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SHORT RATES AND EXPECTED ASSET RETURNS

Kenneth A. Froot

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ABSTRACT

We present evidence that short-term interest rates forecast excess returns on many alternative assets: foreign exchange, stocks, bonds, and commodities. On average, a one-percentage-point increase in short rates is associated with three percent lower annualized excess returns. To test whether this predictability is attributable to time-varying risk, independent measures of excess returns are formed using survey data on expected returns. We find similar predictability in these measures, too. Since the surveys don't include risk premia, the predictable components cannot be attributed to risk. We suggest that when short rates are high (low) investors are excessively optimistic (pessimistic) about alternative-asset returns.

Kenneth A. Froot
Graduate School of Business
Harvard University
Soldiers Field Road
Boston, MA 02163

1. Introduction

It is by now well established that excess returns on many financial assets are partially predictable. Variables useful for explaining returns have been found in each of a growing list of asset markets. It is well known that one type of variable in particular seems to forecast returns across many markets: deviations of asset prices from their fundamental values. Specifically, prices revert toward measures of their own fundamental values more rapidly than would be predicted by models with constant risk premia.[1]

In this paper we focus on another variable that seems to have substantial, but less-well-known predictive power for excess returns: the short-term interest rate.[2] When the U.S. short rate is high, excess returns on foreign exchange, stocks, bonds, and commodities all appear *simultaneously* to be low. This apparent systematic correlation of the short rate with excess returns across assets is very different from the tendency for individual asset prices to revert toward their own fundamentals. Whereas mean reversion might be suggestive of idiosyncratic predictable components in prices, the forecastability of alternative-asset returns using interest rates is evidence of a predictable component that is common across markets. This common component implies that higher short-term interest rates are associated either with relatively *lower* risk of holding other assets, or with excessively optimistic expectations that these alternative assets will match the high promised return

[1] Cutler, Poterba, and Summers (1989) present results on mean reversion for four different markets: stocks, bond, foreign exchange and commodities. Flood, Hodrick and Kaplan (1986), Keim and Stambaugh (1986), Campbell and Shiller (1988), Poterba and Summers (1988), and Fama and French (1988a, b) document the tendency for mean reversion in the stock market. Huisinga (1986) investigates mean reversion in real exchange rates.

[2] Several papers have explored the interest rate's predictive power for certain asset returns. See, for example, Ferson (1989).

on short-term deposits.

This paper first documents the ability of short rates to forecast returns on a number of assets. We present evidence in the foreign exchange, stock, bond, and commodity markets that excess returns are low when short rates are high. The magnitude of the effect appears both substantial and very similar across markets: a one-percentage-point increase in the (annualized) short rate is associated on average with about a three-percentage-point reduction in (annualized) excess returns over the same horizon.[3]

We then go on to report positive cross-market evidence that consistently suggests this forecastability is not generated by risk. Specifically, we use a variety of survey measures of unexpected holding returns in the foreign exchange and bond markets as a complement to the usual approach of studying total holding returns. The survey measures – unlike total excess returns – do not include a risk premium, and therefore can provide independent information on the importance of risk as an explanation for the predictability we document.

The results from these data are striking. First, in almost every case in which surveys are available, we find that the survey expected returns are informative, in that, *ceteris paribus*, they are highly positively correlated with short-term interest rates (as would be suggested by simple models of relative asset pricing). The implication is that when short rates are high, expected nominal returns on alternative assets also tend to be high. Second, for explaining predictable excess returns it makes no difference whether the survey unexpected returns or total excess holding returns are used: *the predictable component in the two measures is the same.* Thus, unless the surveys happen to mismeasure systematically the market's expectation in such a way that the measurement error is perfectly correlated with the risk premium, time-variation in risk cannot explain our results.

Taken together, these findings may suggest that waves of optimism and pessimism can strike numerous speculative markets at one time, rather than merely affecting isolated

[3] Our results are strengthened by the many studies in different markets that demonstrate bias of similar magnitude in the forecasts implied by forward rates. The similarities between those results and ours are discussed in subsequent sections.

assets in an independent way. While with respect to a single asset the short rate might reasonably be thought of as exogenous, in equilibrium short rates are likely to be endogenously determined together with expected returns on all other assets. In this sense, movements in the interest rate, regardless of their ultimate source, may serve as partial indicators of the markets' overall outlook. This view suggests that when short rates are high, investors appear willing to go on holding alternative assets because of high expected returns, and not because the perceived risks of holding those assets are relatively low. Unfortunately, unusual optimism (or pessimism) in investor expected returns has systematically not been validated across markets during our samples.

Naturally, these systematic in-sample expectational errors need not be interpreted as evidence of market irrationality. Results like ours could in principle be explained by peso problems or learning on the part of purely rational investors. If the samples are small and/or otherwise unrepresentative of the ergodic behavior of asset returns, standard inference procedures will be invalid. Such arguments seem increasingly difficult to make, however, as economists are rapidly uncovering more of the same predictability in returns on new instruments, over new time periods, and for new forecast horizons, effectively increasing the size of the statistical sample.

Thus in short, expected returns across assets – as reflected in the short rate – appear excessively volatile. Investors would do better to reduce their expected returns on alternative assets when short rates are high and raise their expected returns when short rates are low.

The paper is organized as follows. Section 2 investigates the ability of the short rate to forecast excess returns on foreign exchange and shows how this predictability can be interpreted as excessive volatility in expected returns on foreign exchange. Sections 3, 4, and 5 follow similar procedures for the stock, bond, and commodity markets, respectively. Section 6 offers interpretations and conclusions.

2. Excess forecast volatility in expected returns on foreign exchange

This section presents evidence that expected returns on foreign exchange are excessively volatile, in that they appear to move too much with current short rates to be rational in the sense of Muth. We will see this implies that interest rates can forecast excess returns on foreign exchange.

The first step is to develop a framework for evaluating whether expected returns are excessively or insufficiently volatile. The expected return on foreign exchange is equal to the expected percentage depreciation of the dollar plus the foreign (j-period) interest rate, or alternatively, the U.S. short (j-period) rate minus a residual, which we term the risk premium on dollar assets:[4]

$$\Delta \mathbf{s}^e_{t+j} + \mathbf{i}^*_t = \mathbf{i}_t - \mathbf{rp}_t, \tag{1}$$

where $\Delta \mathbf{s}^e_{t+j}$ is the expected log percentage change in the spot rate (expressed in dollars per unit of foreign currency) between times t and $t+j$, conditional on all information available at time t. Equation (1) implies that expected depreciation can be written as the interest differential less the risk premium, $\Delta \mathbf{s}^e_{t+j} = \mathbf{i}_t - \mathbf{i}^*_t - \mathbf{rp}_t$.

Consider a regression of the expectational error made by investors in predicting the future spot exchange rate on their expected rate of currency depreciation:

$$\Delta \mathbf{s}_{t+j} - \Delta \mathbf{s}^e_{t+j} = \alpha + \beta \Delta \mathbf{s}^e_{t+j} + \eta_{t+j}, \tag{2}$$

where $\Delta \mathbf{s}_{t+j}$ is the realized log percentage change in the spot rate between times t and $t+j$. To fix ideas, let us for now suppose that the market's expected rate of depreciation is actually observable. In testing (1), we obviously would not impose the rational-expectations restriction that $\Delta \mathbf{s}^e_{t+j}$ is equivalent to the mathematical expectation of $\Delta \mathbf{s}_{t+j}$ conditional on all information at time t. Under that restriction, which is the null hypothesis in (1), $\alpha = \beta = 0$ and the residual, η_{t+j}, is purely random.

[4] In this paper we work with nominal asset-market returns to avoid using poorer-quality indexes of goods-market prices. This implies that the "risk premium," $\mathbf{rp}_t$, is defined so as to include a term which arises entirely from the correlation between returns (e.g. on foreign exchange) and the unexpected change in the price of goods. See the appendix for a complete derivation of this risk premium.

The alternative hypotheses are that expected depreciation displays excessive ($\beta < 0$) or insufficient ($\beta > 0$) forecast volatility.[5] To understand this, suppose that $\beta < 0$. This would imply that when expected depreciation is above (below) its mean, excess returns are systematically lower (higher) than expected. An investor would do better – in that the variance of his forecast error, $\Delta s_{t+j} - \Delta s^e_{t+j}$, would be lower – if he systematically reduced his current expectation of depreciation fractionally toward its mean. This just formalizes the notion that when investors are relatively optimistic, they tend to be too optimistic. Similarly if $\beta > 0$, an investor could improve his forecast by scaling up multiplicatively the deviation between his current expected depreciation and its mean.

In order to distinguish the U.S. rate's role in generating predictable returns a specification similar to (1) can be used, in which the prediction error is regressed on the components of expected deprecation:

$$\Delta s_{t+j} - \Delta s^e_{t+j} = \alpha + \beta_1 i_t + \beta_2 i^*_t + \beta_3 rp_t + \eta_{t+j}, \tag{3}$$

where, once again, the null hypothesis is that $\beta_1 = \beta_2 = \beta_3 = 0$ and the residual is purely random. Equation (3) yields several distinct alternative hypotheses. They are that expected depreciation is excessively volatile with respect to the U.S. short rate, $\beta_1 < 0$, the foreign short rate, $\beta_2 > 0$, and perceptions of risk, $\beta_3 > 0$. Opposite inequalities are associated with insufficient forecast volatility.

While the market's expectation of future depreciation, Δs^e_{t+j}, is unobservable, inferences about its behavior can nevertheless be drawn by using two proxies. The first and most common measure of expected depreciation is the forward discount, $fd_t = f_t - s_t$, where f_t is the log j-period forward rate. By arbitrage, the forward discount is equal to the interest differential, $fd_t = i_t - i^*_t$.[6]

[5] See Froot (1989a) for broader applications of excess forecast volatility. Bilson (1981) has termed the same alternative hypothesis $\beta < 0$ in (1) "excessive speculation."

[6] When using the forward discount or interest differential to measure expected depreciation, the specification in (3) is closely related to standard tests of forward-rate unbiasedness in which the forward-rate prediction error is regressed on the interest differential alone. Hodrick (1988) gives a thorough summary of the literature testing unbiasedness. In addition to estimating (3), we perform standard unbiasedness tests below, with results similar to those found elsewhere.

This measure's obvious advantage is that it is readily observable. Its disadvantage is that by definition it includes any time-varying risk premium that might separate expected depreciation from the interest differential. From (1), $\mathbf{fd}_t = \mathbf{i}_t - \mathbf{i}_t^* = \Delta \mathbf{s}_{t+j}^e + \mathbf{rp}_t$. Thus, under this measure the prediction error on the left-hand side of (3) is just the total excess return on foreign exchange, $\Delta \mathbf{s}_{t+j} - \mathbf{fd}_t = \Delta \mathbf{s}_{t+j} - \Delta \mathbf{s}_{t+j}^e - \mathbf{rp}_t$. As this measure includes $-\mathbf{rp}_t$, the null hypothesis in (3) becomes $\beta_1 = \beta_2 = 0$ and $\beta_3 = -1$.

Our second proxy for the market's expectation of depreciation comes from time-t survey measures of exchange rate expectations. This measure, denoted by $\Delta \mathbf{s}_{t+j}^s$, is useful because it is not contaminated by a risk premium. Although the median survey response is likely to report the market's expectation with error, random measurement errors do not pose a problem for the estimation because $\Delta \mathbf{s}_{t+j}^e$ is on the left-hand side of our regressions. Under this measure of expected depreciation, the dependent variable in (3) is the realized survey prediction error, $\Delta \mathbf{s}_{t+j} - \Delta \mathbf{s}_{t+j}^s$.[7]

Since one of our expectations measures contains a risk premium while the other does not, we will be interested in seeing if the two measures yield similar coefficient estimates when used on the right-hand side of (3). If they don't, then the risk premium remains a potential explanation for the predictability of excess returns. If they do, however, then risk is unlikely to be responsible for our findings.

On the right-hand side of (3), only $\mathbf{rp}_t$ is unobservable. Once again, there are two possible measures. The first assumes that the forward discount is equal to expected depreciation, so that the risk premium is zero (or a constant). This implies that consistent estimates of β_1 and β_2 can be obtained without including the premium in the regressions. The second is the survey measure of the risk premium, given by $\mathbf{rp}_t^s = \mathbf{fd}_t - \Delta \mathbf{s}_{t+j}^s$.

Note that this latter risk measure provides another way of determining whether

[7] See Frankel and Froot (1987) or Froot and Frankel (1989) for a more general discussion of the use of survey data to measure exchange rate expectations. Note that it is not valid to object to the use of surveys on the grounds that they may appear to be inefficient forecasts of the future spot rate. We wish to test whether any forecasting inefficiencies that may exist are consistent with either excess or insufficient forecast volatility. If we were ex ante to disqualify measures of expectations because they turned out ex post to be inefficient, our tests would obviously be biased against rejection.

changes in risk are responsible for the predictable component of $\Delta s_{t+j} - fd_t$. If risk is in fact the culprit, then by including rp_t^s on the right-hand side the interest-rate coefficients should fall to zero. Alternatively, if adding rp_t^s has no effect on the estimates of β_1 and β_2, then risk is unlikely to be responsible for the results.

Before proceeding, several econometric issues should be mentioned. First equation (3) and the other equations that follow can be estimated using OLS with with standard errors calculated using Hansen's (1982) Generalized Method-of-Moments (GMM) using a Newey and West (1987) correction. Where appropriate, the covariance matrix estimators allow for serial as well as contemporaneous correlation. Estimators were computed twice for each regression: once under the assumption of homoskedasticity and again allowing for unknown conditional heteroskedasticity. Due to the downward finite-sample bias of the heteroskedasticity-consistent GMM covariance estimates, we try to be conservative by reporting the larger of two sets of standard errors of the coefficients. While this estimation procedure may be inefficient, the conclusions are unlikely to change using more efficient techniques.

Another potential difficulty in estimating (3) stems from the possibility that interest rates contain a unit root. Although many tests do not reject the unit root hypothesis, the low power of these tests against sensible alternatives suggests that there is little positive evidence for such a unit root. Nevertheless, in order to avoid these issues, an additional version of (3) is estimated with $\beta_1 = -\beta_2$, so that the regressors are the interest differential and measures of the risk premium.[8]

2.1. Results

Table 1 presents estimated versions of (3). In the top panel, we report estimates of (3) in its unrestricted form; in the middle panel we impose the restriction $\beta_1 = -\beta_2$, so that the first regressor is the interest differential; in the bottom panel we impose the restriction $\beta_2 = 0$ to eliminate the foreign interest rate, i_t^*, as a regressor. The data are monthly over

[8] See Baillie and Bollerslev (1989) for evidence that the interest differential (or forward discount) is stationary.

the floating rate period, 1973-1986. Since there are no surveys available over this time period, the forward discount is the only available proxy for expected depreciation. The estimates of β_1 are statistically negative, on average equal to minus three. The estimates of β_2 are generally positive (which implies that foreign residents tend to earn low returns on *dollar* investments when their own interest rates are high), although less statistically significant.

Tables 2a through 2c present estimates of (3), for the sample over which there are survey data, 1981-88. The estimates stack 5 currencies against the dollar (the pound, French franc, Deutsche mark, Swiss franc and yen) in order to save space.[9] The first line of each panel reports the specification in (3), omitting the risk premium on the right-hand side. The second line then adds the survey measure of the premium as an additional independent variable. The third and fourth lines follow the same pattern except, to guarantee stationarity, the interest differential is used as a regressor in place the individual interest rates. The fifth and sixth lines omit the foreign interest rate altogether. Tables 2a, 2b, and 2c report different forecast horizons of 3, 6, and 12 months, respectively.

The top panels of Tables 2a-2c also use the forward discount to proxy for expected depreciation. A quick comparison with Table 1 makes it clear that the specific sample period, forecast horizon, and selection of currencies is of little qualitative importance, as the findings are very similar. Across these tables almost all of the estimates of β_1, and the majority of the estimates of β_2, are statistically different from zero, with the expected signs. Indeed, given the size of the estimated standard errors, the point estimates are surprisingly close, clustering around -3.[10] Since the forward discount contains a risk premium, the results in the lines one, three, and five of the top panels of Tables 2a-2c would usually be ascribed to time-variation in that premium.

[9] This procedure does not obscure much information: the estimates for four of the five individual currencies are not importantly different from either the aggregate measures we report in Tables 2a-2c or the individual-currency estimates reported in Table 1. The estimates for the yen are the only outlier. For that currency no statistically significant effect of interest rates on expectational errors (measured using either the forward rate or survey expectation) was found.

[10] The Durbin-Watson statistics are much lower in Tables 2a-2c because of the usual overlapping-observations problem. Our standard errors correct for this.

Note, however, that when the survey risk premium is included as a regressor (lines two, four, and six of the top panels in Tables 2a-2c) there is no change in the either the point estimates or the standard errors of β_1 or β_2, and that the estimate of β_3 is insignificantly different from zero.[11] If risk were responsible for the findings in odd-numbered lines, then the estimates of β_1 and β_2 in even-numbered lines should be indistinguishable from zero and the estimate of β_3 should be minus one. However, random measurement errors in rp_t^s could also bias the estimates in this direction.

The bottom panel of Tables 2a through 2c duplicates the regressions run in the top panel, only this time using the survey measure of expected depreciation on the left-hand side. Here, even if the surveys do contain random measurement error, the estimates remain unbiased. The results in Tables 2a-2c show without exception that $\beta_1 < 0$ and usually that $\beta_2 > 0$. The similarity of the signs and magnitudes of the coefficients in the top and bottom panels of each table is striking. This again suggests that any risk premium contained in the forward-rate prediction errors is not responsible for the predictability of excess returns. Indeed, even if the surveys were pure noise, they still do not contain a risk premium, and therefore risk cannot explain the similarity between the top and bottom panels.[12]

Neither the point estimates nor the estimated standard errors are importantly affected by the restriction that $\beta_1 = -\beta_2$ in the third and fourth lines of each panel.[13] This indicates that our results in the top two lines are not a direct consequence of any potential nonstationarity in interest rate levels. The estimates of β_1 in the fifth and sixth lines remain negative, but tend to be smaller in magnitude and somewhat less statistically significant.

The similarity of all the reported estimates of β_1 and β_2 supports the view that

[11] The survey data used to form rp_t^s come from the *Economist Financial Report*. Surveys were undertaken each 6 weeks from June 1981 to August 1988 on expectations of the dollar against the same five foreign currencies.

[12] Note that the estimates of β_3 in even-numbered lines increase by one when moving from the top to the bottom panel, whereas the corresponding estimates of β_1 and β_2 remain exactly the same. We expect this to happen because the difference between the top- and bottom-panel dependent variables is just rp_t.

[13] Froot and Frankel (1989) estimate regressions similar to those in the third line of each panel, and then use the surveys to argue, as we do here, that the interest differential's ability to forecast returns does not constitute evidence of a time-varying risk premium.

expected depreciation is excessively volatile, and that Δs^e_{t+j} is excessively sensitive to changes in short rates. When the return on short deposits is high, investors' appear to have overly optimistic expectations about the returns on competing assets like foreign exchange. Investors would do better if they reduced their expectations of depreciation when the short rate is above its mean value (and conversely when the short rate is below).

Even though the surveys don't contain a risk premium, one might ask whether they do contain any information at all, i.e., whether they are informative about expected depreciation. If ex ante rates of return on short bills in different currencies are approximately equalized, one would expect an increase in the interest differential to be associated with a one-for-one increase in expected depreciation. Froot and Frankel (1989) regress the survey measure of expected depreciation on the interest differential and find the coefficient is indeed close to one, while statistically much greater than zero.

3. Excess forecast volatility in expected stock-market returns

This section develops analogous regression tests for expected stock-market returns. To do this, first write the expected return on the market as equal to the short rate plus an "equity premium:"[14]

$$\mathbf{r}^e_{t+j} = \frac{\mathbf{P}^e_{t+j} - \mathbf{P}_t + \mathbf{D}_t}{\mathbf{P}_t} = \mathbf{i}_t + \psi_t, \tag{5}$$

where $\mathbf{P}^e_{t+j}$ is the market's expected stock price at time $t+j$ conditional on all information available at time t, $\mathbf{P}_t$ is the time-t stock price, $\mathbf{D}_t$ is the current dividend payment, and ψ_t is the equity premium. Equation (5) implies that the expected rate of price appreciation is given by the short rate less the dividend yield plus the equity premium,

$$\frac{\mathbf{P}^e_{t+j} - \mathbf{P}_t}{\mathbf{P}_t} = \mathbf{i}_t - \frac{\mathbf{D}_t}{\mathbf{P}_t} + \psi_t. \tag{6}$$

[14] Because we work with nominal returns, ψ_t includes a term attributable to the correlation between unexpected inflation and the excess stock return. See the appendix for more details.

Consider then a regression of the excess (risk-adjusted) return on the components of expected appreciation,

$$\mathbf{r}_{t+j} - \mathbf{r}^e_{t+j} = \alpha + \beta_1 \mathbf{i}_t - \beta_2 \frac{\mathbf{D}_t}{\mathbf{P}_t} + \eta_{t+j}, \tag{7}$$

where the dependent variable is equivalent the unexpected percentage change in prices, $\mathbf{r}_{t+j} - \mathbf{r}^e_{t+j} = \frac{\mathbf{P}_{t+j} - \mathbf{P}^e_{t+j}}{\mathbf{P}_t}$. We do not have any independent measure of the equity premium, so we assume it to be included in the constant term. The hypothesis that expectations are rational thus implies that $\beta_1 = \beta_2 = 0$ and the residual is purely random. The alternative hypothesis is that β_1 is less or greater than zero: that expected stock-market returns are excessively or insufficiently volatile, respectively, with respect to the short rate.

From (5), the expected total stock return is equal to the interest rate plus an equity premium. This suggests an alternative regression to (7), in which the excess return is regressed on the short rate alone. We try this specification in addition to (7) below.

3.1. Results

We use the monthly value-weighted index from the Center for Research in Securities Prices (CRSP) for the stock return data. The series runs from 1926 to 1985. Monthly interest rates on U.S. government securities with approximately one month to maturity come from Ibbotson Associates (1986).[15] In order to take advantage of high frequency stock returns we also used two measures of seven-day interest rates: the rate on eurodollar deposits and repurchase agreements.

Table 3a presents estimates of (7). The results are similar to those in the previous tables, although the coefficients appear different during early portions of the sample. Over most of the post-war sample, however, increases in short rates reliably result in negative excess stock market returns.[16]

[15] This standard data set is used by Marsh and Merton (1987), Fama and French (1987), Campbell and Shiller (1988), and Poterba and Summers (1987), among others.

[16] If short rates contain a unit root, then the standard errors in Table 3a may be biased. The next version of this paper will include probability values from Monte Carlo simulations to address this potential problem.

Table 3b presents similar estimates deleting the dividend yield regressor.[17] Without dividends, estimates from higher-frequency data can be obtained. The first row reports the results from weekly data over the period 1973-84, using the seven-day eurodollar interest rate. The second row of Table 3b uses a seven-day interest rate on repurchase agreements collateralized by U.S. government securities, available from DRI beginning in 1980. Both estimates of β_1 are statistically negative at the one-percent level. In rows 3 through 10 of Table 3b we report estimates for longer horizons (one year and one month) over the full Ibbotson sample and over a number of subsamples. All but one of the estimates of β_1 are less than zero, though none is as large or statistically significant as in the weekly data.

Unfortunately, there is no second measure of expected stock returns to appeal to for further evidence on whether the above results are generated by time-varying risk. However, note that the point estimates in Tables 3a and 3b are similar to those in the foregoing tables.[18]

4. Excess forecast volatility in expected bond returns

Our third set of tests covers expected returns in the bond market. Under the linearized model of the term structure of interest rates, the excess (risk-adjusted) holding period-return is proportional to the market's expectational error in predicting the future interest rate:

$$\mathbf{h}_{t+j}^{(j,k)} - (\mathbf{h}_t^{(j,k)})^e = \frac{d_k}{d_k - d_j}\left(\mathbf{i}_{t+j}^{(k-j)} - (\mathbf{i}_{t+j}^{(k-j)})^e\right), \tag{8}$$

where $\mathbf{h}_{t+j}^{(j,k)}$ is the realized excess holding-period yield obtained from purchasing a k-period bond at time t, holding it for j periods, and then selling it at time $t+j$, $(\mathbf{h}_t^{(j,k)})^e$ is the corresponding market expected excess holding-period yield at time t, $\mathbf{i}_{t+j}^{(k-j)}$ is the realized

[17] Since expected returns are given simply by $\mathbf{i}_t + \psi_t$, the dividend yield $\frac{\mathbf{D}_t}{\mathbf{P}_t}$ does not have to be included on the right-hand side to test for excess volatility of expected returns.

[18] A number of authors have found evidence of negative correlation between short-term nomihal interest rates and subsequent stock market returns, both in the U.S. and in other industrialised countries (see, for example, Fama and Schwert, 1977 and Solnik, 1983). This correlation is usually interpreted as evidence that expected stock returns respond negatively to expected inflation. Of course, under this interpretation higher inflation (and higher nominal short rates) must be associated with lower equity premia on stocks.

rate at time $t+j$ on a $k-j$-period bond $(k > j)$, $(\mathrm{i}_{t+j}^{(k-j)})^e$ is the market's time-t expectation of $\mathrm{i}_{t+j}^{(k-j)}$, d_m is Macaulay's (1938) definition of duration for an m period bond when priced at par, $d_m = \frac{1-(1+\bar{\mathrm{i}})^{-m}}{1-(1+\bar{\mathrm{i}})^{-1}}$, and $\bar{\mathrm{i}}$ is the coupon rate. For pure-discount bonds, such as U.S. Treasury bills, duration is just the time to maturity.[19] Also under the linearized model, the expected future interest rate above the "short" j-period rate can be written as a linear combination of the "long" k-period rate, the short rate, and a term premium:[20]

$$(\mathrm{i}_{t+j}^{(k-j)})^e - \mathrm{i}_t^{(j)} = \frac{d_k}{d_k - d_j}\left(\mathrm{i}_t^{(k)} - \mathrm{i}_t^{(j)}\right) - \theta_t^{(j,k)}. \tag{9}$$

As in the previous sections, consider a regression of the excess holding return for a bill or bond on the components of the expected interest rate change:

$$\mathbf{h}_{t+j}^{(j,k)} - (\mathbf{h}_t^{(j,k)})^e = \alpha + \beta_1 \mathrm{i}_t^{(j)} + \beta_2 \mathrm{i}_t^{(k)} + \beta_3 \theta_t^{(j,k)} + \eta_{t+j}. \tag{10}$$

The null hypothesis in (10) is that $\beta_1 = \beta_2 = \beta_3 = 0$ and the residual is purely random, whereas the alternative hypotheses are that expected interest rate changes are excessively volatile with respect to short rates $(\beta_1 < 0)$, long rates $(\beta_2 > 0)$, and term premia $(\beta_3 > 0)$. Insufficient volatility is associated with the opposite inequalities.

As in section 2, there are two available measures of the expected future interest rate, $(\mathrm{i}_{t+j}^{(k-j)})^e$, which from (8) appears implicitly on the left-hand side of (10). First is the standard one – the time-t forward interest rate on a $k-j$-period instrument to be acquired in j periods: $(\mathrm{i}_{t+j}^{(k-j)})^f = \frac{d_k}{d_k-d_j}\left(\mathrm{i}_t^{(k)} - \mathrm{i}_t^{(j)}\right) + \mathrm{i}_t^{(j)}$. To see that this measure contains the term premium, use (9) to get $(\mathrm{i}_{t+j}^{(k-j)})^f = (\mathrm{i}_{t+j}^{(k-j)})^e + \theta_t^{(j,k)}$. Of course, when using the forward rate to measure the expected future interest rate, the time-t expected excess holding return is zero, so the dependent variable in (10) is simply the total excess holding return, $\mathbf{h}_{t+j}^{(j,k)}$.[21]

[19] For an exposition of the linearised model of the term structure and for evidence on the size of the linearisation error, see Shiller, Campbell and Schoenholtz (1983).
[20] See the appendix for a general definition of this premium.
[21] Regressions of total excess holding returns on the right-hand side variables in (10) are closely related to those used to test the expectations hypothesis. See Shiller (1988) and Campbell and Shiller (1989) for empirical overviews of such tests.

Our second measure of the expected future interest rate, given by $(\mathrm{i}_{t+j}^{(k-j)})^s$, comes from time-$t$ surveys of interest rate expectations of the $k-j$-period rate at time $t+j$. The surveys allow us to compute the realized excess holding return in (8) a second way, which we denote as $\mathrm{h}_{t+j}^{(j,k)} - (\mathrm{h}_t^{(j,k)})^s$. Because this measure of excess holding returns is computed directly from expectations, without reference to forward rates, it does not include a term premium. Thus, any predictable component found in this measure of excess returns cannot be attributed to risk.

On the right-hand side of (10) is the term premium, $\theta_t^{(j,k)}$, which can also be measured two ways. The first comes from the constant-term-premium hypothesis, which implies that we can get consistent estimates of β_1 and β_2 without bothering to include the premium at all in our regressions. The second comes from the survey measure of the term premium, $(\theta_t^{(j,k)})^s$, which can be computed from (9) using the survey expectation in place of $(\mathrm{i}_{t+j}^{(k-j)})^e$. If including $(\theta_t^{(j,k)})^s$ has no effect on the coefficients β_1 and β_2, then risk is unlikely to be the explanation for any predictability based on short and long rates.

4.1. Results

Estimates of equation (10) are reported in Tables 4a through 4f for a number of different instruments and forecast horizons. As in section 2, each table contains two panels; the top uses the total excess holding return, $\mathrm{h}_{t+j}^{(j,k)}$, as the dependent variable, while the bottom uses survey unexpected holding return, $\mathrm{h}_{t+j}^{(j,k)} - (\mathrm{h}_t^{(j,k)})^s$. In order to permit comparison between panels, both samples are constructed on the survey dates (the last Friday of each quarter from 1969 to 1986).[22]

The first line in each panel includes only short and long rates as independent variables; the second line then adds the survey term premium. The third and fourth lines are similar, except that the spread, $\mathrm{i}_t^{(j)} - \mathrm{i}_t^{(k)}$ is used in place of individual short and long rates. If long and short rates are nonstationary but contain a common unit root, then the spread will

[22] As with the exchange rate, there does not appear to be anything unusual about this particular sample. For many instruments and forecast horizons, regressions of $\mathrm{h}_{t+j}^{(j,k)}$ on the difference between the associated long and short rates yield coefficients similar to those reported below.

be stationary.[23] Finally, in the fifth and sixth lines the long rate is omitted altogether.

The interest-rate survey data come from surveys conducted by the *Goldsmith-Nagan Bond and Money Market Letter*, now published in the investor newsletter, *Reporting on Governments*. Expectations are of future interest rates on the Bond Buyer Index, the 30-year mortgage rate, and 12-month Treasury bills.[24] Forecast horizons are three and six months.[25]

Tables 4a and 4b cover the Bond Buyer index at forecast horizons of 3 and 6 months, respectively. Tables 4c and 4d do the same for 30-year mortgages, while Tables 4e and 4f cover 12-month Treasury bills. In each of these six tables, all of the estimates of β_1 are negative. In addition, most are statistically significant and have magnitudes which roughly correspond to those in prior tables. Most estimates of β_2 are statistically positive. Also, the inclusion of $\theta_t^{(j,k)}$ in lines two, four, and six of each panel has no effect on the interest rate coefficients. Finally, the estimates in the top and bottom panels are very similar, suggesting that the term premium is not responsible for the predictability of standard measures of excess holding returns.[26]

The results in Tables 4a through 4f suggest that excess returns on longer-term bills and bonds exhibit excess forecast volatility. In particular, expected returns respond excessively positively to increases in the short rate. When the short rate is high (holding constant the long rate and term premium), future rates expected by investors are low, and therefore investors' expected returns are high. Since on average in such circumstances realized future rates do not turn out to be so low, the realized holding returns on bonds systematically do not turn out to be as high as expected.[27][28]

[23] The regressions in the third line of each panel are equivalent to those used in Froot (1989b) to test the expectations hypothesis. See Campbell and Shiller (1987) for evidence that the spread is stationary.

[24] The Bond Buyer indexes 20 general obligation issues with 20-year maturities. The index is designed to reflect the current yield-to-maturity on new issues.

[25] See Froot (1989b) for more detail on these data.

[26] As in section 2, when the survey risk premium is included as a regressor, there is no change between top and bottom panels in estimates of β_1 and β_2, and the estimate of β_3 increases by one from the top to the bottom panel.

[27] The weakest evidence for this hypothesis is found in Tables 4e and 4f, which use 12-month Treasury bills. Although the signs of the coefficients are as expected, their estimated magnitudes do appear to shift between top and bottom panels.

[28] Many authors, such as Mankiw and Summers (1984), find that the long rate is *insufficiently* volatile with respect to changes in the short rate. The description above should help clarify how their results are consistent with our finding of *excessive variability* of expected returns. If when the short rate rises, the long rate tends to rise as well (instead of remaining

As in section 2, it is consoling to have some positive evidence that the survey measures are indeed informative about expected interest rate changes. Froot (1989b) shows that the survey expected interest rate changes are highly positively correlated with the difference between long and short rates. Holding risk constant, of course, expected future interest rates should increase whenever the long rate rises relative to the short rate.[29]

5. Commodity Markets

Our final application is to commodity markets. As in the earlier examples, we wish to regress the market's expectational error in predicting the future return on commodities on the market's expected rate of return. The market's expected return from time t to $t+j$ can be expressed as the rate of price appreciation less storage costs:

$$\mathbf{c}^{e}_{t+j} - \frac{\mathbf{S}_t}{\mathbf{C}_t} = \mathbf{i}_t + \gamma_t, \tag{11}$$

where $\mathbf{c}^{e}_{t+j}$ is the time-t expected percentage price appreciation from t to $t+j$, $\frac{\mathbf{S}_t}{\mathbf{C}_t}$ is the storage cost over j periods at time t expressed as a percentage of the commodity price $\mathbf{C}_t$, and γ_t is a commodity risk premium.[30]

Thus our regression equation becomes:

$$\mathbf{c}_{t+j} - \mathbf{c}^{e}_{t+j} = \alpha + \beta_1 \mathbf{i}_t + \beta_2 \frac{\mathbf{S}_t}{\mathbf{C}_t} + \eta_{t+j}, \tag{12}$$

where we have left off the risk premium, γ_t, from the right-hand side, as we have no survey data with which to measure it directly. As above, the null hypothesis is that $\beta_1 = \beta_2 = 0$ and the residual is purely random; the alternatives that expected appreciation is excessively volatile imply that $\beta_1 < 0$ and/or $\beta_2 < 0$.

constant), future rates expected by investors are not so low, and therefore expected returns are not so high. Thus, if the long rate were to rise sufficiently when the short rate increases (again holding risk constant), the predictability of returns based on the spread would be eliminated.

[29] One might interpret these results as suggesting that the survey respondents merely report the forward rate instead of their actual expectation. Froot (1989b) shows that the responses differed substantially from forward rates, and that the implied term premia appear sensible (see especially Figure 4 in that paper).

[30] Note that we have dropped the j superscript for the short rate, as it is no longer necessary. See the appendix for more detail on γ_t.

Storage costs can be measured by noting the forward discount on commodities – the difference between log forward and spot rates, which we denote by $\mathbf{cd}_t$ – is equal to the interest rate plus the storage fees: $\mathbf{cd}_t = \mathbf{i}_t + \frac{\mathbf{S}_t}{\mathbf{C}_t}$.

5.1. Results

Table 5 presents our estimates of (11). As above, we estimate (11) both with and without the constraint that $\beta_1 = \beta_2$. That is, we first regress the excess return on the interest rate and storage cost, and then regress it on their sum $\mathbf{cd}_t = \mathbf{i}_t + \frac{\mathbf{S}_t}{\mathbf{C}_t}$.[31] The data are for 3 major metals: lead, nickel, and silver. Spot and 3-month forward prices for these are recorded each month. Unfortunately, the time series are rather short. Although the samples run almost nine years, with a 3-month forecast horizon there are fewer than 36 nonoverlapping observations.

All the estimates of β_1 and β_2 are negative, and a few are statistically significant. This suggests that an increase in the short rate – holding storage costs fixed – is associated on average with lower excess returns. The estimates are also similar in magnitude to what we found in the foreign exchange market, stock, and bond markets above. Although there are no survey data on commodity prices for us to appeal to, the presence of a similar correlation between excess returns and short rates in all of these markets, and the fact that whenever surveys are used these correlations persist, suggest that commodity returns also display excess forecast volatility with respect to the short rate.

[31] With the constraint $\beta_1 = \beta_2$ imposed, the regressions are similar to those reported by Fama and French (1986). Although their selection of commodities differs from ours, their findings are not importantly different for nonseasonal commodities. Seasonal commodities have large predictable components that are clearly associated with movements in storage costs; they tend to yield estimates of β_2 that are near zero.

6. Conclusions

This paper develops simple regression tests capable of distinguishing excessive volatility in expected returns. In doing so we find a striking regularity in excess returns across different asset markets, subsamples and forecast horizons: returns are negatively correlated with short-term interest rates. The size and statistical significance of our estimates of this relation are also comparable across different assets.

In a certain sense, this paper is not the first to document the interest rate's ability to predict excess returns for these markets. Giovannini and Jorion (1987) find direct evidence in the foreign exchange market, while Fama and Schwert (1977) and Keim and Stambaugh (1986), among others, report similar evidence for the stock and bond markets. In addition, there are many studies which find forward discounts to be statistically negatively correlated with excess returns in the foreign exchange, bond, and commodity markets. In the foreign exchange market, the forward discount is equal to the difference between the domestic and foreign interest rates; in the stock market it is equal to the short rate less the dividend yield; in the bond market it is proportional to the difference between the short and long rates; and in commodity markets it is equal to the short rate plus storage costs. Thus, all these results might be interpreted as evidence of the empirical regularity the present paper focuses on.[32]

As always, our results using forward rates alone could in principle be explained by time-varying risk premia. However, we take the additional step of employing survey data on asset return expectations as an independent test of whether the time-varying-risk explanation has merit. Surveys are useful as an alternative measure of expected future returns, one that is free from interference by a risk premium. In virtually every case in which the survey excess returns are used alongside of the actual excess returns, the sign and magnitude of the predictable component of the data remain unchanged. Since the survey excess returns do not include a risk premium, these results suggest that expected returns

[32] The literature in each market documenting the forward discount's ability to forecast returns is far too large to mention here. Representative citations for each market are given in the appropriate sections above.

are excessively volatile with respect to short rate changes.

While it appears that the predictability we document is evidence of systematic in-sample forecast errors, it is important to remember that these errors are not themselves evidence of gross misjudgment by investors. To see that the magnitude of the errors is small, take as an example a \$1000 one-month investment in British pounds (the first regression in Table 1). If the U.S. short rate falls by one percentage point, the results suggest that excess returns rise by 2.488 percent, which implies a one-month unpriced gain of $\frac{\$1000 \times .02488}{12} = \$2.07$ (neglecting compounding). The monthly standard error of the regression, however, is 3.1 percent. Thus, the standard deviation of these extra one-month returns is \$31.00! This is similar to the high risk-return tradeoff implicit in other studies that report predictability of short-horizon excess returns. It may be that there is not a sufficient number of fully-informed, low-transaction-cost traders to fully eliminate the regularity we have identified above.

7. Appendix

The "risk premia" for the assets discussed above are defined as the total required excess returns above the riskless nominal interest rate. As is well-known, as long as unexpected changes in goods prices are correlated with unexpected changes in the asset's price, not all of the required excess return will be attributable to investors' risk aversion. For example, in a world of risk-neutral investors, it is possible for the nominal expected return on an asset to be less than the nominal interest rate. This will occur if those states in which the asset happens to pay off large dollar amounts also turn out to be those in which a dollar purchases more goods.

To see this more formally, consider the following derivation. A necessary condition for the representative investor's intertemporal optimization is the Euler equation:

$$E_t\Big(\frac{m_{t+1}P_t}{P_{t+1}}(1+r^i_{t+1})\Big) = 1, \tag{A1}$$

where $E_t()$ is the expectation operator, applied conditional on all available information at time t, $m_{t+1} = \frac{\delta U'(C_{t+1})}{U'(C_t)}$ is the discounted ratio of marginal utilities of consumption between time $t+1$ and time t, P_t is the time-t dollar price of consumption, and r^i_{t+1} is the nominal return on the ith asset between time $t+1$ and time t. When applied to the riskless rate, r^f_t, condition (A1) for the riskless nominal rate implies that $E_t\left(\frac{m_{t+1}P_t}{P_{t+1}}\right) = 1$. Using this fact, expression (A1), and some simple algebra, we can write the total excess required return on the ith asset as:

$$E_t(r^i_{t+1}) - r^f_t = \frac{\mathrm{cov}_t\left(m_{t+1}, \frac{P_t}{P_{t+1}}(r^i_{t+1} - r^f_t)\right) + E_t\left(m_{t+1}\right) E_t\left(\frac{P_t}{P_{t+1}}(r^i_{t+1} - r^f_t)\right)}{\mathrm{cov}_t\left(m_{t+1}, \frac{P_t}{P_{t+1}}\right) + E_t\left(m_{t+1}\right) E_t\left(\frac{P_t}{P_{t+1}}\right)}, \tag{A2}$$

where $\mathrm{cov}_t()$ is the time-t conditional covariance operator.

We can then split up expression (A2) into two terms: the first due to the correlation between innovations in inflation and asset prices, and the second to risk aversion alone:

$$E_t(r^i_{t+1}) - r^f_t = \frac{E_t\left(\frac{P_t}{P_{t+1}}(r^i_{t+1} - r^f_t - ra^i_t)\right)}{E_t\left(\frac{P_t}{P_{t+1}}\right)} + ra^i_t, \tag{A3}$$

where the term ra_t^i is zero if investors are risk neutral. The first term on the right-hand side of (A3) is derived from (A2) under the assumption that investors are risk neutral.

Let investors' unexpected holding return on the ith asset be given by:

$$r_{t+1}^i - E_t(r_{t+1}^i) = \epsilon_{t+1}^i. \qquad (A4)$$

Then using equations (A3) and (A4) we can express the total realized excess return on asset i as:

$$\begin{aligned} r_{t+1}^i - r_t^f &= \left(r_{t+1}^i - E_t(r_{t+1}^i)\right) + \left(E_t(r_{t+1}^i) - r_t^f\right) \qquad (A5) \\ &= \epsilon_{t+1}^i + \frac{E_t\left(\frac{P_t}{P_{t+1}}(r_{t+1}^i - r_t^f - ra_t^i)\right)}{E_t\left(\frac{P_t}{P_{t+1}}\right)} + ra_t^i, \end{aligned}$$

In the text the latter two terms are referred to as the risk premium. Note however, that the survey unexpected return is a measure of (A4), and therefore, that it contains neither of these latter two terms. In our regressions we find the correlation of (A4) with different regressors to be the same as the correlation of (A5) with those same regressors. Thus, neither of the latter two terms in (A5) is likely to be responsible for the predictability of realized excess returns we document.

8. References

Baillie, R. T., and T. Bollerslev, "Common Stochastic Trends in a System of Exchange Rates," *Journal of Finance,* 44 (March 1989), 167-82.

Bilson, J., "The Speculative Efficiency Hypothesis," *Journal of Business,* 54 (1981), 433-451.

Campbell, J. Y., and R. J. Shiller, "Cointegration and Tests of Present Value Models," *Journal of Political Economy,* 95 (October 1987), 1062-88.

Campbell, J. Y., and R. J. Shiller, "The Dividend-Price Ratio and Expectations of Future Dividends and Discount Factors," *Review of Financial Studies,* 1 (Fall 1988), 195-228.

Campbell, J. Y., and R. J. Shiller, "Yield Spreads and Interest Rate Movements: A Bird's Eye View," Princeton University, 1989.

Cutler, D., J. M. Poterba, and L. H. Summers, "Speculative Dynamics," NBER, November 1989.

Fama, E.F., and K.R. French, "Detecting Spot Price Forecasts in Futures Prices," *Journal of Business* 59 (1986), 39-54.

Fama, E.F., and K.R. French, "Permanent and Temporary Components of Stock Prices," *Journal of Political Economy,* 96 (April 1988), 246-73. (a)

Fama, E.F., and K.R. French, "Dividend Yields and Expected Stock Returns," *Journal of Financial Economics,* 22 (October 1988), 3-27. (b)

Fama, E.F. and G.W. Schwert, "Asset Returns and Inflation," *Journal of Financial Economics,* 5 (March 1977), 319-38.

Ferson, Wayne E., "Changes in Expected Security Returns, Risk, and the Level of Interest Rates," *Journal of Finance,* 44 (December 1989), 1191-1218.

Flood, R.P., R.J. Hodrick, and P. Kaplan, "An Evaluation of Recent Evidence on Stock Market Bubbles," NBER Working Paper No. 1971, July 1986.

Frankel, J.A., and K.A. Froot, "Using Survey Data to Test Standard Propositions Regarding Exchange Rate Expectations," *American Economic Review,* 77 (March 1987), 133-153.

Froot, K.A., "Excess Forecast Volatility in Asset Prices and Expected Returns," Harvard University, 1989. (a)

Froot, K.A., "New Hope for the Expectations Hypothesis of the Term Structure of Interest Rates," *Journal of Finance,* 44 (June 1989), 283-305. (b)

Froot, K.A., and J.A. Frankel, "Forward Discount Bias: Is it an Exchange Risk Premium?," *Quarterly Journal of Economics,* 416 (February 1989) 139-161.

Giovannini, A. and P. Jorion, "Interest Rates and Risk Premia in the Stock Market and in the Foreign Exchange Market," *Journal of International Money and Finance,* 6 (March 1987), 107-24.

Hansen, L.P., "Large Sample Properties of Generalized Method of Moments Estimators," *Econometrica,* Vol. 50, pp. 1029-1054, 1982.

Hodrick, R.J., *The Empirical Evidence on the Efficiency of Forward and Futures Foreign Exchange Markets,* in Fundamentals of Pure and Applied Economics, (Cur, Switzerland: Harwood Academic Publishers), 1988.

Huizinga, J., "An Empirical Investigation of the Long Run Behavior of Real Exchange Rates," University of Chicago, 1986.

Ibbotson Associates, *Stocks, Bonds, Bills and Inflation: 1986 Yearbook,* Ibbotson Associates, Chicago, 1986.

Keim D.B., and R.F. Stambaugh, "Predicting Returns in the Stock and Bond Markets," *Journal of Financial Economics,* 17 (1986), 357-90.

Macaulay, F.R., *Some Theoretical Problems Suggested by Movements of Interest Rates, Bond Yields and Stock Prices in the United States Since 1856,* New York, NBER (1938).

Mankiw, N.G. and L.H. Summers, "Do Long-Term Interest Rates Overreact to Short-Term Interest Rates?," *Brookings Papers on Economic Activity* (1984), 223-242.

Marsh, T., and R. Merton, "Dividend Variability and Variance Bounds Tests for the Rationality of Stock Market Prices," *American Economic Review,* 76 (1986), 483-498.

Newey, W., and K. West, "A Simple, Positive Semi-definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix," *Econometrica,* 55 (May 1987), 703-8.

Poterba, J.M. and L.H. Summers, "The Persistence of Volatility and Stock Market Fluctuations," *American Economic Review,* 76 (1986), 1142-1151.

Poterba, J.M. and L.H. Summers, "Mean Reversion in Stock Prices: Evidence and Implications," *Journal of Financial Economics,* 22 (October 1988) 27-60.

Shiller, R.J., "The Term Structure of Interest Rates," forthcoming in the *Handbook of Monetary Economics,* North-Holland (1988).

Shiller, R.J., J.Y. Campbell, and K.L. Schoenholtz, "Forward Rates and Future Policy: Interpreting the Term Structure of Interest Rates," *Brookings Papers on Economic Activity* (1983), 223-242.

Solnik, B., "The Relation Between Stock Prices and Inflationary Expectations," *Journal of Finance,* 38 (March 1983), 35-48.

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Table 1
Regressions of:
$$\Delta s_{t+j} - fd_t = \alpha + \beta_1 i_t + \beta_2 i_t^* + \eta_{t+j}$$
1-month forecast horizon

Currency	β_1	β_2	F-prob $\beta_i = 0$	DW	$\bar{R}^2$	DF
pound	-2.488^b	3.464^c	0.00	2.04	.08	159
	(0.990)	(1.021)				
French franc	-3.422^c	1.517^a	0.00	2.26	.06	159
	(1.124)	(0.765)				
mark	-3.817^c	1.924^a	0.00	2.26	.05	159
	(1.266)	(1.210)				
yen	-2.941^c	0.175	0.00	1.93	.04	159
	(1.036)	(0.499)				
pound	-2.933^c	2.933^c	0.00	2.05	.08	160
	(0.789)	(0.789)				
French franc	-1.775^b	1.775^b	0.00	2.18	.04	160
	(0.738)	(0.738)				
mark	-2.925^c	2.925^c	0.00	2.18	.04	160
	(1.070)	(1.070)				
yen	−0.670	0.670	0.00	1.85	.01	160
	(0.596)	(0.596)				
pound	−1.564		0.06	1.93	.01	160
	(1.090)					
French franc	-2.400^b		0.00	2.19	.03	160
	(1.094)					
mark	-2.810^c		0.00	2.27	.04	160
	(1.107)					
yen	-2.903^c		0.00	1.93	.05	160
	(1.039)					

Notes are on following page.

Notes: Data are sampled monthly from April 1973 to December 1986, with forecast horizons of 1 month. The middle panel of estimates use $i_t - i_t^*$ as the independent variable, thus imposing the restriction $\beta_1 = -\beta_2$. Interest rates are the average of the bid and ask on 1-month eurodeposits. Standard errors are calculated using GMM allowing for arbitrary serial correlation and heteroskedasticity where appropriate. Intercept terms were included but are not reported. Superscripts ^{a}, ^{b}, and ^{c}, represent statistical significance at the 10, 5 and 1 percent levels, respectively.

Table 2a
Regressions of:
$$\Delta s_{t+j} - \Delta s^e_{t+j} = \alpha + \beta_1 i_t + \beta_2 i^*_t + \beta_3 rp_t + \eta_{t+j}$$
3-month forecast horizon

Dependent Variable	β_1	β_2	β_3	F-prob $\beta_i = 0$	DW	$\bar{R}^2$	DF
$\Delta s_{t+j} - fd_t$	-3.318^b	2.222^a		0.00	1.01	.07	308
	(1.321)	(1.292)					
	-3.255^b	2.130^a	0.194	0.00	1.15	.07	292
	(1.286)	(1.298)	(0.536)				
	-2.916^c	2.916^c		0.00	1.01	.07	309
	(1.036)	(1.036)					
	-2.835^c	2.835^c	0.274	0.00	1.17	.07	293
	(1.016)	(1.016)	(0.525)				
	-2.100^a			0.00	0.95	.05	294
	(1.281)						
	-2.104^a		0.155	0.00	1.08	.05	293
	(1.248)		(0.560)				
$\Delta s_{t+j} - \Delta s^e_{t+j}$	-3.589^b	1.953		0.00	0.91	.08	293
	(1.417)	(1.360)					
	-3.255^b	2.130^a	1.194^b	0.00	1.15	.15	292
	(1.286)	(1.298)	(0.536)				
	-2.992^c	2.992^c		0.00	0.89	.06	294
	(1.095)	(1.095)					
	-2.835^c	2.835^c	1.274^b	0.00	1.17	.14	293
	(1.016)	(1.016)	(0.525)				
	-2.521^a			0.00	0.85	.06	294
	(1.344)						
	-2.104^a		1.155^b	0.00	1.08	.12	293
	(1.248)		(0.560)				

Notes follow Table 2c.

Table 2b
Regressions of:
$$\Delta s_{t+j} - \Delta s^e_{t+j} = \alpha + \beta_1 i_t + \beta_2 i^*_t + \beta_3 rp_t + \eta_{t+j}$$
6-month forecast horizon

Dependent Variable	β_1	β_2	β_3	F-prob $\beta_i = 0$	DW	$\bar{R}^2$	DF
$\Delta s_{t+j} - fd_t$	-3.975^c	2.137^a		0.00	0.56	.20	303
	(1.124)	(1.078)					
	-3.739^c	2.124^a	0.898^b	0.00	0.83	.23	287
	(1.034)	(1.058)	(0.403)				
	-3.586^c	3.586^c		0.00	0.48	.14	289
	(1.083)	(1.083)					
	-3.334^c	3.334^c	1.124^a	0.00	0.85	.19	288
	(0.959)	(0.959)	(0.586)				
	-2.726^c			0.00	0.51	.17	304
	(1.046)						
	-2.515^c		0.834	0.00	0.76	.19	288
	(0.997)		(0.580)				
$\Delta s_{t+j} - \Delta s^s_{t+j}$	-4.100^c	1.905^a		0.00	0.50	.21	288
	(1.236)	(1.150)					
	-3.739^c	2.124^b	1.898^c	0.00	0.83	.31	287
	(1.034)	(1.058)	(0.403)				
	-3.549^c	3.549^c		0.00	0.54	.15	304
	(1.010)	(1.010)					
	-3.334^c	3.334^c	2.124^c	0.00	0.85	.28	288
	(0.959)	(0.959)	(0.586)				
	-2.988^c			0.00	0.47	.18	289
	(1.102)						
	-2.515^c		1.834^c	0.00	0.76	.28	288
	(0.997)		(0.580)				

Notes follow Table 2c.

Table 2c

Regressions of:

$$\Delta s_{t+j} - \Delta s^e_{t+j} = \alpha + \beta_1 i_t + \beta_2 i^*_t + \beta_3 rp_t + \eta_{t+j}$$

12-month forecast horizon

Dependent Variable	β_1	β_2	β_3	F-prob $\beta_i = 0$	DW	$\bar{R}^2$	DF
$\Delta s_{t+j} - fd_t$	-3.160^c	0.702		0.00	0.30	.27	283
	(0.944)	(0.765)					
	-3.179^c	0.664	-0.296	0.00	0.28	.28	272
	(0.929)	(0.799)	(0.638)				
	-3.080^c	3.080^c		0.00	0.25	.14	284
	(0.948)	(0.948)					
	-3.078^c	3.078^c	0.316	0.00	0.37	.14	273
	(0.974)	(0.974)	(0.741)				
	-2.727^c			0.00	0.29	.27	284
	(0.890)						
	-2.785^c		-0.356	0.00	0.35	.27	272
	(0.976)		(0.726)				
$\Delta s_{t+j} - \Delta s^s_{t+j}$	-3.215^c	0.406		0.00	0.32	.30	273
	(0.939)	(0.770)					
	-3.179^c	0.664	0.704	0.00	0.28	.32	272
	(0.929)	(0.799)	(0.638)				
	-3.129^c	3.129^c		0.00	0.29	.13	274
	(0.944)	(0.944)					
	-3.078^c	3.078^c	1.316^a	0.00	0.37	.19	273
	(0.974)	(0.974)	(0.741)				
	-2.967^c			0.00	0.32	.30	274
	(0.878)						
	-2.785^c		0.644	0.00	0.35	.31	272
	(0.976)		(0.726)				

Notes are on following page.

Notes to Tables 2a-2c: The top and bottom panels use, respectively, forward-rate prediction errors and exchange-rate survey prediction errors as dependent variables. Within each panel, the third and fourth sets of estimates use $i_t - i_t^*$ as the first independent variable, thus imposing the restriction that $\beta_1 = -\beta_2$. Interest rates are the average of the bid and ask on 12-month eurodeposits. Data are sampled each 6 weeks from June 1981 to August 1988. Data for 5 currencies (pound, French franc, mark, Swiss franc, and yen) are stacked. Each currency is given its own intercept term (not reported). Standard errors are calculated using GMM allowing for arbitrary serial and cross-sectional correlation, and for heteroskedasticity where appropriate. Superscripts [a], [b], and [c], represent statistical significance at the 10, 5 and 1 percent levels, respectively.

Table 3a

Regressions of:

$$r_{t+j} - r^{e}_{t+j} = \alpha + \beta_1 i_t + \beta_2 \frac{D_t}{P_t} + \eta_{t+j}$$

Data set	β_1	β_2	F-prob $\beta_i = 0$	DW	$\bar{R}^2$	DF
yearly, 1936-85	−1.185	5.188[b]	0.01	1.84	.10	57
	(0.809)	(2.352)				
monthly, 1936-85	−1.364[b]	4.124[b]	0.00	2.00	.02	597
	(0.710)	(2.007)				
monthly, 1956-85	−4.909[b]	17.111[b]	0.00	1.87	.06	357
	(1.064)	(4.190)				
monthly, 1976-85	−5.287[c]	24.474[c]	0.00	1.99	.08	118
	(1.710)	(8.708)				
monthly, 1966-75	−15.385[c]	25.372[b]	0.00	1.96	.12	118
	(3.878)	(12.401)				
monthly, 1956-65	−8.306	1.493	0.15	1.81	.01	118
	(5.061)	(7.654)				
monthly, 1946-55	0.838	5.822	0.08	1.96	.00	118
	(7.784)	(4.091)				
monthly, 1936-45	76.357	2.628	0.08	2.13	.02	118
	(36.141)	(4.999)				

Notes: Stock return data are from CRSP; interest rates are 1-month rates from Ibbotson. Dividend yields are constructed by averaging dividend payments over the previous 12 months. Standard errors are calculated using GMM allowing for arbitrary serial correlation and heteroskedasticity where appropriate. Intercept terms were included but are not reported. Superscripts [a], [b], and [c], represent statistical significance at the 10, 5 and 1 percent levels, respectively.

Table 3b
Regressions of:

$$\mathbf{r}_{t+j} - \mathbf{r}^{e}_{t+j} = \alpha + \beta_1 \mathbf{i}_t + \eta_{t+j}$$

Data set	β_1	F-prob $\beta_i = 0$	DW	$\bar{R}^2$	DF
weekly, 1973-84	-3.975^c (1.287)	0.00	1.99	.01	602
weekly, 1980-86	-6.237^c (2.027)	0.00	1.98	.03	310
yearly, 1936-85	-1.556^b (0.818)	0.03	1.91	.04	57
monthly, 1936-85	-1.631^b (0.709)	0.00	2.01	.01	598
monthly, 1956-85	-1.995^b (0.870)	0.01	1.88	.01	358
monthly, 1976-85	-2.924^c (3.878)	0.00	1.96	.12	119
monthly, 1966-75	-10.347^c (3.488)	0.15	1.92	.06	119
monthly, 1956-65	-8.720^a (4.577)	0.03	1.22	.02	119
monthly, 1946-55	3.370 (8.117)	0.82	1.97	.00	119
monthly, 1936-45	74.856^c (45.512)	0.02	2.15	.02	119

Notes: Stock return data are from CRSP. Interest rates for line 1 are weekly eurodollar deposit rates; those for line 2 are seven-day repurchase agreements. Monthly and annual estimates use 1-month interest rates from Ibbotson. Standard errors are calculated using GMM allowing for arbitrary serial correlation and heteroskedasticity where appropriate. Intercept terms were included but are not reported. Superscripts ^{a}, ^{b}, and ^{c}, represent statistical significance at the 10, 5 and 1 percent levels, respectively.

Table 4a

Regressions of:

$$\mathbf{h}_{t+j}^{(j,k)} - (\mathbf{h}_t^{(j,k)})^e = \alpha + \beta_1 \mathbf{i}_t^{(j)} + \beta_2 \mathbf{i}_t^{(k)} + \beta_3 \theta_t^{(j,k)} + \eta_{t+j}$$

Excess returns on bond buyer index held for 3 months

(k=240 months, j=3 months)

Dependent Variable	β_1	β_2	β_3	F-prob $\beta_i = 0$	DW	$\bar{R}^2$	DF
$\mathbf{h}_{t+j}^{(j,k)}$	−5.141[b]	7.293[a]		0.00	2.35	.12	67
	(2.284)	(2.335)					
	−4.992[c]	7.844[b]	0.282	0.01	2.35	.12	66
	(2.361)	(2.280)	(0.311)				
	−5.252[b]	5.252[b]		0.06	2.42	.10	68
	(2.704)	(2.704)					
	−5.224[b]	5.224[b]	0.070	0.01	2.42	.09	67
	(2.735)	(2.735)	(0.277)				
	−1.132			0.45	2.00	.00	68
	(1.860)						
	−1.082		0.031	0.75	2.00	.00	67
	(2.041)		(0.313)				
$\mathbf{h}_{t+j}^{(j,k)} - (\mathbf{h}_t^{(j,k)})^e$	−5.668[a]	5.344[a]		0.02	2.18	.09	67
	(2.826)	(2.863)					
	−4.992[c]	7.844[b]	1.282[c]	0.00	2.35	.27	66
	(2.361)	(2.280)	(0.311)				
	−5.651[b]	5.651[b]		0.05	2.17	.10	68
	(2.858)	(2.858)					
	−5.224[b]	5.224[b]	1.070[c]	0.00	2.72	.25	67
	(2.735)	(2.735)	(0.277)				
	−2.731			0.02	1.99	.06	68
	(2.068)						
	−1.082		1.031[c]	0.00	2.00	.16	67
	(2.041)		(0.313)				

Notes follow Table 4f.

Table 4b

Regressions of:

$$\mathbf{h}_{t+j}^{(j,k)} - (\mathbf{h}_t^{(j,k)})^e = \alpha + \beta_1 \mathbf{i}_t^{(j)} + \beta_2 \mathbf{i}_t^{(k)} + \beta_3 \theta_t^{(j,k)} + \eta_{t+j}$$

Excess returns on bond buyer index held for 6 months

(k=240 months, j=6 months)

Dependent Variable	β_1	β_2	β_3	F-prob $\beta_i = 0$	DW	$\bar{R}^2$	DF
$\mathbf{h}_{t+j}^{(j,k)}$	-5.955^c	7.906^c		0.00	1.40	.29	66
	(1.674)	(1.325)					
	-5.780^c	8.017^c	0.204	0.00	1.39	.28	65
	(1.701)	(1.357)	(0.363)				
	-6.093^c	6.093^c		0.00	1.46	.26	67
	(1.435)	(1.435)					
	-6.169^c	6.169^c	−0.101	0.00	1.46	.25	66
	(1.549)	(1.549)	(0.361)				
	−1.195			0.29	0.90	.01	67
	(1.391)						
	−1.216		−0.018	0.57	2.00	.00	66
	(1.438)		(0.454)				
$\mathbf{h}_{t+j}^{(j,k)} - (\mathbf{h}_t^{(j,k)})^e$	-6.810^c	7.360^c		0.00	1.36	.27	66
	(1.888)	(1.495)					
	-5.780^c	8.017^c	1.204^c	0.00	1.39	.38	65
	(1.701)	(1.357)	(0.363)				
	-6.850^c	6.850^c		0.00	1.39	.28	67
	(1.530)	(1.530)					
	-6.169^c	6.169^c	0.898^b	0.00	1.46	.35	66
	(1.549)	(1.549)	(0.361)				
	-2.380^a			0.05	0.92	.07	67
	(1.458)						
	−1.216		0.982^b	0.02	2.00	.13	66
	(1.438)		(0.454)				

Notes follow Table 4f.

Table 4c

Regressions of:

$$\mathbf{h}_{t+j}^{(j,k)} - (\mathbf{h}_t^{(j,k)})^e = \alpha + \beta_1 \mathbf{i}_t^{(j)} + \beta_2 \mathbf{i}_t^{(k)} + \beta_3 \theta_t^{(j,k)} + \eta_{t+j}$$

Excess returns on 30 year mortgages held for 3 months

(k=360 months, j=3 months)

Dependent Variable	β_1	β_2	β_3	F-prob $\beta_i = 0$	DW	$\bar{R}^2$	DF
$\mathbf{h}_{t+j}^{(j,k)}$	-6.588^b	8.104^c		0.01	2.79	.08	67
	(2.253)	(2.569)					
	-7.215^b	10.021^c	0.518	0.01	2.78	.10	66
	(2.825)	(3.132)	(0.383)				
	-7.335	7.335		0.13	2.88	.08	68
	(4.795)	(4.795)					
	-7.807^c	7.807^c	0.194	0.01	2.91	.08	67
	(2.831)	(2.831)	(0.327)				
	0.329			0.95	2.49	.00	68
	(2.281)						
	0.258		-0.036	0.99	2.49	.00	67
	(2.737)		(0.499)				
$\mathbf{h}_{t+j}^{(j,k)} - (\mathbf{h}_t^{(j,k)})^e$	-5.378	4.403^a		0.20	2.52	.02	67
	(2.974)	(2.991)					
	-7.215^b	10.021^c	1.518^c	0.00	2.78	.19	66
	(2.825)	(3.132)	(0.383)				
	-4.897	4.897		0.09	2.47	.03	68
	(2.953)	(2.953)					
	-7.807^c	7.807^c	1.194^c	0.00	2.91	.17	67
	(2.831)	(2.831)	(0.327)				
	-1.620			0.37	2.43	.00	67
	(2.203)						
	0.258		0.964^b	0.01	2.49	.08	67
	(2.737)		(0.499)				

Notes follow Table 4f.

Table 4d

Regressions of:

$$\mathbf{h}_{t+j}^{(j,k)} - (\mathbf{h}_t^{(j,k)})^e = \alpha + \beta_1 \mathbf{i}_t^{(j)} + \beta_2 \mathbf{i}_t^{(k)} + \beta_3 \theta_t^{(j,k)} + \eta_{t+j}$$

Excess returns on 30 year mortgages held for 6 months

(k=360 months, j=6 months)

Dependent Variable	β_1	β_2	β_3	F-prob $\beta_i = 0$	DW	$\bar{R}^2$	DF
$\mathbf{h}_{t+j}^{(j,k)}$	-6.652^c	7.382^c		0.00	1.43	.20	66
	(2.973)	(2.893)					
	-6.639^c	8.122^c	0.407	0.00	1.42	.21	65
	(1.972)	(2.041)	(0.351)				
	-7.147^c	7.147^c		0.00	1.50	.20	67
	(2.321)	(2.321)					
	-7.138^c	7.137^c	0.140	0.00	1.51	.20	66
	(2.260)	(2.260)	(0.302)				
	0.119^a			0.98	1.03	.00	67
	(1.282)						
	-0.027		-0.086	0.99	1.03	.00	66
	(1.312)		(0.441)				
$\mathbf{h}_{t+j}^{(j,k)} - (\mathbf{h}_t^{(j,k)})^e$	-6.682^c	5.563^b		0.02	1.33	.12	66
	(2.320)	(2.225)					
	-6.639^c	8.122^c	1.407^c	0.00	1.42	.31	65
	(1.972)	(2.041)	(0.351)				
	-5.924^b	5.924^b		0.02	1.21	.12	67
	(2.425)	(2.425)					
	-7.138^c	7.137^c	1.140^c	0.00	1.51	.29	66
	(2.260)	(2.260)	(0.302)				
	-1.580			0.02	0.19	.02	67
	(1.211)						
	-0.027		0.914^b	0.04	1.03	.10	66
	(1.312)		(0.441)				

Notes follow Table 4f.

Table 4e

Regressions of:

$$\mathbf{h}_{t+j}^{(j,k)} - (\mathbf{h}_t^{(j,k)})^e = \alpha + \beta_1 \mathbf{i}_t^{(j)} + \beta_2 \mathbf{i}_t^{(k)} + \beta_3 \theta_t^{(j,k)} + \eta_{t+j}$$

Excess returns on 12 month Treasury bills held for 3 months

(k=15 months, j=3 months)

Dependent Variable	β_1	β_2	β_3	F-prob $\beta_i = 0$	DW	$\bar{R}^2$	DF
$\mathbf{h}_{t+j}^{(j,k)}$	-1.924^b	2.870^c		0.00	2.24	.27	67
	(0.827)	(0.822)					
	−1.447	2.131^a	−0.346	0.00	2.32	.26	66
	(0.989)	(1.173)	(0.392)				
	−2.481	2.481		0.01	2.33	.08	68
	(1.780)	(1.780)					
	−0.788	0.788	-0.929^c	0.00	2.51	.24	67
	(0.940)	(0.940)	(0.239)				
	0.861^b			0.02	1.85	.15	68
	(0.359)						
	0.271		-0.853^b	0.00	2.25	.24	67
	(0.493)		(0.442)				
$\mathbf{h}_{t+j}^{(j,k)} - (\mathbf{h}_t^{(j,k)})^e$	−0.545	0.736		0.56	2.37	.00	67
	(0.834)	(0.840)					
	−1.447	2.131^a	0.654^a	0.25	2.32	.02	66
	(0.989)	(1.173)	(0.392)				
	−0.658	0.658		0.53	2.49	.00	68
	(0.828)	(0.828)					
	−0.788	0.788	0.071	0.78	2.51	.00	67
	(0.940)	(0.940)	(0.239)				
	0.170			0.56	2.29	.00	68
	(0.329)						
	0.271		0.147	0.73	2.25	.00	67
	(0.493)		(0.442)				

Notes follow Table 4f.

Table 4f

Regressions of:

$$\mathbf{h}_{t+j}^{(j,k)} - (\mathbf{h}_t^{(j,k)})^e = \alpha + \beta_1 \mathbf{i}_t^{(j)} + \beta_2 \mathbf{i}_t^{(k)} + \beta_3 \theta_t^{(j,k)} + \eta_{t+j}$$

Excess returns on 12 month Treasury bills held for 6 months

(k=18 months, j=6 months)

Dependent Variable	β_1	β_2	β_3	F-prob $\beta_i = 0$	DW	$\bar{R}^2$	DF
$\mathbf{h}_{t+j}^{(j,k)}$	-2.069^c	2.642^c		0.00	1.46	.35	66
	(0.733)	(0.699)					
	-1.305	1.461	-0.857^a	0.00	1.59	.37	65
	(0.841)	(0.977)	(0.478)				
	-2.828^c	2.828^c		0.00	1.70	.21	67
	(0.969)	(0.969)					
	-1.164	1.164	-1.083^c	0.00	1.65	.38	66
	(0.706)	(0.706)	(0.251)				
	0.624^c			0.00	1.04	.17	67
	(0.195)						
	-0.107		-1.426^c	0.00	1.50	.36	66
	(0.218)		(0.320)				
$\mathbf{h}_{t+j}^{(j,k)} - (\mathbf{h}_t^{(j,k)})^e$	-1.177^a	1.263^b		0.10	1.60	.05	66
	(0.655)	(0.625)					
	-1.305	1.461	0.143	0.22	1.59	.04	65
	(0.841)	(0.977)	(0.478)				
	-1.291^b	1.291^b		0.09	1.69	.06	67
	(0.581)	(0.581)					
	-1.164	1.164	-0.083	0.09	1.65	.05	66
	(0.706)	(0.706)	(0.251)				
	0.111			0.56	1.34	.00	67
	(0.162)						
	-0.107		-0.426	0.32	1.50	.00	66
	(0.218)		(0.320)				

Notes are on following page.

Notes to Tables 4a-4f: The top and bottom panels use, respectively, forward-rate prediction errors and interest-rate survey prediction errors as dependent variables. Within each panel, the third and fourth sets of estimates use $i_t^{(j)} - i_t^{(k)}$ as the first independent variable, thus imposing the restriction that $\beta_1 = -\beta_2$. Data are sampled each quarter from 1969 to 1986. Intercept terms were included but are not reported. Standard errors are calculated using GMM allowing for arbitrary serial correlation and for heteroskedasticity where appropriate. Superscripts ^{a}, ^{b}, and ^{c}, represent statistical significance at the 10, 5 and 1 percent levels, respectively.

Table 5

Regressions of:

$$c_{t+j} - c^e_{t+j} = \alpha + \beta_1 i_t + \beta_2 \frac{S_t}{C_t} + \eta_{t+j}$$

3-month forecast horizon

Commodity	β_1	β_2	F-prob $\beta_i = 0$	DW	$\bar{R}^2$	DF
lead	−3.164	−0.246	0.04	0.65	.04	100
	(2.008)	(0.395)				
nickel	−6.174	−0.885	0.02	0.37	.04	100
	(4.861)	(0.650)				
silver	−5.858[a]	−4.783[a]	0.05	0.74	.03	95
	(3.477)	(2.912)				
lead	−0.484		0.09	0.67	.01	101
	(0.371)					
nickel	−1.035[a]		0.02	0.37	.03	101
	(0.644)					
silver	−5.284[b]		0.01	0.75	.04	96
	(2.702)					

Notes: Spot and 3-month forward commodity prices are monthly from DRI, 1981 to 1989. Interest rates are the average of the bid and ask on 3-month dollar eurodeposits. Standard errors are calculated using GMM allowing for serial correlation and heteroskedasticity where appropriate. Intercept terms were included but are not reported. Superscripts [a], [b], and [c], represent statistical significance at the 10, 5 and 1 percent levels, respectively.