to know the transition probabilities of the *unobseved* process governing transitions between models and, as we shall discuss next, the estimated number of transitions is very small.

BP exchange rates. The resulting estimates imply very few expected transitions.[22] One might suspect that the small number of transitions might be an artifact of the simplicity of the two random-walk setup. However, estimates of transition probabilities obtained in studies with more general two-model structures also imply very few transitions.[23] In contrast to these studies, the results in the preceding section suggest that the structural-change process in the foreign exchange market entails many transitions (structural breaks) between many different models. This can seen in table 3, where most of the estimates of the slope coefficient are significantly different from one (and zero) and some of these estimates are positive, while others are negative.

One way to reconcile the apparent inconsistency between our structural-change findings and those generated by the Engel and Hamilton model is to argue that, despite the apparent variation among the estimates in table 3, our results are due to small sample biases of the estimates of the slope parameter and/or its standard error and that, in fact, the underlying set of different models for the exchange rate is quite small, say two as in the Engel and Hamilton model. Given the substantial magnitudes of the estimates of the slope parameter in table 3, this would appear to be an unlikely explanation. In any case, to explain away our results as due to small sample biases, one would have to demonstrate that the estimated pattern of positive and negative values of the slope parameter during different subperiods presented in table 3, is in fact consistent with a stationary process such as the Markov switching between two (or some feasibly small number of) *pre-specified* sub-models characterizing the exchange rate change in each "regime".

But even if the two-state Markov switching model were found to be inconsistent with the observed structural change, the question would still remain as to whether a model of switching between n (feasibly small, but greater than two) number of models might characterize reasonably well the temporal instability of the exchange rate process. Such n-state Markov switching models might result in estimates that imply a greater number of transitions than in the Engel and Hamilton model. Since the literature does not contain examples of an n-state Markov switching model, it is not known how well

[22] This is implied by the large estimated probabilities of remaining in each regime from one period to the next, i.e. large diagonal elements of the transition matrix of the two-state Markov switching model.

[23] Bekaert and Hodrick (1993) specify the two models of the exchange rate (and the forward premium) as more general ARIMA processes and Evans and Lewis (1995) extend the Engel and Hamilton switching model to allow for jumps in the exchange rate.

such extensions will work. However, while the extended model might fit better the structural change during some *past* period spanned by the sample, the economist would need to specify sub-models ("regimes") capturing the impact of *all future* policy rules on the behavior of the exchange rate.

There is also another difficulty with the two-state (or n-state) Markov switching model. This model assumes that after each transition to a "new" policy regime, government policy either remains at the new regime or reverts back to the "old" regime. It would seem natural to suppose that because each future administration (or each successive year of the same administration) is at least somewhat different than those in the past, implying that as time proceeds, government policy will "transit" to rules different (in terms of variables and/or values of parameters) than those characterizing policy in the past. Thus, even if structural change is caused primarily by changes in government policy, we would expect to find more than two different models are needed to characterize the behavior of exchange rates over three decades of floating, and that the number of switches between models substantially exceeds the number of transitions implied by the estimates of the two-state Markov switching model.

Beyond its inherent problems as a model of structural change, the (two-state) Markov switching model has provided little help in explaining the excess returns puzzle under the REH. After estimating the Markov switching model, Engel and Hamilton (1990) conclude that "in the absence of a plausible story about foreign exchange risk premia, we conclude that there are long swings in the dollar and the *markets do not know it* (p.711, emphasis)." Bekaert and Hodrick (1993) modify the Engel and Hamilton model to allow for a richer class of ARIMA models of the exchange rate and the forward premium. To reduce the small sample bias, Bekaert and Hodrick reestimate the BF regression under the constraint that the exchange rate and the forward premium follow the two-state Markov process. Their estimate of β is still highly negative, which leads them to conclude: "After considering alternative sources of bias, our conclusion is that the evidence against the unbiasedness hypothesis using the rational expectations econometrics is very strong (Bekaert and Hodrick (1993, p. 132)." Evans and Lewis (1995) extend the Engel and Hamilton model to allow for jumps in the exchange rate and introduce a risk premium into their REH-based econometric specification of the exchange rate process. Using simulations of their switching process to generate realizations of the exchange rate process, they re-estimate the BF regression and argue that "some of the anomalous behavior [of foreign exchange returns] can be explained by rational expectations about the shifts between appreciating and depreciating processes (p. 734)."

Bekaert and Hodrick and Evans and Lewis carry out their analyses under the assumption that the transition probabilities of the Markov switching model are constant. The study by Kaminsky (1993) suggests that such probabilities are likely to be temporally unstable.[24] In addition to assuming Markov switching between two random-walk models of the exchange rate, as in Engel and Hamilton (1990), Kaminsky introduces a two-state Markov switching model governing the transition between two "informational states," dubbed "correct" or "wrong". The two informational states arise from the assumption that additional information is available to participants in the foreign exchange market, such as Fed announcements, that provides either the "correct" or "wrong" signal on the actual, but unobservable, regime. Thus, in the Kaminsky extension of the Engel and Hamilton model, the estimation of the model involves two Markov-transition matrices (rather than one). Kaminsky shows that an estimate of the probability of the Markov model governing the transition between the two random walks depends on the specification of the Markov process governing the transition between the informational states. Moreover, the transition probabilities between informational states are *temporally unstable*: the values of these probabilities are different for the period when Burns and Miller were chairman of the Fed as compared with the period of the Volker and Greenspan chairmanships.

In contrast to the REH-based approach to modeling structural change and the bias of the forward rate, the IKE approach is predicated on the presumption that the temporal instability arising in the foreign exchange market (and therefore the excess returns puzzle as modified in this paper) cannot be explained by a *pre-specified, overarching* model of structural change. We consider it unlikely that an overarching quantitative model could be found to characterize economic change, even if it were to be confined to the temporally unstable relationships in just one market (e.g., the foreign exchange market) and even if the change was *solely* a consequence of changes in government policy rules, as in the REH framework.[25] This scepticism results

[24]Kaminsky motivates her study by the observation that the Engel and Hamilton model makes an unrealistic assumption that agents use only information contained in past observations of the exchange rate in estimating the current unobservable state of the process and in forming their forecasts of the change in the exchange rate.

[25]We note that the characterization of government policy by a *fixed* rule is also likely itself to involve parameter drift and large and discontinuous structural change, as policy

from the fact that *pre-specification* of future government policy, both monetary and fiscal, seems to require quantitative modeling of the behavior of *future and unknown* policy officials. Nevertheless, these arguments do not preclude the possibility that the REH approach may eventually be able to deal with the kind of temporal instability that appears to be an inherent feature of macroeconomic phenomena, particularly in the asset markets.

We also note that in addition to temporal instability in macroeconometric models arising from changes in government policy, the IKE framework implies that temporal instability will also arise from the partly autonomous, continuous and discontinuous, updating of agents' expectations in a world of imperfect knowledge.[26] Thus in *our world of imperfect knowledge*, the potential universe of *future* models has to remain *open* and structural change procedures, such as the ones we used in the preceding section, should accompany any empirical analysis of macroeconometric models, especially in asset markets. In the next section, we show how such an open-ended approach to the analysis of the time-series of the exchange rate and the forward premium over the floating-rate period can provide an explanation of the excess returns puzzle, as modified in this paper.

5 Towards a Resolution of the Excess-Returns Puzzle

In this section we first outline our IKE-based model of the uncertainty premium.[27] This model implies that persistent swings in the exchange rate both

officials update their necessarily imperfect views of the world. For evidence that some popular "policy rules" are subject to temporal instability see for example, Frydman and Rappoport (1987) and more recently Clarida, Gali and Gertler (2000).

[26] The neglect of imperfect knowledge might be a more important problem in macroeconomics than has been recognized. As Phelps put it in his recent reflections on the state of macroeconomics

> "What got lost in monetary macroeconomics amid its considerable achievements in modeling imperfect information is the imperfect knowledge, or uncertainty in the sense of Knight, that pervades the more entrepreneurial of the market economies" (Phelps (2002))

[27] For a complete treatment of our model of the uncertainty premium see Frydman and Godberg (2001, 2002, 2003a).

away from and toward long-run benchmark levels will be associated with particular swings in expected excess returns. We then show how this implied behavior of expected excess returns helps to explain not only the temporal instability reported in tables 2 and 3 and figures 1 though 3, but also the particular pattern of positive and negative biases reported in table 3 for the major exchange rate swings of the 1970's, 1980's and 1990's.

5.1 Imperfect Knowledge Expectations, Myopic Loss Aversion and the Uncertainty Premium

Our model of the expected excess return on foreign exchange assumes imperfect knowledge expectations, heterogeneity of beliefs and myopic loss aversion. With IKE and the assumption of individual rationality, expectation functions will be temporally unstable, and as we showed in section 2 above, this will lead to temporal instability in the relationship between the uncertainty premium and the forward premium and, therefore, to temporal instability of β in the BF regression. In addition to temporal instability, we showed in Goldberg and Frydman (1996a) and Frydman and Goldberg (2001,2002) that with IKE, otherwise standard models of the exchange rate can generate long-swings in the exchange rate. This long swings result leads us to develop a dynamic version of myopic loss aversion that accounts for the evolution of agents' assessments of losses during episodes of persistent movements of the exchange rate from benchmark levels.

The assumption of myopic loss aversion due to Kahneman and Tversky (1979) and Benartzi and Thaler (1995) maintains that agents are more sensitive to reductions than to increases in wealth, the level of wealth is of second order of importance and agents monitor returns on their investments relatively frequently. If we had assumed risk neutrality, agents' decisions to buy or sell foreign exchange would depend only on their assessments of expected returns. But with myopic loss aversion, agents' decisions to buy or sell foreign exchange depend on their assessments of the *expected utility* of these decisions. Moreover, since agents are more sensitive to losses than to gains, they need to be compensated for this added sensitivity. Thus, *both* buyers and sellers of foreign exchange demand a premium (over the risk-neutral return) on taking open positions in the foreign exchange market.

Consider first the group of buyers of foreign exchange and let $E^{\mathrm{B}}(R)$ denote this group's expected one-period return on buying foreign exchange (i.e., $E^{\mathrm{B}}(R) = E^{\mathrm{B}}(\Delta s_{t+1}) - fp_t$), where aggregation is accomplished using wealth shares. If the foreign exchange market was populated by buyers only, then the exchange rate would be bid to the point where the *prospective* or expected utility of this return equaled zero. But because the utility functions of buyers place greater weight on potential losses, a prospective return of zero necessarily implies a positive excess return, $E^{\mathrm{B}}(R) > 0$, i.e., buyers as a group are unwilling to bid the exchange rate up to the risk-neutral point where the expected yields on domestic- and foreign-currency positions are equal. As such, with myopic loss aversion, uncovered interest rate parity (UIP) does not hold and buyers of foreign exchange expect to earn a premium on their open positions. Using the utility function of Benartzi and Thaler [1995], which has become standard in the literature, we show in Frydman and Goldberg (2001,2002, 2003a) that this expected (excess) return is a function of the "expected loss" of buyers, L^{B}, and their degree of loss aversion, which is captured by the parameter $\lambda > 1$, i.e., $E^{\mathrm{B}}(R) = (1 - \lambda)\, L^{\mathrm{B}}$.[28] It is important to note that for buyers, losses from foreign currency speculation represent negative realizations of R, so that $L^{\mathrm{B}} < 0$ and $E^{\mathrm{B}}(R) > 0$. We will refer to this expected excess return as the uncertainty premium of buyers.

Analogously, let $E^{\mathrm{S}}(R)$ and L^{S} denote the expected return and expected loss of the group of sellers, respectively, giving $E^{\mathrm{S}}(R) = (1 - \lambda)\, L^{\mathrm{S}} < 0$.[29] Note that for sellers of foreign exchange, losses represent positive realizations of R, so that $L^{\mathrm{S}} > 0$ and with $\lambda > 1$, $E^{\mathrm{S}}(R) < 0$. Again, if the market contained only sellers, this group would be unwilling to bid the exchange rate down to the point where UIP would hold. They therefore expect to earn a premium on selling foreign exchange. We will refer to this expected excess return as the uncertainty premium of sellers.

In the aggregate, the expected excess return on foreign exchange, which we call the aggregate uncertainty premium and denote by U, is the summation of the uncertainty premia of the buyers and sellers, i.e., $U = E^{\mathrm{B}}(R) + E^{\mathrm{S}}(R) = (1 - \lambda)\,(L^{\mathrm{B}} + L^{\mathrm{S}})$. Thus, positive values of U represent the premium required by loss-averse buyers of foreign currency in *excess* of the premium

[28] The expected loss is defined as the average of potential losses based only on the truncated loss part of the subjective distribution governing returns, so that $E^{\mathrm{B}}(R) = G^{\mathrm{B}} + L^{\mathrm{B}}$, where G^{B} denotes the expected gain. The literature has reported λ to be in excess of 2. See Benartzi and Thaler (1995) and references therein.

[29] We also note that an agent will sometimes belong to the group of buyers and at other times to the group of sellers. It seems reasonable, therefore, to assume that the average λ for both groups is the same.

required by loss-averse sellers. One of the immediate implications of our model of U is that if the weight of buyers of foreign exchange (sellers) is consistently greater than that of sellers (buyers) during any subperiod, such that $E^{\mathrm{B}}(R)$ is consistently greater than $E^{\mathrm{S}}(R)$, we would expect such subperiods to be characterized by rising (falling) values of foreign exchange (falling) and a generally positive (negative) aggregate uncertainty premium. Hence, transitions from subperiods characterized by a generally rising (falling) exchange rate to ones characterized by a generally falling (rising) exchange rate will be associated with sign reversals in expected excess returns.[30]

In Frydman and Goldberg (2001,2002,2003a) we develop a dynamic model for equilibrium movements in the aggregate uncertainty premium, U. This entails specifying how L^{B} and L^{S} are connected to the evolution of exchange rate expectations and the exchange rate during exchange rate swings. If agents are aware of the tendency of floating exchange rates to exhibit long swings, then they will view the gap in the exchange rate from long-run benchmark levels as an important factor in assessing the potential losses from currency speculation. We assume, therefore, that an increase in the gap between the exchange rate and the perceived long-run benchmark level leads buyers (sellers) to revise upwards (downwards) the absolute value of their assessments of the potential losses. To see the intuition behind this assumption assume the group of buyers and sellers believe that the value of foreign exchange is greater than the perceived benchmark level and that the value of foreign exchange rises further. Since both buyers and sellers are aware of the long-swings nature of the exchange rate, buyers become less confident of a further movement away and sellers become more confident of a countermovement. As such, buyers (sellers) revise upwards (downwards) their assessments of the potential losses, i.e., $-L^{\mathrm{B}}$ rises and L^{S} falls, implying a rise in U.

We show in Frydman and Goldberg (2001,2002,2003a) that this reasoning leads to the following implication: the absolute value of U will rise (fall) persistently during any subperiod characterized by a persistent swing in the exchange rate away from (back to) the long-run benchmark level. In Frydman and Goldberg [2001] we use survey data from MMS and provide preliminary evidence that expected excess returns are indeed related to the gap from

[30]For the inability of standard models of the risk premium under the REH to match up with the observed sign reversals in excess returns see Lewis (1995) and Mark and Wu (1998).

benchmark levels, as measured by PPP.[31]

5.2 Explaining the Modified Excess Returns Puzzle

We are now ready to use our IKE model of the uncertainty premium to shed light on the temporal instability of the BF regression and the pattern of positive and negative biases reported in table 3 for the major swings away from PPP during the 1970's, 1980's and 1990's. The swings away during the 1980's, which involved persistently rising values of the U.S. dollar away from PPP, implies that the weight of buyers of dollars (taken here to be the foreign currency) consistently dominated during this subperiod. According to our model, then, the aggregate excess returns on the dollar relative to the BP, DM and JY should have been consistently positive during these swings, as buyers required an uncertainty premium (in excess of the premium required by sellers) to compensate them for the potential losses and uncertainty associated with a countermovement back to PPP. Moreover, our model implies that these positive excess returns on the dollar should have risen persistently during these swings as the gaps from PPP widened and as buyers (sellers) of dollars increased (decreased) their assessments of the potential losses. Similar predictions hold for the swings away during the 1970's and 1990's, although during these swings, sellers of dollars obviously dominated, implying that the excess returns on dollars during these subperiods should have been negative and trending down.

Figures 10 through 12 plot the *ex post* excess return (R) on the U.S. dollar along with trend lines during all of the persistent swings away from PPP examined in section 3. As predicted by our model, R is generally negative and trending down during the exchange rate swings in the 1970's and 1990's for all currencies and generally positive and trending up during the exchange rate swings in the 1980's for all currencies.

The extent to which the trending of excess returns in figures 10-12 implies biases in forward-premium predictions depends on the behavior of the forward premium during these swings. For each of these swings, if fp trends persistently in one direction during some portion or all of the swing, then this will lead to a non-zero correlation between the R and fp, generating a

[31] Canova (1991), Bekaert (1994) and Gokey (1994) all report evidence suggesting that deviations from uncovered interest rate parity are related to deviations from PPP. Engel (1996) argues, however, that the decompositions used in these papers preclude insight into this problem.

bias in forward-premium predictions during that portion of the exchange rate swing involving a trending fp. Moreover, given that our model generates a clear prediction concerning the sign and direction of change of R during each of the swings away, our model also generates clear predictions concerning the signs of the biases reported in table 3.

Consider first the DM swing away from PPP during the period from 1982:06 to1985:02. Our IKE-based model of the uncertainty premium predicts a negative correlation between U and fp and therefore a negative bias during 1982:11-1984:06, since it predicts that the excess return on the dollar should be trending up during the exchange rate swing and because fp happens to trend down during 1982:11-1984:06 as shown in figure 5. The same logic implies negative biases of forward-premium predictions for the BP and JY during the 1980's swings, since during these swings, the exchange-rate and forward-premium trends were in the same directions as for the DM. As reported in table 3, these predictions of negative biases are borne out for all three currencies, with estimates of the slope coefficient that are all significantly less than one and zero despite the small sample sizes. The fact that all of the slope estimates are negative is not surprising because the contribution of exchange rate changes to *ex post* excess returns consistently dwarfs the contribution of the forward premium over the period of floating rates, implying that a trend in *ex post* excess returns implies a similar trend in the change of the exchange rate.

In similar fashion, our theory predicts positive correlations between U and fp and therefore positive biases during substantial portions of the swings in the 1970's and 1990's for all currencies. This is because it predicts that the excess return on the dollar should be trending down during these swings and because fp also happens to trend down during substantial portions of these time periods. As reported in table 3, these predictions of positive biases are borne out for all currencies during the swings in both the 1970's and 1990's, with estimates of the slope coefficients which are all greater than one and despite the small sample sizes, significantly so for four out of the five cases.

In addition to providing an explanation of the pattern of biases reported in table 3, our IKE-based model of the uncertainty premium also provides an explanation of the other key feature of the modified excess-returns puzzle: the widespread temporal instability of the BF regression. Section 2 showed that temporal instability in the relationship between exchange rate expectations and fp would lead to instability in the correlation between U and fp and thus, in the slope coefficient in the BF regression. The assumption of individual rationality in a world of imperfect knowledge implies that expectation functions will be temporally unstable, suggesting that this channel will be a major source of instability in the BF regression.

Most exchange rate theory posits a causal connection between relative interest rates and the exchange rate. For example, the Dornbusch (1976) and Frankel (1979) (DF) sticky-price monetary model implies that following a one-time overshooting of PPP, the exchange rate movement back will involve particular trends in both real and nominal exchange rates and interest rates. Frankel [1985] and Dornbusch and Frankel [1988], among others, use this logic to argue that the rise in the value of the U.S. dollar during the first half of the 1980's was due in large measure to the rise in U.S. long-term real interest rates relative to the rest of the world that occurred during this period. But as noted by these authors, it is difficult to reconcile the DF model with persistent movements of the exchange rate away from PPP. Referring to the swings of the 1980's, Dornbusch and Frankel remark that, "the dollar overshot the overshooting equilibrium (Dornbusch and Frankel (1998), p. 17)." Without a theory of long swings it is difficult to talk about exchange rate movements.

We show in Goldberg and Frydman (1996a) and Frydman and Goldberg (2002) that with IKE, the DF model not only generates long swings, but that a persistent rise in the exchange rate away from PPP may be associated with a rising or falling nominal interest rate differential, depending in part on the weights agents attach to the set of fundamental variables in forming expectations.[32] When rational agents coping with imperfect knowledge inevitably update their expectations functions, any relationship between nominal exchange rates and nominal interest rates that may have existed over some subperiod may either disappear or even switch sign.

As suggested by the time plots of excess returns in figures 10 through 12 and of the forward premium in figures 5, 8 and 9, such instability seems to be an integral part of the behavior in the foreign exchange market. Consider the exchange rate swings of the 1980's. For all three currencies, the swings away from PPP began by mid 1982 (1982:06 for the BP, 1982:06 for the DM and 1981:02 for the JY) and continued until the beginning of 1985. Inspection of figures 5, 8 and 9 reveals that for all three currencies the downward trends

[32]The DF model with IKE does imply a stable long-run relationship between swings in the real exchange rate and swings in the real interest rate differential. This implication of IKE is used in Frydman, Goldberg and Juselius (2003) to explain the PPP puzzle.

in fp set in only after the exchange rate swings began and reversed direction in mid 1984, well before the exchange rate swings also reversed direction. Given this behavior on the part of fp and given that our IKE-based model of the uncertainty premium implies an upward trending U throughout the exchange rate swings away, our model implies that the correlation between U and fp should be unstable over the full subperiods involving the swings. This logic leads to two predictions. First, the BF regression should experience a structural break in 1984:07 for all three currencies As we reported in table 3, this is indeed the case at p-values lower than .001 for *all three* currencies. Second, the biases reported in table 3 should all fall and perhaps lose their significance if the BF regression were run over the full subperiods involving the swings. Table 5 reports the results of such regressions. In each case, the slope estimates become much less negative (for the BP, the slope coefficient becomes positive and for the DM and JY the slope increases from -19.8 to -6.61 and from -22.29 to -4.31, respectively!) and all estimates become insignificantly different from one or zero. In contrast to earlier studies suggesting that the 1980's are characterized by negative biases, the null hypothesis of unbiasedness cannot be rejected over these subperiods in the 1980's.

Inspection of the time plots of excess returns and the forward premium for the swings in the 1970's and 1990's, as well as the evidence of temporal instability reported in table 3, suggests that similar logic and behavior is at work during these other subperiods of exchange rate swings. Hence, taken as a whole, our empirical findings lead to the following conclusion: the temporal instability of expectation functions, through its impact on the correlation between the uncertainty premium and the forward premium, provides a reasonable explanation of the observed temporal instability of the BF regression.

6 Concluding Remarks

This paper has used a new model of the risk premium under imperfect knowledge expectations, called an uncertainty premium, to explain the modified excess-returns puzzle, i.e., the observation that the predictions of the forward premium are sometimes negatively biased, positively biased, unbiased or possess no predictive content, depending on the subperiod examined. It is important to re-emphasize that the observed pattern of positive and negative values of the slope coefficient may also be due to the transient correlation between the forecast errors and the forward premium. Hence, one of the issues that remains to be examined is the question of the relative contributions of the uncertainty premium and forecast errors in explaining the excess returns puzzle, as modified in this paper.

The inability of REH-based models of the risk premium and structural change to explain the findings of the BF regression has led to a troublesome suggestion that economic agents are grossly irrational. Our IKE-based model of structural change and the uncertainty premium holds out the promise that the foreign exchange market may be informationally efficient after all. It seems to us, therefore, that macroeconomists face a choice: either they can continue trying to explain asset market anomalies using the REH and live with the specter of gross irrationality in asset markets or they can recognize that imperfect knowledge, which is an inherent feature of the environment in which agents have to form expectations, may be the key to reconciling the movement of asset prices with the postulate of individual rationality.

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Table 1[a]
The BF Regression

		$\hat{\alpha}$	$\hat{\beta}$	R^2
BP	1973:5-1996:12	-.540 (.24)	**-1.54***** (.68)	.016
	1975:1-1996:12	-.497 (.25)	**-1.51***** (.77)	.014
DM	1973:5-1996:12	-.256 (.23)	**-.52*** (.85)	.002
	1975:1-1996:12	.238 (.24)	**-.62*** (.88)	.003
Yen	1973:5-1996:12	-.271 (.20)	*.035* (1.16)	.000
		-	*.42* (1.05)	
	1975:1-1996:12	-.687 (.21)	**-3.97***** (1.60)	.025

[a]Numbers in parentheses are standard errors corrected for heteroskedasticity using White (1980). Bold numbers for $\hat{\beta}$ indicate significance from one, with three stars denoting significance at the one percent level, two stars at the five percent level, one star at the ten percent level and no stars at the fifteen percent level. Numbers in italics denote insignificance from one.

Table 2[a]
The Structural Instability of the BF Regression

		$\hat{\alpha}$	$\hat{\beta}$	R^2	Structural Change Test[b]
BP	1973:5-1979:12 (80:1-89:12)[c]	-.170 (.539)	*-.04* (1.12)	.000	190.63 (.00)
		-	*.25* (.65)	-	
	1980:1-1989:12 (90:01-96:12)	-.908 (.323)	**-4.49***** (1.114)	.106	92.721 (.24)
	1980:1-1984:12 (85:01-89:12)	-1.135 (.369)	**-2.60***** (1.16)	.056	113.7 (.00)
	1985:1-1989:12 (90:01-94:12)	-2.067 (1.363)	**-9.01**** (4.34)	.076	80.34 (.04)
	1990:1-1996:12	.28 (.416)	*1.20* (1.86)	.006	
		-	*.59* (1.36)	-	
DM	1973:5-1979:12 (80:01-89:12)	-.340 (.577)	*.85* (2.02)	.003	174.33 (.00)
		-	*1.67* (1.23)	-	
	1980:1-1989:12 (90:01-96:12)	-1.794 (.886)	**-5.57**** (2.81)	.062	(122.32 (.00)
	1980:1-1984:12 (85:01-89:12)	.232 (1.234)	*-1.913* (3.43)	.012	99.61 (.00)
	1985:1-1989:12 (90:01-96:12)	-3.516 (1.220)	**-10.43**** (4.96)	.051	240.77 (.00)
	1990:1-1996:12	-.057 (.355)	*.16* (1.52)	.000	
		-	*.06* (1.46)	-	
Yen	1973:5-1979:12 (80:01-89:12)	-.054 (.326)	*1.63* (1.09)	.033	177.6 (.00)
		-	*1.65* (1.11)	-	
	1980:1-1989:12 (90:01-96:12)	-1.305 (.496)	**-6.41**** (2.95)	.047	67.57 (.90)
	1980:1-1984:12 (85:01-89:12)	-.629 (.705)	**-3.98** (3.35)	.031	49.13 (.07)
	1984:1-1989:12 (90:01-92:12)	-2.344 (.730)	**-16.33***** (5.85)	.085	52.23 (.04)
	1990:01-1996:12	-.277 (.331)	**-7.16**** (3.77)	.043	
		-	**-6.83**** (3.80)	-	

[a]Numbers in parentheses are standard errors corrected for heteroskedasticity using White (1980). Bold numbers for $\hat{\beta}$ indicate significance from one, with three stars denoting significance at the one percent level, two stars at the five percent level, one star at the ten percent level and no stars at the fifteen percent level. Numbers in italics denote insignificance from one.
[b]Forecast chi square test due to Hendry (1980). Numbers in parentheses are p-values.
[c] The parentheses shows the forecast period.

Table 3[a]
The BF Regression
Major Swings Away from PPP

		$\hat{\alpha}$	$\hat{\beta}$	R^2	Structural Change Test[b]
BP	1975:06-1976:10 (79:12)	-	**3.59**** (1.09)	-	85.90 (.000)
	1983:03-1984:06 (79:12)	-.10 (.96)	**-13.02**** (6.65)	.191	183.12 (.000)
DM	1976:12-1978:10 (79:12)	-	**5.13**** (2.01)	.128	49.3 (.000)
	1982:11-1984:06 (89:12)	-6.42 (2.64)	**-19.80***** (6.37)	.239	123.04 (.000)
	1993:09-1995:04 (96:12)	-1.45 (.78)	**6.91** (3.91)	.168	65.42 (.000)
JY	1977:05-1978:10 (79:12)	-	***12.46***** (4.77)	-	61.56 (.000)
	1983:01-1984:06 (89:12)	-2.53 .95	**-22.29***** (6.06)	.226	258.70 (.000)
	1991:04-1995:04 (96:12)	-1.20 (.51)	*7.25* (7.15)	.035	61.46 (.000)

[a]Numbers in parentheses are standard errors corrected for heteroskedasticity using White (1980). Bold numbers for $\hat{\beta}$ indicate significance from one, with three stars denoting significance at the one percent level, two stars at the five percent level, one star at the ten percent level and no stars at the fifteen percent level. Numbers in italics denote insignificance from one.
[b]Forecast chi square test due to Hendry (1980). Numbers in parentheses are p-values.

Table 4[a]
BF Regression
Other Subsamples: 1970's and 1990's

		$\hat{\alpha}$	$\hat{\beta}$	R^2
BP	1973:05-1975:05	-1.78 (.61)	**-2.68***** (1.08)	.123
	1976:10-1979:12	-	**-2.19***** (.59)	-
	1990:1-1996:12	.28 (.416)	*1.20* (1.86)	.006
		-	*.59* (1.36)	-
DM	1973:05-1976:11	-.16 (.72)	*.99* (3.37)	.003
		-	*1.44* (2.38)	-
	1978:11-1979:12	-5.17 (4.56)	*-9.93* (9.92)	.071
		-	*.09* (1.78)	-
	1990:01-1993:08	-.57 (.74)	*1.66* (2.43)	.008
		-	*.33* (1.71)	-
	1995:05-1996:12	-1.13 (5.02)	*-11.66* (29.92)	.013
		-	*-4.72* (3.47)	-
JY	1973:05-1977:04	-.06 (.28)	*2.28* (1.06)	.197
		-	*2.25* (1.01)	
	1978:11-1979:12	-.73 (.3.29)	*-9.99* (12.76)	.041
		-	**-7.54**** (3.81	
	1990:01-1991:03	-.17 (1.13)	*-5.55* (13.65)	.009
		-	*-6.70* (11.66)	-
	1995:05-1996:12	-	**-14.71***** (4.94)	-

[a]Numbers in parentheses are standard errors corrected for heteroskedasticity using White (1980). Bold numbers for $\hat{\beta}$ indicate significance from one, with three stars denoting significance at the one percent level, two stars at the five percent level, one star at the ten percent level and no stars at the fifteen percent level. Numbers in italics denote insignificance from one.

Table 5[a]
The BF Regression
The Swings of the 1980's

		$\hat{\alpha}$	$\hat{\beta}$	R^2
BP	1982:06-1985:02	-1.49 (.48)	*.08* (3.24)	.000
DM	1982:06-1985:02	-1.40 (2.22)	*-6.61* (5.60)	.055
JY	1981:05-1985:02	-.32 (.96)	*-3.42* (4.91)	.016

[a]Numbers in parentheses are standard errors corrected for heteroskedasticity using White (1980). Bold numbers for $\hat{\beta}$ indicate significance from one, with three stars denoting significance at the one percent level, two stars at the five percent level, one star at the ten percent level and no stars at the fifteen percent level. Numbers in italics denote insignificance from one.

Figure 1
One-Step Chow Test
BP

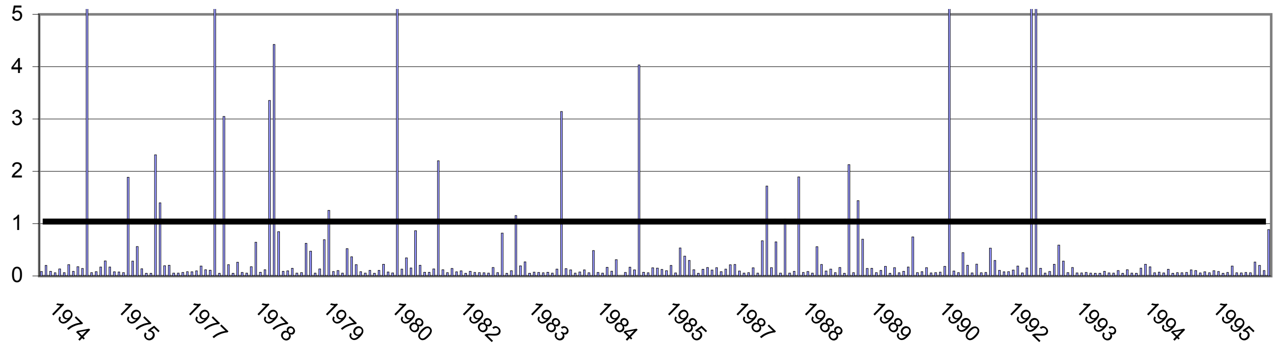


Figure 2
One-Step Chow Test
DM

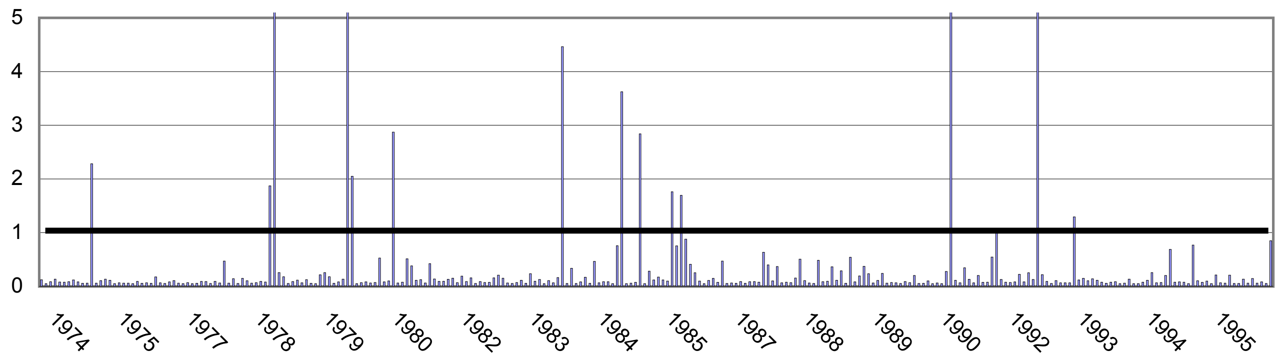


Figure 3
One-Step Chow Test
JY

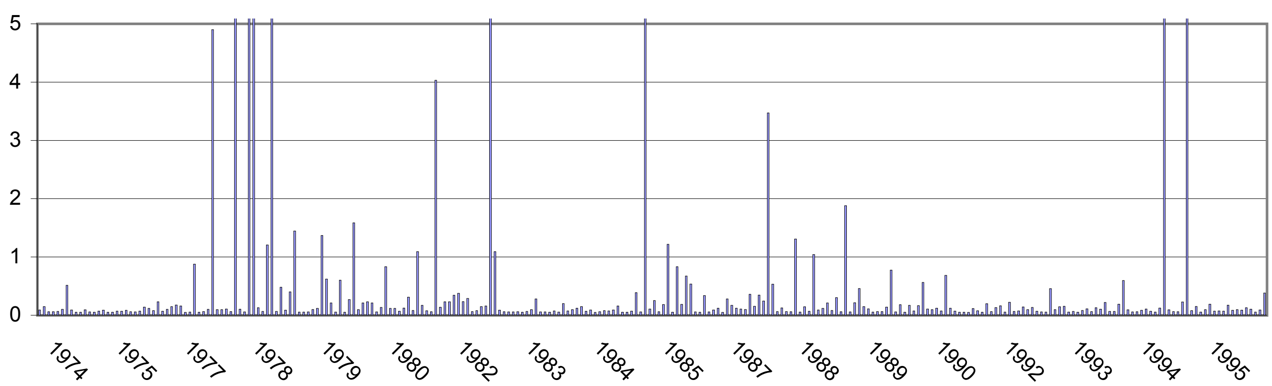


Figure 4

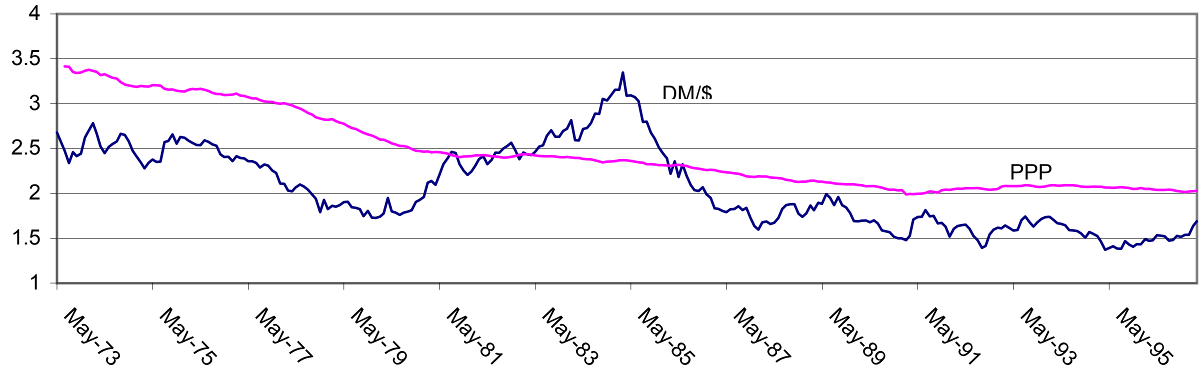


Figure 5
DM Forward Premium
(a)

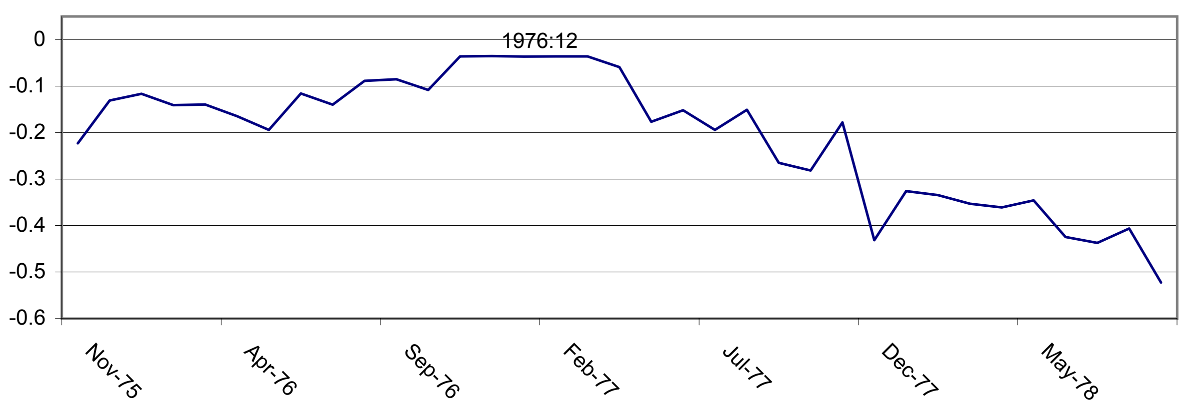


Figure 5
(b)

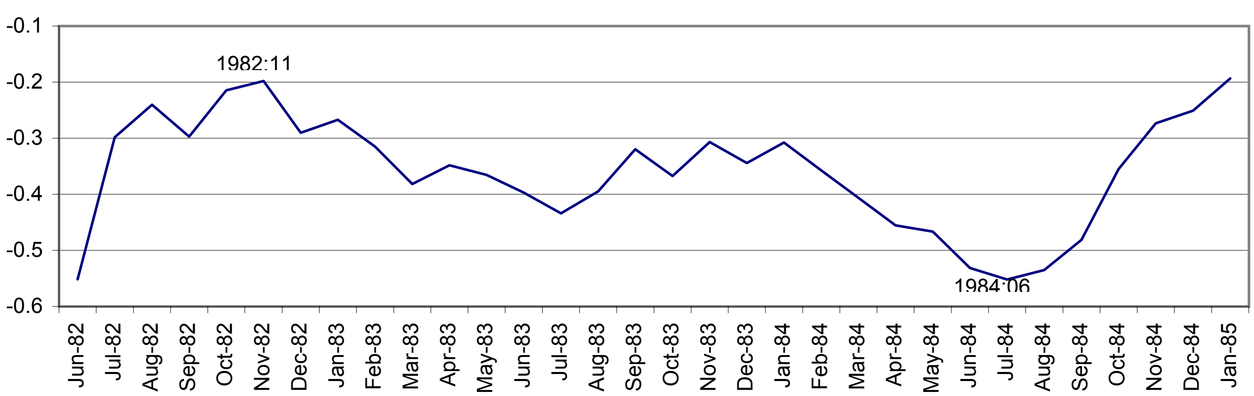


Figure 5
(c)

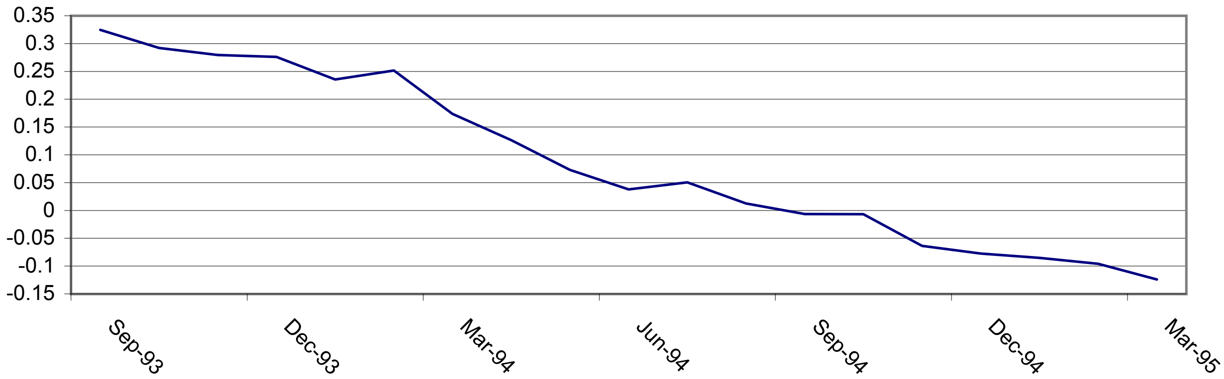


Figure 6

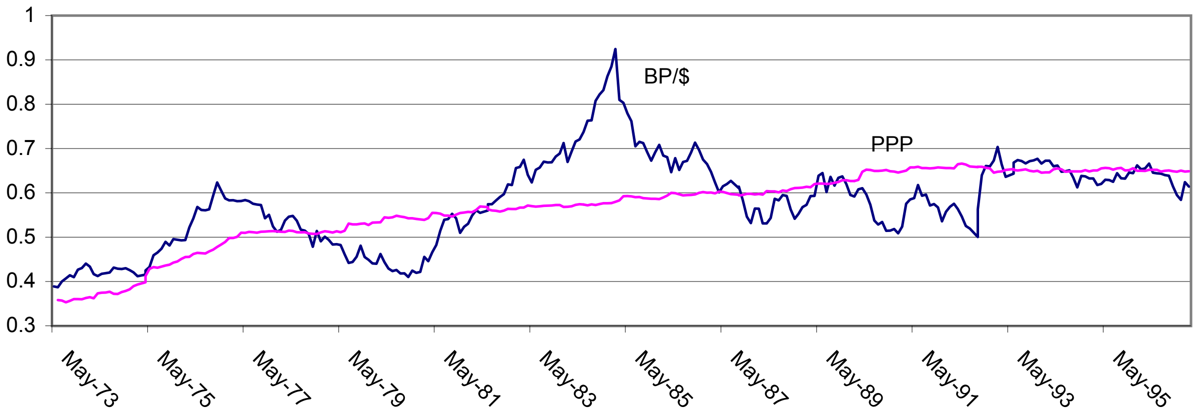
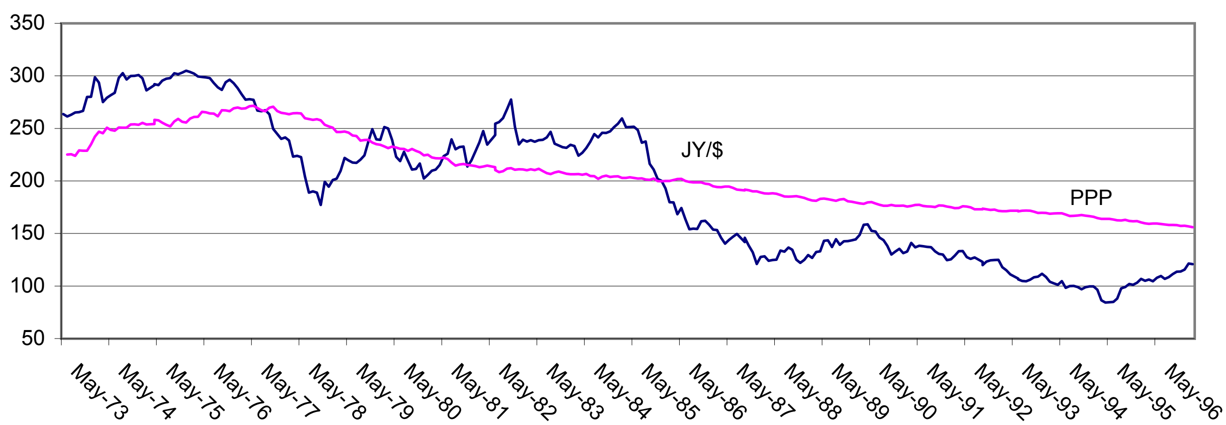


Figure 7



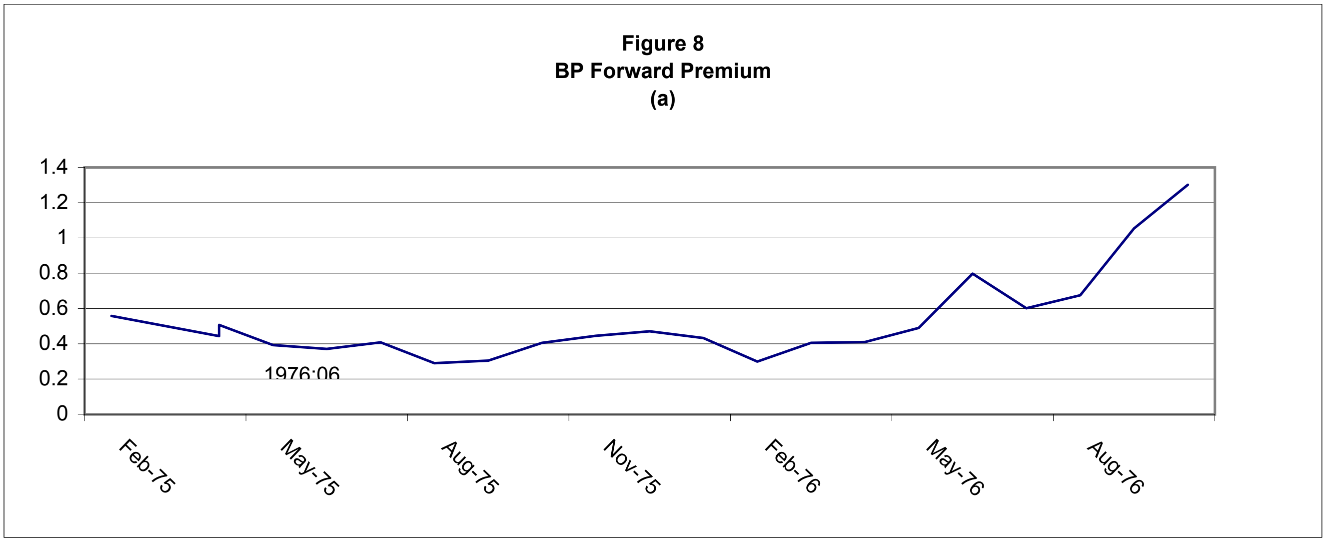


Figure 8
BP Forward Premium
(a)
1.4
1.2
1
0.8
0.6
0.4
0.2
0
1976:06
Feb-75
May-75
Aug-75
Nov-75
Feb-76
May-76
Aug-76

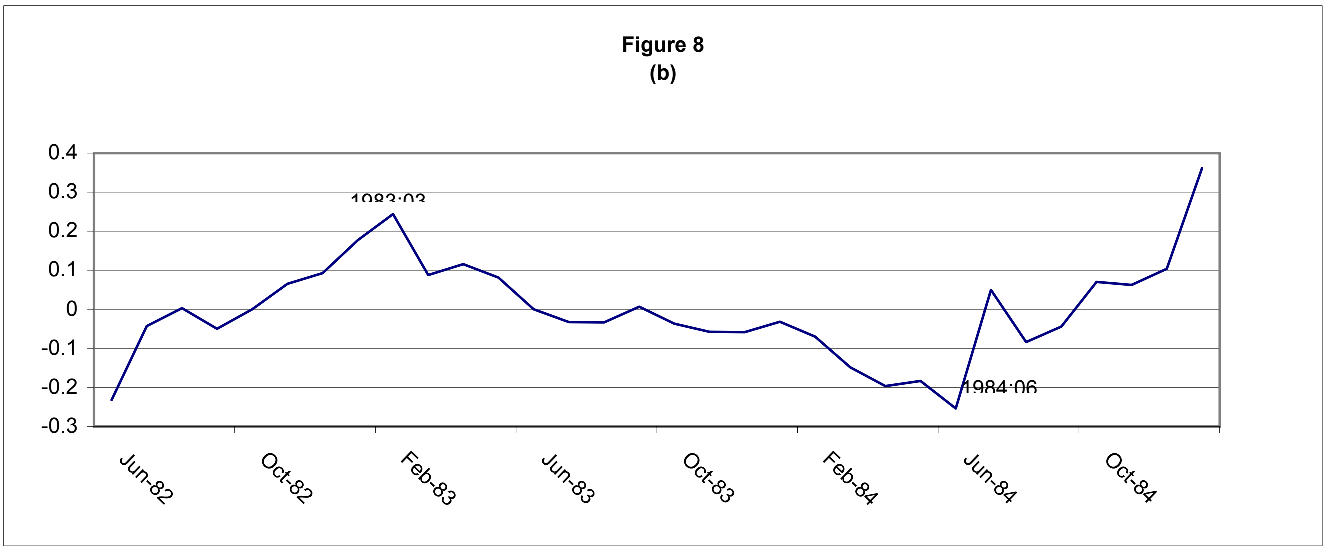


Figure 8
(b)
0.4
0.3
0.2
0.1
0
-0.1
-0.2
-0.3
1983:03
1984:06
Jun-82
Oct-82
Feb-83
Jun-83
Oct-83
Feb-84
Jun-84
Oct-84

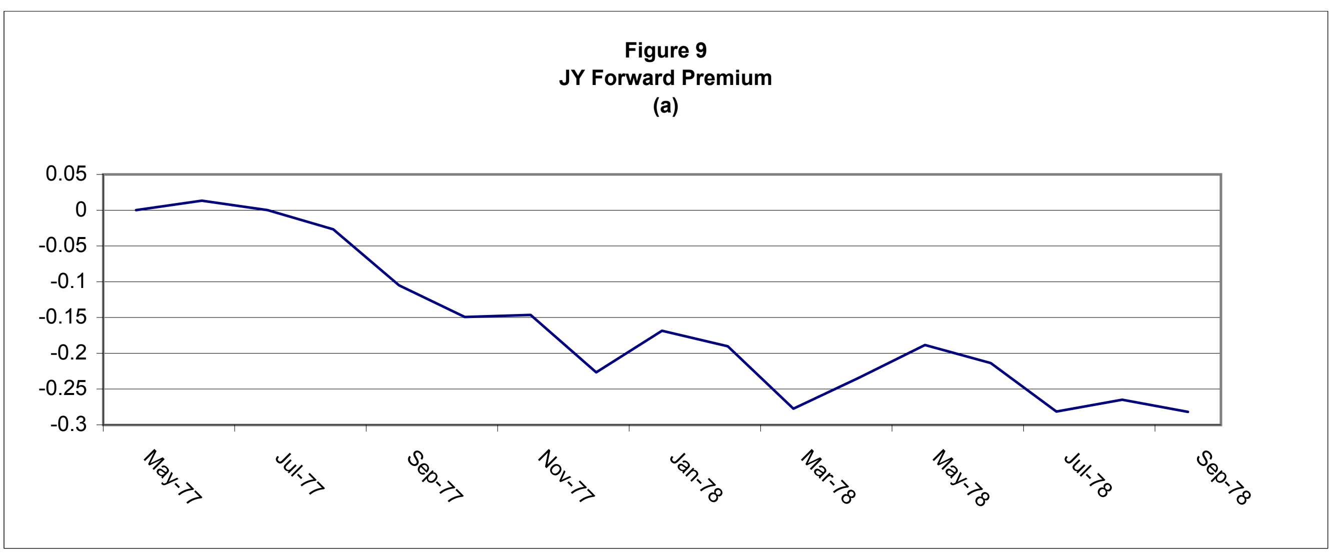


Figure 9
JY Forward Premium
(a)
0.05
0
-0.05
-0.1
-0.15
-0.2
-0.25
-0.3
May-77
Jul-77
Sep-77
Nov-77
Jan-78
Mar-78
May-78
Jul-78
Sep-78

Figure 9
(b)

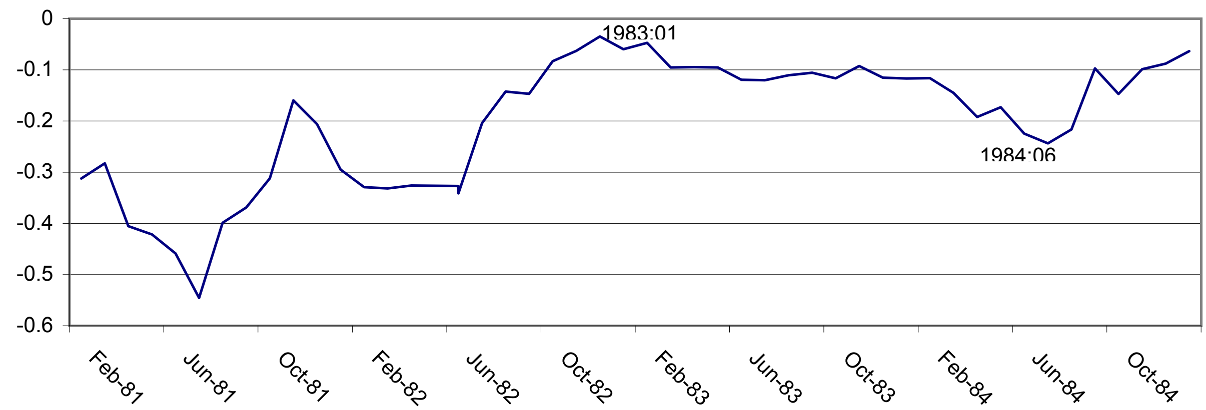


Figure 9
(c)

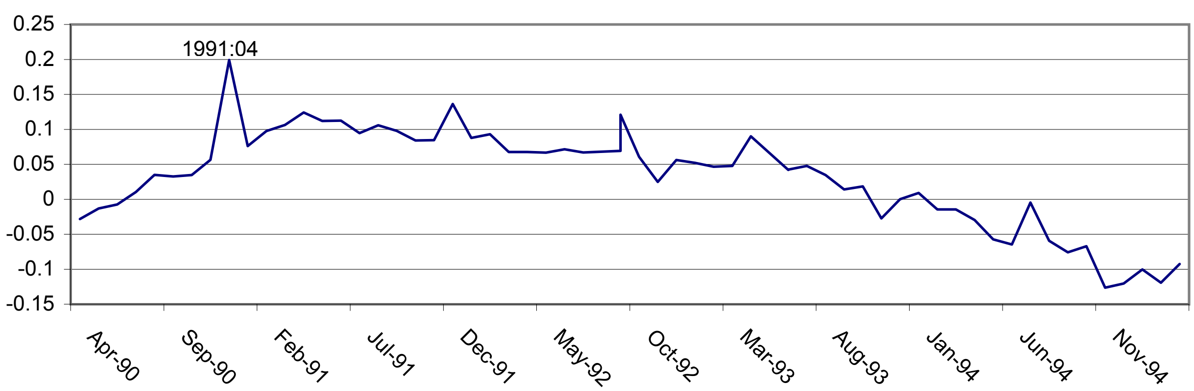


Figure 10
Excess Return BP/$
(a)

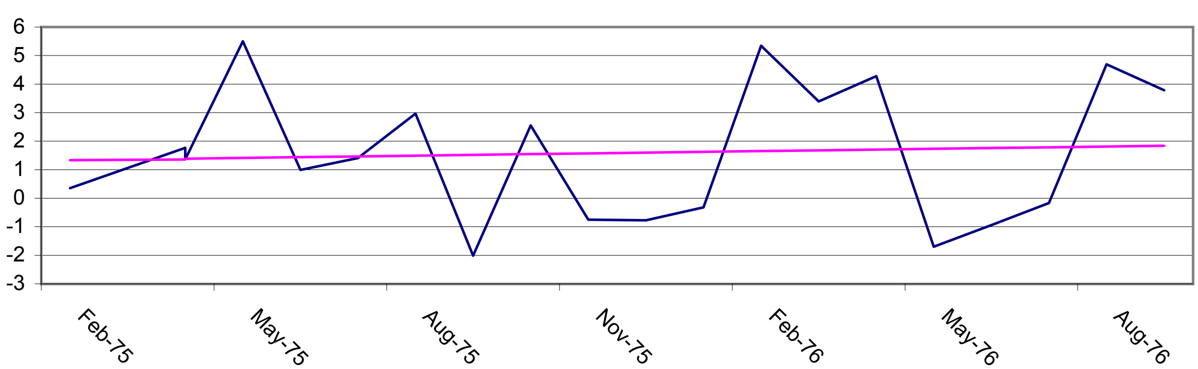


Figure 10
(b)

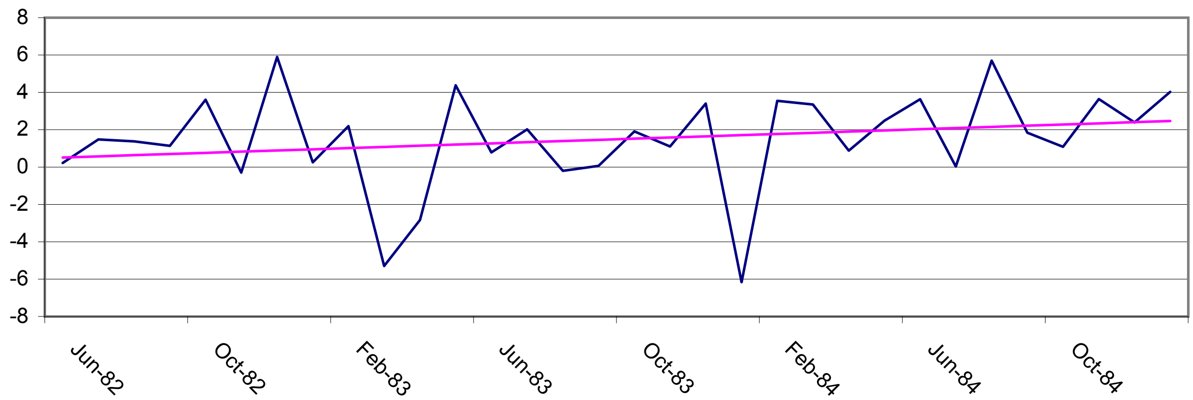


Figure 11
Excess Return: DM/$
(a)

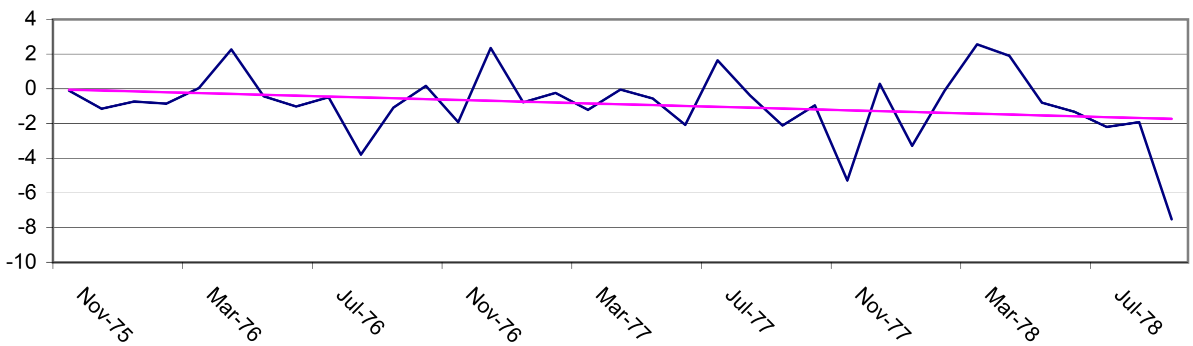


Figure 11
(b)

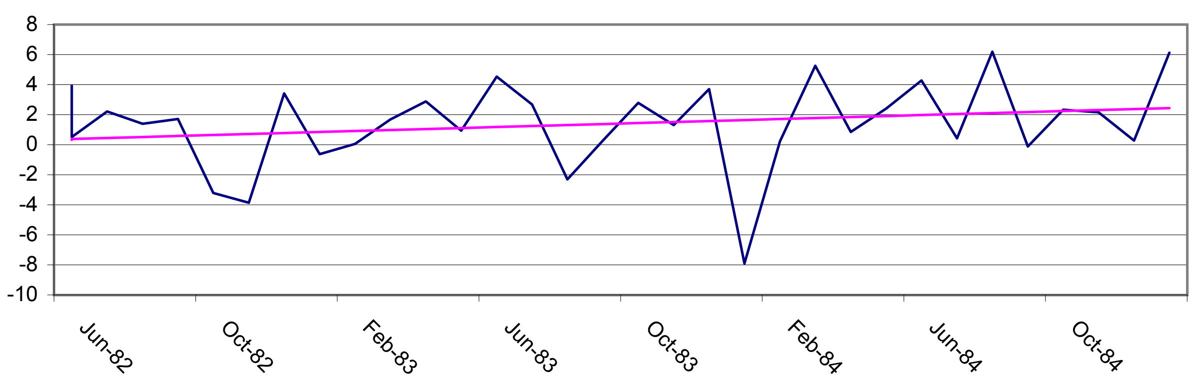


Figure 11
(c)

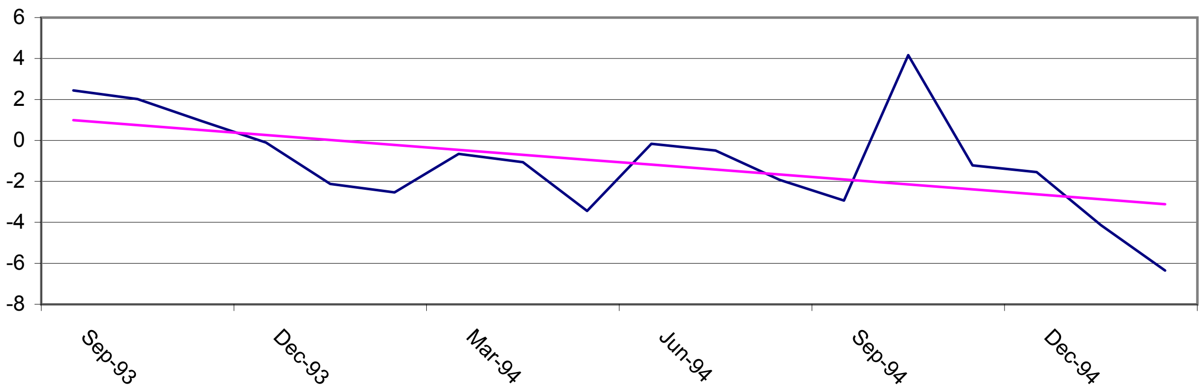


Figure 12
Excess Returns: JY/$
(a)

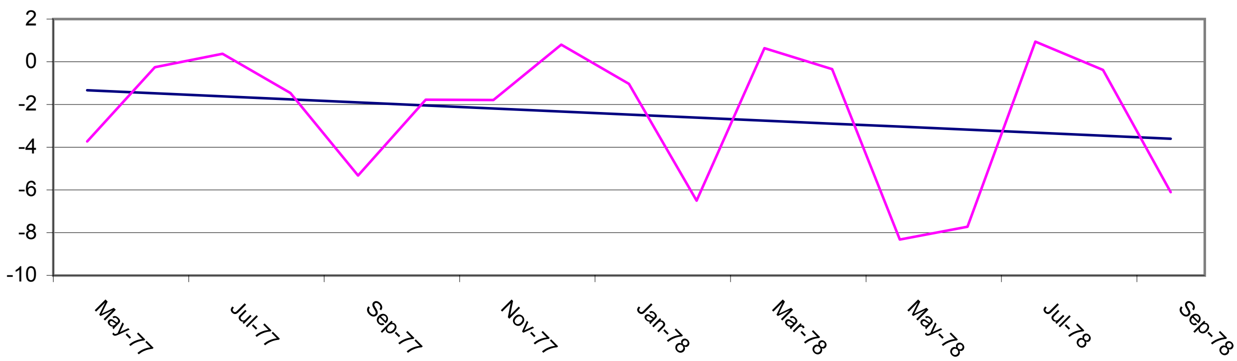


Figure 12
(b)

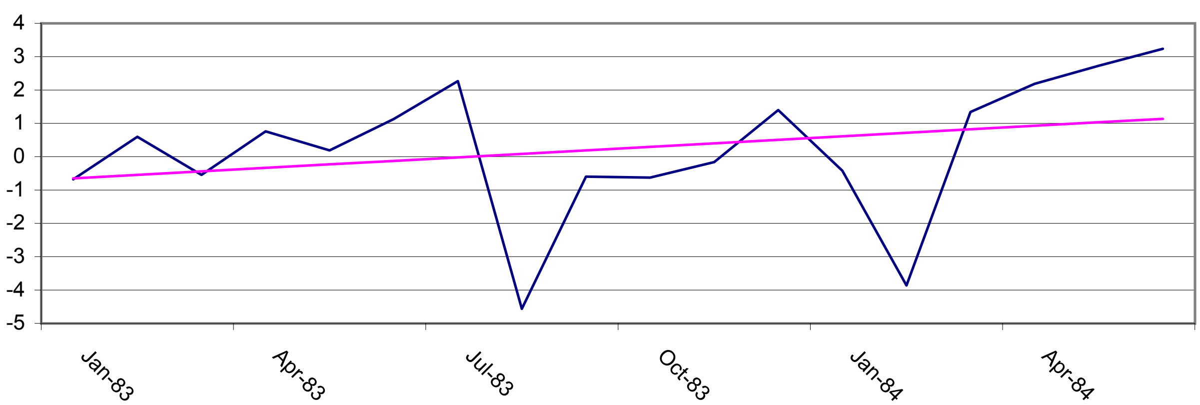
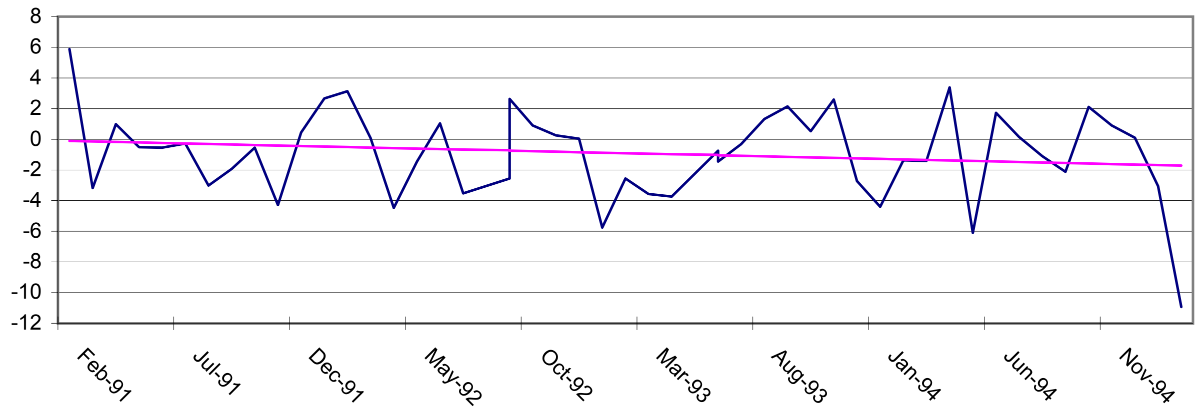


Figure 12
(c)


8
6
4
2
0
-2
-4
-6
-8
-10
-12
Feb-91
Jul-91
Dec-91
May-92
Oct-92
Mar-93
Aug-93
Jan-94
Jun-94
Nov-94