# THE CROSS-SECTION OF FOREIGN CURRENCY RISK PREMIA AND US CONSUMPTION GROWTH RISK

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**ABSTRACT**

The risk premia produced by the Consumption-CAPM line up with predictable excess returns in currency markets. Consumption growth risk explains why low interest rate currencies do not appreciate as much as the interest rate differential. We take the perspective of a US household that invests in foreign T-bills, and we sort these investments into portfolios based on the nominal interest rate differential with the US. US investors earn low excess returns on low interest rate currencies and high excess returns on high interest rates currencies. We find that US consumption growth risk explains much of this variation in returns across these portfolios because low interest rate currencies provide US investors with a hedge against aggregate consumption risk. This pattern arises because the conditional correlation of foreign consumption growth with US consumption growth decreases in these countries, rendering these currencies riskier for a US investor.

**Keywords**: Exchange Rates, Asset Pricing, Risk Sharing.

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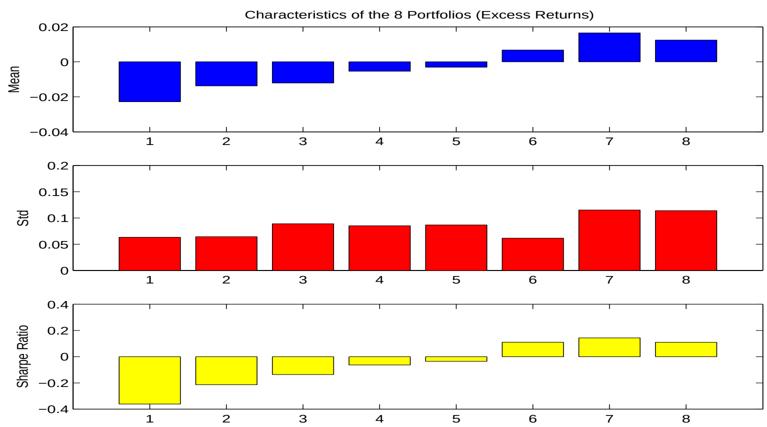
## I. Introduction

This paper applies standard tools in empirical finance to an old question: why do high interest rate currencies tend to appreciate or depreciate less than the interest rate differential on average? To answer that question, we take the perspective of a US household that invests in foreign T-bills, we sort these investments into portfolios based on the nominal interest rate differentials with the US at the end of the previous period, and then explain the variations in average excess returns on these portfolios, not on the currencies themselves. On average, US investors earn low excess returns on low interest rate currencies and high excess returns on high interest rate currencies. The relation is almost monotonic, as shown in figure 1.

**Figure 1.** 8 Currency Portfolios 1953-2002 sorted by current interest rate: Currencies are allocated to portfolios on the basis of the interest rate differential with the US at the end of the previous year. The data is annual.




Why do interest rates predict risk premia on currencies? We find that US consumption growth risk explains much of this variation in forex risk premia, because high interest rate currencies typically depreciate when US consumption growth is low, while low interest rate currencies appreciate. On average, low interest rate currencies hedge US investors against US aggregate consumption growth risk while high interest rate currencies expose them to more consumption growth risk. All our results build on this basic finding. The market price of currency risk is large, but comparable to the one typically found when equity risk is studied. In that sense, the forward premium puzzle looks like a standard asset pricing puzzle. The key variables are the covariances between foreign excess returns and US consumption growth, i.e. consumption growth betas. In addition, we find that ad hoc factors that price bond and equity risk, do not price currency risk, while macroeconomic factor models that exploit conditioning information do.

What gives rise to the monotonic relation between the consumption growth betas of exchange rates and interest rates? There are two possible mechanisms: negative correlation between the first and second moment of the foreign (nominal) stochastic discount factor (SDF), and/or a lower correlation of the SDFs between high interest rate currencies and the US. In a sample of 12 developed countries, we find that the conditional correlation between foreign and US consumption growth decreases with the interest rate gap.

**Currency Portfolios** The foreign currency portfolios built in this paper serve three purposes. First, this method creates a large average spread of up to five hundred basis points between the low and high interest rate portfolios, an order of magnitude larger than the average spread for any two countries. Second, it keeps the number of covariances to be estimated low while allowing us to use data from the largest possible set of countries. Third, it enables us to continuously expand the number of countries studied as additional financial markets open up to international investors.

We consider two large classes of pricing models. The first class uses returns as pricing factors (e.g. Fama & French (1992) and Santos & Veronesi (2001)). The second class introduces measures of macroeconomic undiversifiable risk. They have proven successful in explaining the cross-sectional variation in US stock returns (e.g. Bansal, Dittmar, & Lundblad (2002), Lettau & Ludvigson (2001), Santos & Veronesi (2001), Lustig & VanNieuwerburgh (2002), and Cochrane (2001) for an overview). These macroeconomic models introduce time-variation in the market price of aggregate consumption growth risk by conditioning on other scaling variables.

We test the US investor's Euler equation in two ways. First, we minimize the pricing errors on these currency portfolios. Second, we check the robustness of our results for a smaller set of countries by using the interest rate differential directly as an instrument and checking the Euler equation errors for these country-specific managed portfolios. To save space, we report in this paper results obtained through the first method (GMM) on annual and quarterly data for the periods 1953-2002 and 1971-2002.

In this framework, we show that at annual frequencies, the Consumption Capital Asset Pricing Model (henceforth CCAPM) explains up to eighty percent of the variation in currency risk premia across these eight portfolios. At quarterly frequencies, scaled versions of the CCAPM that introduce additional macroeconomic conditioning information explain up to seventy percent. Thus, the scaled CCAPM explains much of the variation in average excess returns across these portfolios. The estimated coefficient of risk aversion is around 50 for the CCAPM, and the estimated price of aggregate consumption growth risk is mostly positive and significant. Moreover, the price of consumption growth risk in currency markets is not significantly different from that in US equity markets. If we estimate the models only on US domestic stock portfolios sorted by book-to-market and size, we can still explain the variation in currency premia quite well.

**Related Literature** This paper is motivated by three distinct strands of the exchange rate literature. First, interest rate differentials are not unbiased predictors of subsequent exchange rate changes. In fact, high interest rate differentials seem to lead to further appreciations on average (Hansen & Hodrick (1980) and Fama (1984)).[1] Fama (1984) argues that time-varying-risk premia can explain these findings only if (1) risk premia are more volatile than expected future exchange rate changes, and (2) the risk premia are negatively correlated with the size of the expected depreciation. Many authors have concluded that this sets the bar too high, and they have ruled out a risk-based explanation. Froot & Thaler (1990) conclude their survey of this literature as follows:

> A rational efficient markets paradigm provides no satisfactory explanation for the observed results. The conclusion we draw from the tests completed so far is that there is no positive evidence that the forward discount bias is due to risk (as opposed to expectational errors). Risk premia which are derived from economists asset pricing models show no sign of being systematically related to the predictable excess returns derived from econometricians regressions. Taken as a whole, the evidence suggests that explanations which allow for the possibility of market inefficiency should be seriously investigated.

Our work shows currency risk premia predicted by standard asset pricing models do line up with the predictable excess returns in currency markets.

Second, most traditional exchange rate models have proven largely unsuccessful in explaining and/or predicting exchange rates. Meese & Rogoff (1983) conclude that a random walk outperforms most, if not all, of these models in terms of forecasting ability. This is reassuring, because it seems like we are less likely to miss important information in the investor's information set by focusing only on interest rate differentials. [2]

[1] Hodrick (1987), Lewis (1995) and Verdelhan (2004) provide extensive surveys and updated regression results.

[2] In more recent work, Gourinchas & Rey (2003) argue that a measure of current account imbalance predicts returns on US assets held by foreigners, and hence exchange rates.

Third, there is a large, recent literature that tries to explain the volatility and persistence of real exchange rates. This literature builds on earlier work by Lucas (1982) and Stockman (1988). Chari, Kehoe, & McGrattan (2002) rely on price-stickiness to replicate the volatility of the real exchange rate in a general equilibrium model, but they fail to find a link between the ratio of consumption and the real exchange rate, as predicted by the model. This motivates work by Alvarez, Atkeson, & Kehoe (2002): They generate volatile, persistent real exchange rates in a Baumol-Tobin model with endogenously segmented markets, effectively severing the link between the real exchange rate and aggregate consumption growth. Our results suggest that this may be too radical a remedy. Conditional on the interest rate, there appears to be a strong link between consumption growth and exchange rates.

In fact, our results provide empirical support for work by Verdelhan (2003). Verdelhan (2003) replicates the forward discount bias in a model with external habits and provides estimates to support this mechanism. Brandt, Cochrane, & Santa-Clara (2002) point out that the percentage change in the real exchange rate equals the difference between the domestic and the foreign stochastic discount factor when markets are complete. They conclude that real exchange rates are actually less volatile than the size of the Sharpe ratio on equity suggests, assuming there is very little risk sharing between countries. We simply test whether the restrictions imposed by a US investor's Euler equation on the joint conditional distribution of interest rates and nominal exchange rates are satisfied. This Euler equation has to hold regardless of the span of the menu of traded assets. The foreign stochastic discount factor and the foreign price level do not enter into our analysis. We do not have to measure the real exchange rate.

Our paper is closest to recent work by Hollifield & Yaron (2001). They find that real factors drive most of the predictable variation in currency risk premia. In a general class of affine models Backus, Foresi, & Telmer (2001) show that the state variables need to have asymmetric effects on the state prices in different currencies. Our findings are consistent with theirs. Finally, Sarkissian (2003) finds that the cross-sectional variance of consumption growth across countries helps somewhat to explain currency risk premia, but his focus is on explaining unconditional moments of currency risk premium, on a currency-by-currency basis.

The second section outlines our empirical framework, while the third section presents the overall theory behind our estimation strategy. The fourth section presents the asset pricing results obtained on our foreign currency portfolios. The fifth section details the economic mechanism at the core of our results.

## II. Framework

### A. Environment

We focus on a US investor. This investor can trade foreign T-bills. These bills are claims to a unit of foreign currency one period from today in all states of the world. $R_{t+1}^{i,\$}$ denotes the risky dollar return from buying a foreign T-bill in country $i$, selling it after one period and converting the proceeds back into dollars: $R_{t+1}^{i,\$} = R_{t,t+1}^{i,£} \frac{e_{t+1}^i}{e_t^i}$, where $e_t^i$ is the exchange rate in $\$/£$ and $R_{t,t+1}^{i,£}$ is the risk-free one-period return in units of foreign currency $i$. $R_{t,t+1}^{\$}$ is the US currency risk-free rate while $R_{t,t+1}$ is the risk-free rate in units of US consumption.

**Euler equation** We use $m_{t+1}$ to denote the US investor's real stochastic discount factor (henceforth SDF). This discount factor prices payoffs in units of US consumption. In the absence of short-sale constraints or other frictions, the US investor's Euler equation for foreign currency investments holds for each currency $i$:

$$E_t \left[ m_{t+1} R_{t+1}^i \right] = 1, \tag{1}$$

where $R_{t+1}^i$ denotes the return in units of US consumption from investing in T-bills of currency $i$: $R_{t+1}^i = R_{t+1}^{i,\$} \frac{p_t}{p_{t+1}}$, and $p_t$ is the dollar price of a unit of the US consumption basket. The dollar SDF $m_{t+1}^{\$}$ prices dollar returns:

$$E_t \left[ m_{t+1}^{\$} R_{t+1}^{i,\$} \right] = 1, \tag{2}$$

These Euler equations impose testable restrictions on the joint distribution of the US SDF, the exchange rate and interest rates, regardless of the span of traded assets. In the case of complete markets, these restrictions are much tighter.

**Complete Markets** When markets are complete in dollars, the change in the real exchange rate $q_t^i$ equals the ratio of the domestic and the foreign stochastic discount factor:

$$m_{t+1} \frac{q_{t+1}^i}{q_t^i} = m_{t+1}^{i,*}. \tag{3}$$

The Hansen-Jagannathan bounds imply that the standard deviation of $\log m$ is on the order of fifty percent. Large exchange rate changes are implied by large shocks to marginal utility growth at home and abroad, unless the home and foreign SDF are strongly correlated. Brandt et al. (2002) exploit the restrictions imposed by complete markets on the real exchange rates to argue that there is more risk sharing among countries than commonly thought.

**Incomplete Markets** We do not make any assumptions about the span of markets, and we ignore the foreign SDF altogether. Instead we concentrate on the US investor's Euler equation:

$$E_t\left[m_{t+1}\left(R^{i,\$}_{t+1}\frac{p_t}{p_{t+1}} - R^{\$}_{t,t+1}\frac{p_t}{p_{t+1}}\right)\right] = 0. \tag{4}$$

The US investor does not care about the foreign price level and the real exchange rate $q^i_t$ is not relevant here. Instead, the US investor cares about the spread in returns in units of US consumption. This US investor's Euler equation, together with its foreign equivalent, restricts the exchange rate process, but does not uniquely pin it down.

Does the SDF $m_{t+1}$ that prices the returns on US securities also price foreign currency risk? The answer in the literature is *no*, with the exception of Brandt et al. (2002). Our answer is a qualified *yes*: the factors that explain the variation in returns for US assets can explain the variation in returns across different *types of currencies*, i.e. low interest rate vs. high interest rate currencies.

**Currency Portfolios** To analyze the risk-return trade-off for a US investor investing in foreign currency markets, we construct currency portfolios. At the end of each period $t$ we allocate countries to $N^p$ portfolios on the basis of the nominal interest rate differential $R^{i,\$}_{t,t+1} - R^{\$}_{t,t+1}$, observed at the end of period $t$. The low interest rate differential portfolios and high interest rate differential portfolios are ranked from 1 to $N^p$. We compute dollar excess returns of foreign T-bill investments $R^{j,e}_{t+1}$ for each portfolio $j$ by taking (weighted) averages across the different countries in a portfolio. When using annual data, the 3-month T-bill rate was used instead of the one-year rate, simply because fewer countries issue bills at the one year maturity. As data become available, new countries are added to these portfolios. The composition of the portfolio as well as the number of countries in a portfolio obviously changes from one period to the next.

The spread in average excess returns $E_T\left[R^{j,e}_{t+1}\right], j = 1, \ldots, N^p$ across portfolios is much larger than the spread in average excess returns across countries $E_T\left[R^{i,e}_{t+1}\right], i = 1, \ldots, N^c$, simply because the average interest rate differential with the US tend to be rather small for most countries.

**Sample** We always use a total number of eight portfolios. Given the limited number of countries, we do not want too many portfolios. We consider two different samples. First, the longest sample ranges from 1953 to 2002 and spans a number of different exchange rate arrangements. For our purposes, that in itself does not present a problem.[3] Second, we consider a shorter sample ranging from 1971 to 2002. The sample starts with the demise of Bretton-Woods. We use a cutoff value of 20 for Quinn's capital account liberalization index. The estimates for annual data are reported first; the second section discusses the quarterly results, covering a shorter subsample.

**Default** To compute the actual returns on a T-bill investment after default, we used the dataset compiled by Reinhardt, Rogoff, & Savastano (2003) to identify defaults.[4] The (ex ante) recovery rate we applied to T-bills after default is seventy percent. This number reflects two sources: Singh (2003) and Moody's Investors Service (2003). When using quarterly data, we simply assume a country always defaults in the 1st quarter and drop the country from the sample after that.

**Capital Account Liberalization** The restrictions imposed by the Euler equation on the joint distribution of exchange rates and interest rates only make sense if foreign investors can in fact purchase local T-bills. Quinn (1997) has built indices of openness based on the coding of the IMF Annual Report on Exchange Arrangements and Exchange Restrictions. This report covers fifty-six nations from 1950 onwards and 8 more starting in 1954-1960. Quinn (1997)'s capital account liberalization index ranges from zero to one hundred. We chose a cut-off value of 20. This means we eliminate countries where approval of both capital payments and receipts are rare, or when payments or receipts are at best only infrequently granted.

## B. Forward Premium Puzzle

**Uncovered interest rate parity** If investors are risk-neutral, the interest rate differential should be an unbiased predictor of changes in the exchange rate. This means that the slope coefficient $\alpha_1$ in the projection of exchange rate changes on interest rate differentials should be equal to one :

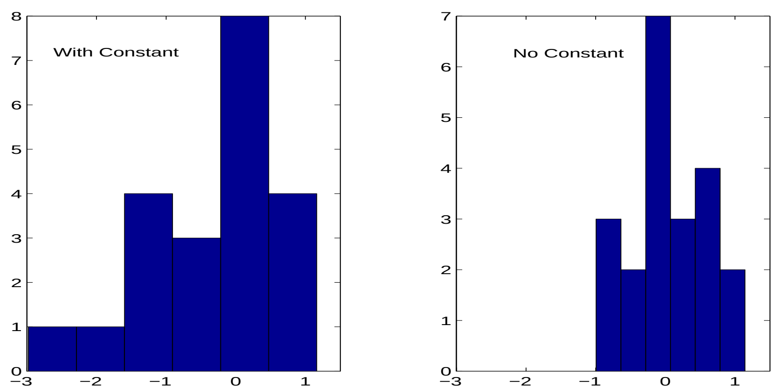
[3]When dealing with fixed exchange rates, investors might allow for a low probability event, and ask for a premium above the one implied by US consumption growth risk.

[4]We would like to thank Carmen Reinhardt for generously sharing those data with us.

$$\Delta \log(e^{-1,i}_{t+1}) = \alpha_0 + \alpha_1 \left[ R^{i,£}_{t+1} - R^{\$}_{t,t+1} \right] + \varepsilon^i_{t+1}, \tag{5}$$

But in the data this slope coefficient $\alpha_1$ is usually not only smaller than one, but more often than not it is negative. Figure 2 plots the histogram of slope coefficients for 16 currencies between 1971.1 and 2002.4, with and without a constant. All of the slope coefficients are smaller than one, but most of the slope coefficients are negative. If we do not include a constant in the regression, the distribution of the slope coefficients shifts to the right and the mean of the distribution is close to zero. The first picture indicates that currencies with *unusually* low interest rates depreciate while the second picture indicates that low interest rate currencies do not appreciate as much as the interest rate differential. This suggests investors must have known they can *make money* by chasing high interest rates.

**Figure 2.** Histogram of UIP slope coefficients: OLS Regression of exchange rate changes at quarterly frequency on interest rate differentials for 16 countries. The sample is 1971.1-2002.4 and the frequency is quarterly. The left panel shows the results for the regression that includes a constant, the right panel excludes the constant. The sample includes Australia, Belgium, Barbados, Canada, Germany,France, the United Kingdom, Ireland, Italy, Jamaica, Japan, Malaysia, the Netherlands, Sweden, Trinidad and Tobago, Vietnam




**Forex Excess Returns** What does the failure of U.I.P. imply for the excess returns on these currency portfolios? Table I lists the mean excess return, the standard deviation and the Sharpe ratio for foreign T-bill investments. The largest spread exceeds five percentage points for the 1971-2002 subsample. The average annual returns are almost monotonic in the interest rate differential, except for the last portfolio. This last one is comprised of high inflation currencies; UIP tends to work much better at high inflation levels. This non-monotonicity has been documented extensively by Bansal & Dahlquist (2000). We will re-visit this issue in the next section. Most surprising are the negative Sharpe ratios of up to minus forty percent for the lowest interest rate currency portfolios.

The same pattern is found for developed countries (see Table X in the Appendix on the authors' web sites), although the spread in excess returns drops from five to two and a half percentage points. Especially at quarterly frequency, the non-monotonicity in excess returns has all but disappeared, confirming the findings of Bansal & Dahlquist (2000): the forward premium puzzle all but disappears for developed countries, at high levels of the nominal interest rate.

**Table I**
**Statistics for 8 Currency Portfolios**

Reports the mean, standard deviation and the Sharpe ratio for the real excess return on investing in foreign T-Bills for each of the eight portfolios $i$. These portfolios are constructed by ranking currencies into portfolios at time $t$ based on the nominal interest rate differential with the US at the end of period $t-1$. Portfolio 1 is the portfolios of currencies with the smallest interest rate differential. In Panel A, the frequency of the data is annual. In Panel B, (annualized) quarterly results are reported. The sample includes all countries in a given year which are assigned a Quinn capital account liberalization index that exceeds 20.

| | 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 |
|---|---|---|---|---|---|---|---|---|
| **Panel A: Annual Returns** | | | | | | | | |
| *1953-2002* | | | | | | | | |
| *Mean* | −0.023 | −0.014 | −0.012 | −0.0053 | −0.0031 | 0.0067 | 0.016 | 0.012 |
| *Std.* | 0.063 | 0.064 | 0.089 | 0.085 | 0.087 | 0.061 | 0.12 | 0.11 |
| *Sharpe Ratio* | −0.36 | −0.21 | −0.14 | −0.063 | −0.035 | 0.11 | 0.14 | 0.11 |
| *1971-2002* | | | | | | | | |
| *Mean* | −0.029 | −0.0076 | −0.0037 | −0.0015 | −0.0091 | 0.012 | 0.022 | 0.0042 |
| *Std.* | 0.078 | 0.066 | 0.088 | 0.1 | 0.11 | 0.075 | 0.14 | 0.14 |
| *Sharpe Ratio* | −0.37 | −0.12 | −0.042 | −0.014 | −0.084 | 0.16 | 0.16 | 0.031 |
| **Panel B: Quarterly Returns** | | | | | | | | |
| *1953.1-2002.4* | | | | | | | | |
| *Mean* | −0.028 | −0.004 | −0.014 | 0.0092 | −0.0023 | 0.00011 | 0.025 | 0.0055 |
| *Std.* | 0.13 | 0.094 | 0.16 | 0.14 | 0.14 | 0.12 | 0.13 | 0.16 |
| *Sharpe Ratio* | −0.22 | −0.042 | −0.09 | 0.065 | −0.017 | 0.00092 | 0.19 | 0.034 |
| *1971.1-2002.4* | | | | | | | | |
| *Mean* | −0.029 | −0.0029 | −0.0036 | 0.016 | −0.006 | −0.0033 | 0.035 | −0.0032 |
| *Std.* | 0.13 | 0.12 | 0.14 | 0.17 | 0.17 | 0.15 | 0.15 | 0.2 |
| *Sharpe Ratio* | −0.22 | −0.025 | −0.025 | 0.095 | −0.036 | −0.022 | 0.22 | −0.016 |

**Slope coefficients in Cross-Section** At first sight, U.I.P is not a bad fit for the cross-section (results are reported in the Appendix). The slope of the regression line is .85 if we include the eight portfolios. But the estimated slope coefficient drops to .4 if we exclude the highest interest rate portfolios. We know that U.I.P starts to work better for developing countries with high interest rates and this explains the pattern.[5]

It appears that the simple investment strategy our currency portfolios analyze produce very large spreads in risk premia of up to five percent. The next section first introduces the class of linear factor models we focus on and then it explores the relation between the conditional risk premia and interest rates.

## III. Do we Need a new Theory for Currency Risk?

We argue that forward premia are fully consistent with standard asset pricing theory. Our toolbox includes linear factor models that have proved useful in pricing currency risk. This section presents these models and show why some might account for the forward premia.

### A. Linear Factor Models with Time-Varying Coefficients

Our objective is to link currency risk premia to standard asset pricing factors in a linear pricing framework:

$$m_{t+1} = c_t + d_t' F_{t+1}, \tag{6}$$

where $c_t$ and $d_t$ can depend on the vector of scaling variables $x_t$. Time-variation in risk premia plays a key role for currency risk premia and the scaling variables will help us to capture this. We restrict this relation to be linear: $c_t = \gamma_0 + \gamma_1 x_t$ and $d_t = \eta_0 + \eta_1 x_t$. We consider two large classes of pricing models. The first class uses returns as pricing factors. In this group are the Fama & French (1992) and the Santos & Veronesi (2001)' models. The second class of models directly introduces measures of the undiversifiable, macroeconomic risk that investors are compensated for. The Consumption-CAPM (henceforth CCAPM) and its scaled versions belongs to this class. Table II summarizes these two classes of linear factor models.

**Return factors** First, we consider the Fama-French equity pricing factors: the CRSP value-weighted excess return $R^{vw}$, the small-minus-big return $R^{SB}$ and the high-minus-low return $R^{HL}$. These factors explain the variation in returns along the book-to-market and size dimensions relatively well (Fama & French (1992)). We refer to this model as the $FF$-CAPM for equity. Second, Fama & French (1992) also construct two bond pricing factors; the first one is the difference between the long term government bond return and the risk free rate $R^{long}$ and the

[5]The estimated slope coefficient drops to .2 if we only consider developed countries over the same sample! So, the U.I.P appears to work well in cross-section, only because of developing countries with high interest rate differentials.

**Table II**

**Linear Factor Models: The upper panel contains models with returns as factors; the lower panel contains consumption-based models**

| | $f_1$ | $f_2$ | $f_3$ | $f_4$ | $f_5$ |
|---|---|---|---|---|---|
| *FF-CAPM equity* | $R^{vw}$ | $R^{SB}$ | $R^{HL}$ | | |
| *FF-CAPM bonds* | | | | $R^{long}$ | $R^{corp}$ |
| *FF-CAPM bonds+equity* | $R^{vw}$ | $R^{SB}$ | $R^{HL}$ | $R^{long}$ | $R^{corp}$ |
| *y-CAPM* | $R^{vw}$ | $\frac{l}{c}$ | $R^{vw}\frac{l}{c}$ | | |
| *CCAPM* | $\Delta(\log c_t)$ | | | | |
| *HCAPM* | $\Delta(\log c_t)$ | | | $A_{t-1}\Delta(\log \rho_t)$ | |
| *cay-CCAPM* | $\Delta(\log c_t)$ | $\Delta(\log c_t)cay_{t-1}$ | | | |
| *my-CCAPM* | $\Delta(\log c_t)$ | $\Delta(\log c_t)my_{t-1}$ | | | |
| *my-HCAPM* | $\Delta(\log c_t)$ | $\Delta(\log c_t)my_{t-1}$ | $A_{t-1}\Delta(\log \rho_t)$ | $A_{t-1}\Delta(\log \rho_t)my_{t-1}$ | |

second one is the spread between the return on a long-term corporate bond index and a long term government bond $R^{corp}$. Fama & French (1993) argue that these factors proxy for the underlying undiversifiable macroeconomic risk. Third, Santos & Veronesi (2001) propose a scaled version of the standard CAPM; the scaling variable is the labor income share. An increase in the labor income share reduces the stand-in investor's exposure to equity risk and this reduces the market price of risk. Santos & Veronesi (2001) show this conditional version of the CAPM explains a large share of the cross-sectional variation in average returns. We refer to this model as the $y$-CAPM.

$$m_{t+1} = b_0 + b_1 \Delta \log \left(R^m_{t+1}\right) + b_2 x_t \Delta \log \left(R^m_{t+1}\right)$$

**Scaled Consumption-CAPM** We consider the standard Consumption-CAPM with only aggregate consumption growth risk and the Housing-CAPM (see Piazzesi, Schneider, & Tuzel (2002)), henceforth HCAPM, which introduces rental price growth risk in addition to aggregate consumption growth risk. Finally, we consider two different scaled versions of the consumption CAPM. To allow for time-variation we follow Lettau & Ludvigson (2001) in proposing a linearized version of the standard Breeden-Lucas stochastic discount factor:

$$m_{t+1} = b_0 + b_1 \Delta \log \left(c_{t+1}\right) + b_2 x_t \Delta \log \left(c_{t+1}\right) \tag{7}$$

Two scaling factors $x_t$ are considered. Lettau & Ludvigson (2001) introduce the consumption wealth ratio ($cay$) as a scaling variable to capture the variation in the conditional moments that the standard CCAPM cannot deliver. Although it is not explicitly derived as such, Lettau & Ludvigson (2001) motivate this scaling by appealing to habit formation. Campbell & Cochrane (2000) argue that scaled models will outperform the CCAPM in the case of habit formation.

Lettau & Ludvigson (2001) choose the consumption-wealth ratio because it summarizes the agent's expectations about future returns in a wide class of models. This becomes apparent when one loglinearizes the budget constraint. We refer to this model as the *cay*-CCAPM.

Lustig & VanNieuwerburgh (2002) derive (7) in an economy with heterogenous agents in which the housing collateral ratio $my$ governs the amount of risk sharing. When the housing collateral ratio is low, it is harder for households to share idiosyncratic risk. This increases the market price of aggregate consumption growth risk. In our empirical work we rescale ($my$) to keep it positive as follows: $my_{max} - my$; the scaling variable is an indicator of collateral scarcity.[6] Lustig & VanNieuwerburgh (2002) explain how the ratio of collateralizable wealth is measured empirically as the residual from a cointegrating relationship between labor income and housing wealth, along the lines of the computation by Lettau & Ludvigson (2001) for the consumption-wealth ratio. We refer to this model as the *my*-CCAPM, or the *my*-HCAPM, if we allow for non-separabilities.

The relative success of the models proposed by Santos & Veronesi (2001), Lettau & Ludvigson (2001) and Lustig & VanNieuwerburgh (2002) in pricing domestic stock returns suggest that the Fama-French asset pricing factors do proxy for underlying macroeconomic risk. We will show that the macroeconomic factor models can price both currency risk and domestic equity risk, while the Fama-French factors cannot.

**Unconditional Factor Model** As explained by Lettau & Ludvigson (2001), the conditional factor model maps into unconditional factor model where the unconditional factor $f^u$ is given by:

$$F_t^u = [f_t; x_{t-1} * F_t]$$

This unconditional factor model is the one we will estimate. In principle, one could allow the constant $c_t$ in the SDF in equation (6) to be time-varying. This would introduce the scaling variable itself as a pricing factor, but we decided to only include the interaction terms, since there is no compelling economic reason to expect scaling variable risk to be priced.

This means the US stochastic discount factor is affine in $F_{t+1}$, a $K \times 1$ vector of pricing factors:

$$m_{t+1} = b' F_{t+1}^u$$

[6] We use three different measures of housing collateral, one based on residential wealth $myrw$, one based on fixed assets $myfa$, and, finally, one based on outstanding mortgages, $mymo$.

## B. Theoretical Consumption Growth Betas

Risk premia are time-varying and they increase on average with the interest rates. To explain these pattern in the data, the conditional currency risk premia need to switch signs, depending on whether the currency has low or high nominal interest rates, to explain the negative slope coefficients. We work out the log-normal case to illustrate this theoretical mechanism.

**Log-normality** Assume $\log m_{t+1}$ and $\log R^i_{t+1}$ are jointly, conditionally normal. Then the Euler equation can be restated to obtain an expression for the log of the real currency risk premium:

$$\log E_t R^i_{t+1} - \log R_{t,t+1} = -Cov_t \left( \log m_{t+1}, \log R^{i,\$}_{t+1} - \Delta \log p_{t+1} \right).$$

We refer to the log premium as $crp^i_t$. It is determined by the covariance between the log of the SDF $m$ and the real returns on investing in foreign T-bills.

Assume that the log of the real SDF is linear in the log of the pricing factors $F$:

$$\log m_{t+1} = b_0 + \sum_{j=1}^{n} b_j(x_t) \log F_{j,t+1},$$

where the coefficients $b$ depend on time $t$ information variables $x_t$. Substituting in for the factors, we can restate the log currency risk premium as:

$$log(crp^i_{t+1}) = -\sum_{j=1} b_j(x_t)[Cov_t \left( \log F_{j,t+1}, \Delta \log e^i_{t+1} \right) + Cov_t \left( \log F_{j,t+1}, \Delta \log p_{t+1} \right)]$$

The first term is pure currency risk compensation. The second part is inflation risk compensation.

Our benchmark model is the scaled CCAPM. We examine what restrictions are implied on the joint distribution of consumption growth and exchange rates by this increasing pattern of currency risk premia in interest rates.

**Consumption Growth and Exchange Rates** In the scaled CCAPM, the single pricing factor is consumption growth $F_{t+1} = \frac{c_{t+1}}{c_t}$ and we assume the parameter $b(x)$ is affine in $x_t$ with $b_2 > 0$ and $x_t > 0$. If there is a stand-in agent with CRRA preferences who consumes aggregate US consumption, $b_1 = \gamma$ and $b_2 = 0$. This is the standard CCAPM. But, in general, the scaling variable $x$ can introduce time-variation in the market price of consumption growth risk.

It follows that the log currency risk premium is determined by the sign of the conditional covariance between consumption growth and the change in the exchange rate (usually known in the finance literature as consumption growth betas):

$$\log\left(crp^i_{t+1}\right) = -\left(b_1 + b_2 x_t\right)\left[Cov_t\left(\Delta(\log c_{t+1}), \Delta \log e^i_{t+1}\right) - Cov_t\left(\Delta(\log c_{t+1}), \Delta \log p_{t+1}\right)\right]$$

We can abstract from the inflation compensation term $\Delta \log(p_t)$, because it cannot explain any of the cross-sectional variation. This equation conveys two essential insights about what is needed to explain the forward premium/discount[7]:

1. the consumption growth betas of currencies need to be positive when foreign interest rates are low and negative when interest rates are high;

2. the size of the risk premia increases when high interest rate currencies are more sensitive to US consumption growth in bad times, when $x$ is large.

First, currencies that are expected to appreciate when US consumption growth is high and depreciate when US consumption growth is low, earn a positive conditional risk premium. In the data, the risk premium $\left(crp^i_{t+1}\right)$ is positively correlated with foreign interest rates $R^{i,£}_{t,t+1}$: low interest rate currencies earn negative risk premia and high interest rate currencies earn positive risk premia. To match this fact, the following necessary condition needs to be satisfied:

$$Cov_t\left(\Delta \log c_{t+1}, \Delta \log e^i_{t+1}\right) > 0 \textit{ when } R^{i,£}_{t,t+1} \textit{ is low}$$
$$Cov_t\left(\Delta \log c_{t+1}, \Delta \log e^i_{t+1}\right) < 0 \textit{ when } R^{i,£}_{t,t+1} \textit{ is high}$$

Moreover, the $Cov_t(\cdot,\cdot)$ needs to switch signs over time for a given currency. The next section explains how this pattern arises when the conditional volatility of the (nominal) SDF increases when (nominal) interest rates are low.

Second, if high interest rate currencies are more correlated with aggregate US consumption growth in bad times, when $x_t$ is large, this increases the size of the risk premium.

## C. Preliminary empirical betas

We provide some empirical evidence to support both of the claims above. First, we examine the consumption growth betas of exchange rates in each of these portfolios. Second, we examine the

[7]We are grateful to Andy Atkeson for this clarification.

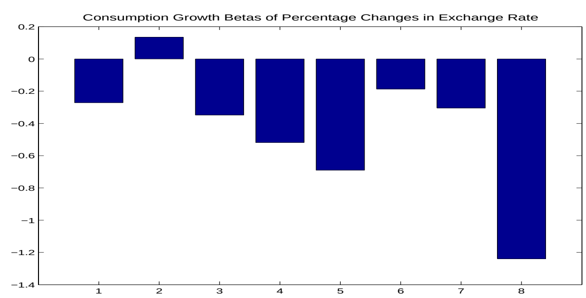
conditional consumption growth betas.

**Unconditional Consumption Growth Betas of Foreign Currencies** To check whether this necessary condition is satisfied in the data, we compute the unconditional consumption growth betas by regressing the deflated average change in the exchange rate on US consumption growth, for each of 8 currency portfolios:

$$\Delta(\log e^i_{t+1}) - \Delta \log p_{t+1} = \alpha_0 + \alpha_1 \Delta(\log c^{US}_{t+1}) + \epsilon_{t+1}$$

For the quarterly data, Figure 3 shows positive or small negative betas for low interest rate currencies and large negative betas for high interest currencies. On average, low interest rate currencies hedge US investors against aggregate consumption growth risk while high interest rate currencies expose them to more consumption growth risk. In Figure 4, the annual data show a similar pattern, at least for the post Bretton-Woods sample. All our results build on this basic finding. The returns on foreign cash holdings in high interest rate countries expose US investors to more consumption growth risk, while the returns on foreign cash holdings in low interest rate currencies provide a hedge.

**Figure 3.** Consumption Growth Betas of Exchange Rates Estimated slope coefficients in regression of percentage exchange rate changes on US consumption growth for 8 currency portfolios: $\Delta \log(e^{-1,i}_{t+1}) - \Delta \log(p_{t+1}) = \beta_0 + \beta_1 \left[\Delta \log c^{US}_{t+1}\right]$. We use the 1971.1-2002.4 sample. Quarterly Data.




**Conditional Consumption Growth Betas of Foreign Currencies** To evaluate the conditional consumption growth betas, we run the following regression of exchange rates on US consumption growth and US consumption growth interacted with the scaling variable:

$$\Delta(\log e^i_{t+1}) - \Delta \log p_{t+1} = \alpha_0 + \alpha_1 \Delta(\log c^{US}_{t+1}) + \alpha_2 x_t \Delta(\log c^{US}_{t+1}) + \epsilon_{t+1}$$

**Figure 4.** Consumption Growth Betas of Exchange Rates Estimated slope coefficients in regression of percentage exchange rate changes on US consumption growth for 8 currency portfolios: $\Delta \log(e_{t+1}^{-1,i}) - \Delta \log(p_{t+1}) = \beta_0 + \beta_1 \left[\Delta \log c_{t+1}^{US}\right]$. We used the 1971-2002 sample. Annual Data.

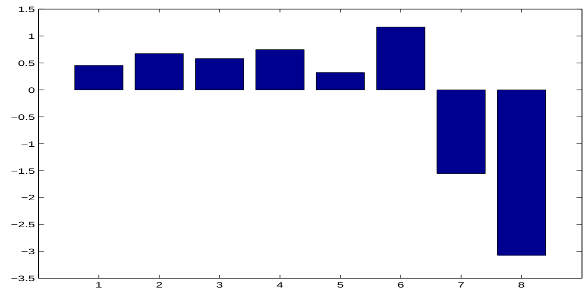
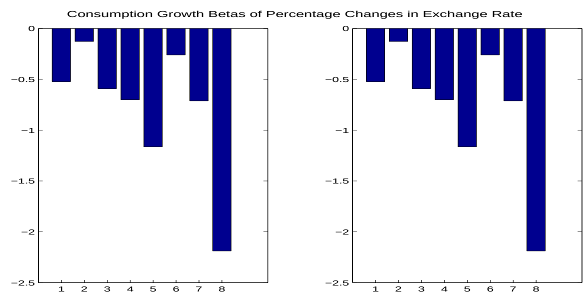


**Figure 5.** Consumption Growth Betas of Exchange Rates in Bad Times Estimates from regression of percentage exchange rate changes on consumption growth using 8 currency portfolios as test assets. The left panel uses the 1953.1-2002.1 sample. The right panel uses the 1971.1-2002.4 sample. Quarterly Data.




The sensitivity of high interest rate currencies increases in bad times, when $x$ is large. We consider the Coll-CCAPM: the housing collateral ratio $my$ is the scaling variable. Figure 5 plots the same consumption growth betas but evaluated at $my$ one half standard deviation above its sample mean. The consumption growth betas for high interest rate currencies nearly double relative to the unconditional values, when housing collateral is scarce in the US. High interest rate currencies become much riskier in bad times!

# IV. Estimation

We use three different estimation strategies, the first relies on a G.M.M estimator, the second is a two-step linear regression procedure due to Fama & MacBeth (1973) commonly used in cross-sectional asset pricing, the third is equivalent to the pricing of managed portfolios.

**GMM** SDF is $m_{t+1} = 1 - b'F_{t+1}$. The moments conditions are the sample analog of the population pricing errors:

$$g_T(b) = E_T(m_t R_t^e) = E_T(R_t^e) - E_T(R_t^e F_t^{u,\prime})b$$

where $R_t^e = [R_t^{1,e}\ R_t^{2,e}\ ..\ R_t^{Np,e}]'$ in the first stage of GMM we use the identity matrix as the weighting matrix, $W = I$, while in the second stage we use $W = S^{-1}$ where $S$ is the covariance matrix of the pricing errors in the first stage: $S = \sum_{-\infty}^{\infty} E[(m_t R_t^e)(m_{t-j} R_{t-j}^e)']$. Since we focus on linear factor models, the first stage is equivalent to an OLS-cross-sectional regression of average returns on the second moment of returns and factors, while the second stage is a GLS cross-sectional regression of average excess returns on the second moment of returns and factors ( see e.g. Cochrane (2001), chapter 13).

Essentially, we gauge how much of the variation in average returns across portfolios can be explained by variation in the betas. If the model is successful, this means that we can claim success in explaining exchange rate changes, conditional on whether the country is a low or high interest rate currency.

## A. Currency Portfolios as Test Assets

When it comes to pricing currency risk, the factors that directly measure macroeconomic, undiversifiable risk outperform the factors constructed only from asset returns. In a first step, we only use the currency portfolios as test assets. In the next section we introduce additional test assets.

**CCAPM** The standard CCAPM can explain between sixty percent for 1971-2003 and eighty percent between 1953 and 2002 of the cross-sectional variation in average excess returns US investors earn on our constructed currency portfolios. Figure 6 plots the actual sample average of the excess return $E_T\left[R_{t+1}^{j,e}\right]$ on the vertical axis against the predicted excess return $\beta_j'\lambda$ on the horizontal axis, for each of the eight currency portfolios $j$; the right panel of the figure plots the CCAPM results with predicted excess return $\beta_c^j \lambda_c$, the panel on the left plots the CAPM results, with predicted excess return $\beta_R^j \lambda_R$. Variation in market betas hardly explains any of the variation in returns, while the variation in consumption betas explain eighty percent.

**Figure 6.** CAPM and CCAPM: Predicted vs. Actual Excess Return for 8 Currency Portfolios between 1953-2002. Predicted excess return on horizontal axis. GMM estimates using 8 currency portfolios as test assets. Annual Data.

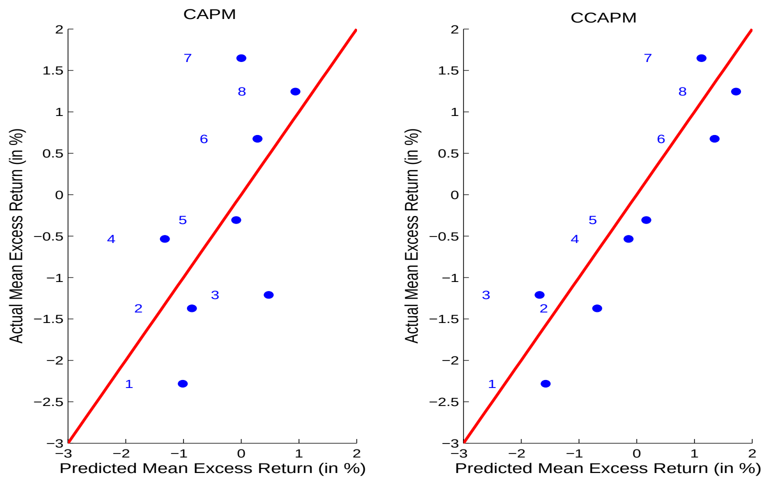


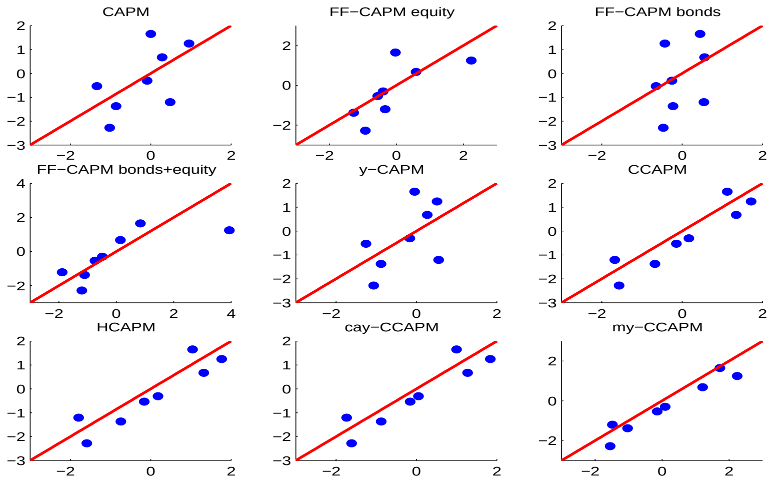

Table III reports the estimated market prices of risk and the p-value for the $\chi^2$-test. The estimated price of consumption growth risk $\lambda_c$ is positive, but not significant in all cases. $\lambda_c$ is large, around five. An asset with a consumption growth beta of one yields an average risk premium of five percent.

The bottom panel of Table III reports estimates using quarterly returns on eight currency portfolios that are re-balanced each quarter instead of each year. The results confirm our findings for the annual returns. In the quarterly data, we observe a similar pattern, but now the standard CCAPM explains only fourty percent of the variation in returns, but the HCAPM and the *my*-HCAPM explain up to seventy percent. As before, the price of consumption growth risk $\lambda_c$ in the CCAPM is large; the US investor earns a quarterly excess return of 2.5 percent on an asset with a consumption growth beta of one. The price of scaled consumption growth risk $\lambda_{c,x}$ is positive but not significant.

To assess whether individual factors have explanatory power, we also report the coefficient estimates $b$ in a separate Appendix in Table XI. The implied coefficient of risk aversion in the CCAPM is around 56! This is line with estimates of the coefficient of risk aversion the literature, but those are based on stocks.

Finally, Figure 7 plots the predicted against actual excess return for 4 factor models and 5 consumption-based models. The single-factor CCAPM clearly does as well or better than the

**Figure 7.** Predicted vs. Actual Excess Return for 8 Currency Portfolios between 1953-2002. Predicted excess return on horizontal axis. GMM estimates using 8 currency portfolios as test assets. Annual Data. Each panel plots the results for one of the 9 linear factor models. The filled dots are the currency portfolios.



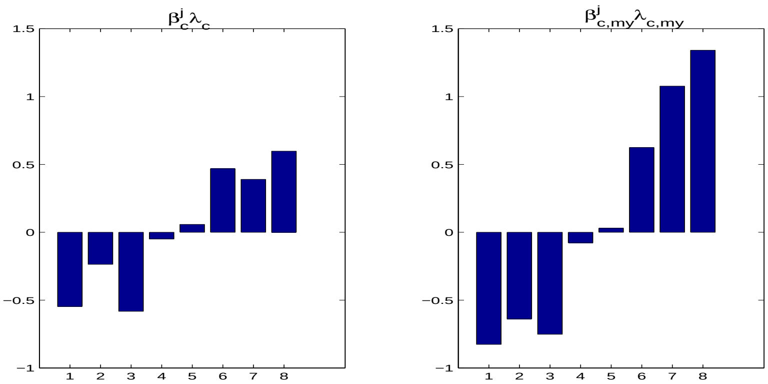

multi-factor models without consumption growth. This is discussed in some detail in the next susbsection.

**Scaled CCAPM** The scaled versions of the CCAPM capture quite some of the variation in currency risk premia, especially at quarterly frequencies, because (1) the consumption growth betas of exchange rates switch signs between high and low interest rate episodes and (2) these betas increase in absolute value when the scaling variable is large, i.e. in bad times. Recall that the expected return on currency portfolio $j$ predicted by the model consists of two parts:

$$E[R^{e,j}] = \beta_c^j \lambda_c + \beta_{c,x}^j \lambda_{c,x}$$

The first part is the consumption growth risk premium; the second part is the risk premium for consumption growth risk in bad times. The estimated price of scaled consumption growth risk $\lambda_{c,x}$ is positive as well, as predicted by the theory. This means the price of consumption growth risk increases in bad times, when $x$ is large. Collateral-consumption growth risk plays a role in explaining currency risk. Figure 8 depicts the consumption growth risk premium and the consumption-growth-collateral risk premium for each of the eight currency portfolios. For low interest rate currencies, there is -.60 percentage points due to consumption growth risk and about -.75 percentage points due to consumption-growth-collateral risk. For high interest rate currencies, there is .5 percentage points due to consumption growth risk and about 1.5 percentage points due to consumption-growth-collateral risk.

**Figure 8.** *my*-CCAPM: Risk Premia 1953-2002. GMM estimates using 8 currency portfolios as test assets. Annual Data.




The coefficient estimates reported in table XI in the Appendix reveal whether individual factors have explanatory power for currency risk premia, rather than wether the risk is priced. The estimated coefficients $b_{c,x}$ for the interaction term with the scaling variable are positive and significant for the *my*-CCAPM and *my*-HCAPM, but not for the *cay*-CCAPM.

**CAPM** On the other hand, the basic CAPM explains only 36 percent of the variation in annual excess returns, compared to eighty percent over the same sample for the CCAPM. Adding other return-based factor does not help much. The Fama-French equity factors explain only a small fraction of the variation in average excess returns across these portfolios. The estimated price of market risk is large: an asset with a beta of one earns an annual excess return of between 9 and 16 percent.

Using quarterly re-balanced portfolios, the standard CAPM hardly explains any of the variation in quarterly currency returns, but the Fama-French bond factors do. These explain up to eighty percent.

In addition, these ad hoc factor models will have trouble pricing both currency and bond/equity risk, because some of the estimated risk prices have the wrong sign, as reported in Table IV.

## Table III
## CCAPM Risk Price Estimates for Currency Portfolios

GMM estimates using 8 currency portfolios as test assets. The first column contains $\lambda_c$, the second column $\lambda_x$, the third column $\lambda_{c,x}$, the fourth column $\lambda_\rho$ and the final column $\lambda_{\rho,x}$, where $x$ denotes the scaling factor. The upper panel reports estimates based on annual data, the lower reports estimates based on quarterly data. The collateral measure for the *my*-CCAPM and *my*-HCAPM is $myfa$. We used the optimal lag length to estimate the spectral density matrix (Andrews, 1991).

| *Model* | $\lambda_c$ | $\lambda_{c,x}$ | $\lambda_\rho$ | $\lambda_{\rho,x}$ | *p value* | $R^2_{adj}$ | $R^2$ |
|---|---|---|---|---|---|---|---|
| **Panel A: Annual Data** | | | | | | | |
| *1953-2002* | | | | | | | |
| *CCAPM* | 5.3 | | | | 0.77 | 0.81 | 0.81 |
| *s.e.* | (2.8) | | | | | | |
| *HCAPM* | 5.3 | | 0.46 | | 0.79 | 0.76 | 0.8 |
| *s.e.* | (2.9) | | (0.31) | | | | |
| *cay-CCAPM* | 5.3 | 1.1 | | | 0.61 | 0.79 | 0.82 |
| *s.e.* | (3.7) | (0.8) | | | | | |
| *my-CCAPM* | 7.7 | 1.2 | | | 0.74 | 0.79 | 0.82 |
| *s.e.* | (7.4) | (1.1) | | | | | |
| *my-HCAPM* | 11 | 1.7 | 0.63 | 0.075 | 0.78 | 0.78 | 0.87 |
| *s.e.* | (26) | (4) | (1.9) | (0.22) | | | |
| *1971-2002* | | | | | | | |
| *CCAPM.* | 5.8 | | | | 1 | 0.68 | 0.68 |
| *s.e.* | (0.15) | | | | | | |
| *HCAPM.* | 2.9 | | 0.76 | | 1 | 0.62 | 0.67 |
| *s.e.* | (0.023) | | (0.0098) | | | | |
| *cay-CCAPM.* | 5 | 1.1 | | | 1 | 0.62 | 0.68 |
| *s.e.* | (0.075) | (0.015) | | | | | |
| *my-CCAPM.* | 4.5 | 0.64 | | | 1 | 0.61 | 0.67 |
| *s.e.* | (0.22) | (0.038) | | | | | |
| *my-HCAPM.* | 3.3 | 0.29 | 0.51 | 0.077 | 1 | 0.68 | 0.82 |
| *s.e.* | (0.056) | (0.0014) | (0.011) | (0.0011) | | | |
| **Panel B: Quarterly Data** | | | | | | | |
| *1953.1-2002.4* | | | | | | | |
| *CCAPM* | 2.9 | | | | 0.15 | 0.4 | 0.4 |
| *s.e.* | (2) | | | | | | |
| *HCAPM* | 2.4 | | −0.044 | | 0.57 | 0.62 | 0.67 |
| *s.e.* | (1.6) | | (0.1) | | | | |
| *cay-HCAPM* | 2.7 | 0.71 | | | 0.1 | 0.31 | 0.41 |
| *s.e.* | (2) | (0.5) | | | | | |
| *my-CCAPM* | 2 | 0.14 | | | 0.11 | 0.4 | 0.49 |
| *s.e.* | (1.4) | (0.15) | | | | | |
| *my-HCAPM* | 6.3 | 0.62 | 0.11 | −0.0052 | 0.22 | 0.47 | 0.69 |
| *s.e.* | (10) | (1.1) | (0.43) | (0.043) | | | |
| *1971.1-2002.4* | | | | | | | |
| *CCAPM* | 1.5 | | | | 0.12 | 0.3 | 0.3 |
| *s.e.* | (0.65) | | | | | | |
| *HCAPM* | 2.5 | | −0.014 | | 0.73 | 0.41 | 0.5 |
| *s.e.* | (2.1) | | (0.14) | | | | |
| *cay-CCAPM* | 1.9 | 0.52 | | | 0.15 | −0.02 | 0.13 |
| *s.e.* | (0.89) | (0.25) | | | | | |
| *my-CCAPM* | 2 | 0.19 | | | 0.082 | 0.13 | 0.25 |
| *s.e.* | (1) | (0.13) | | | | | |
| 7.4 | 0.62 | | | | | | |
| *my-HCAPM* | 0.018 | −0.012 | 0.66 | 0.15 | 0.52 | | |
| *s.e.* | (16) | (1.4) | (0.77) | (0.11) | | | |

## Table IV
## CAPM Risk Price Estimates for Currency Portfolios

GMM estimates using 8 currency portfolios as test assets. The first column contains $\lambda_{R^{vw}}$, the second column $\lambda_{R^{SB}}$, the third column $\lambda_{R^{HL}}$, the fourth column $\lambda_{R^{long}}$, the fifth column $\lambda_{R^{corp}}$. For the $y$-CAPM, the second column contains $\lambda_{l/c,R^{vw}}$ We consider two samples: 1953-2003 and 1971-2002, at annual and quarterly frequency. We used the optimal lag length to estimate the spectral density matrix (Andrews, 1991).

| *Model* | $\lambda_1$ | $\lambda_2$ | $\lambda_3$ | $\lambda_4$ | $\lambda_5$ | *p value* | $R^2_{adj}$ | $R^2$ |
|---|---|---|---|---|---|---|---|---|
| **Panel A: Annual Data** | | | | | | | | |
| *1953-2002* | | | | | | | | |
| *CAPM* | 16 | | | | | 0.42 | 0.36 | 0.36 |
| *s.e.* | (5) | | | | | | | |
| *FF-CAPM e.* | 22 | $-34$ | 29 | | | 0.29 | 0.31 | 0.51 |
| *s.e.* | (28) | (49) | (46) | | | | | |
| *FF-CAPM b.* | | | | 3.9 | 1.7 | 0.45 | $-0.065$ | 0.087 |
| *s.e.* | | | | (2.5) | (1.3) | | | |
| *y-CAPM* | 19 | 21 | | | | 0.41 | 0.2 | 0.32 |
| *s.e.* | (7.3) | (8.2) | | | | | | |
| *1971-2002* | | | | | | | | |
| *CAPM* | 9.1 | | | | | 0.69 | 0.25 | 0.25 |
| *s.e.* | (3.5) | | | | | | | |
| *FF-CAPM e.* | 6.6 | 14 | $-6.4$ | | | 0.92 | $-0.7$ | $-0.21$ |
| *s.e.* | (5.4) | (6.8) | (9) | | | | | |
| *FF-CAPM b.* | | | | 7.7 | 1.7 | 0.64 | 0.3 | 0.4 |
| *s.e.* | | | | (6.5) | (1.8) | | | |
| *y-CAPM* | 9.2 | 9.8 | | | | 0.58 | $-2.7$ | $-0.6$ |
| *s.e.* | (3.7) | (3.9) | | | | | | |
| **Panel B: Quarterly Data** | | | | | | | | |
| *1953.1-2002.4* | | | | | | | | |
| *CAPM* | 5.4 | | | | | 0.011 | 0.12 | 0.12 |
| *s.e.* | (4) | | | | | | | |
| *FF-CAPM e.* | 7.5 | 0.39 | 4.5 | | | 0.054 | $-0.33$ | 0.052 |
| *s.e.* | (6.6) | (2.8) | (2.9) | | | | | |
| *FF-CAPM b.* | | | | 2.2 | 5.3 | 0.94 | 0.76 | 0.8 |
| *s.e.* | | | | (4.7) | (2.9) | | | |
| *y-CAPM* | 4 | 5.6 | | | | 0.0046 | 0.0048 | 0.15 |
| *s.e.* | (5.2) | (7.7) | | | | | | |
| *1971.1-2002.4* | | | | | | | | |
| *CAPM* | 7.6 | | | | | 0.025 | 0.095 | 0.095 |
| *s.e.* | (4) | | | | | | | |
| *FF-CAPM e.* | 20 | 0.93 | 7.8 | | | 0.71 | 0.041 | 0.32 |
| *s.e.* | (12) | (4) | (4.9) | | | | | |
| *FF-CAPM b.* | | | | 6.5 | 0.085 | 0.1 | $-0.21$ | $-0.04$ |
| *s.e.* | | | | (3.8) | (1) | | | |
| *y-CAPM* | 12 | 18 | | | | 0.028 | $-0.25$ | $-0.07$ |
| *s.e.* | (5.5) | (8.1) | | | | | | |

**Pricing Errors** In a separate Appendix in Table XII we report the pricing errors for all the models we have tested. Clearly the consumption-based models and the $y - CAPM$ do much better than the ad hoc factor models: the average pricing errors for the consumption-based models are only half the size of those for the ad hoc factor models. Nonetheless, all models overpredict the risk premium on the first portfolio by at least seventy basis points. This is still a large pricing error. However, the ad hoc factor models misprice the first portfolio by at least 1.2 percentage points, nearly twice as large.

**Post-Bretton Woods** The demise of the gold standard obviously increases the volatility of exchange rates in most developed countries. This may affect the distribution of the moments of our currency portfolio returns and affect our estimates. To guard against this possibility, we focus only on the post Bretton-Woods sample. These results are reported in the lower part of Table III. Very similar results are obtained. The CCAPM explains almost seventy percent of the variation in returns and the point estimates are sensible: the estimated price of consumption growth risk is around five; the price of scaled consumption growth risk is positive and significant.

## B. Stocks as Test Assets

A key question is whether currency risk is priced differently from equity risk and bond risk. We examine whether the compensation for aggregate risk in currency markets differs from the one applied in domestic equity markets, again from the perspective of a US investor by adding the 25 size and book-to-market portfolios constructed by Fama and French (see annex) to the eight currency portfolios. These FF-portfolios sort stocks according to size and book to market quintiles, because both size and book-to-market predict returns. We want to look at an interesting source of variation for domestic returns and check whether these returns can be priced by the same stochastic discount factor that prices currency risk.

We use a total of 33 moments to estimate the model: 25 equity moments and 8 currency moments. Figure 9 plots the predicted excess return on the horizontal axis against the actual excess return on the vertical axis. The filled dots represent the eight currency portfolios, while the empty dots represent the 25 Fama-French portfolios. The sample runs from 1953 to 2002.

The three factor model for stock pricing developed by Fama & French (1993) fails to price both the equity risk and the currency risk. The two factor model for bonds does much better, but all five of these factors are needed to get a reasonably close match between predicted and actual excess returns. The models that explicitly introduce macroeconomic risk invariably do much better at pricing both currency risk and equity risk in a parsimonious way. The average

**Table V**

**CCAPM Risk Price Estimates for Equity and Currency Portfolios**

GMM estimates using 8 currency and 25 equity portfolios as test assets. The first column contains $\lambda_c$, the second column $\lambda_x$, the third column $\lambda_{c,x}$, the fourth column $\lambda_\rho$ and the final column $\lambda_{\rho,x}$, where $x$ denotes the scaling factor. The upper panel reports estimates based on annual data, the lower reports estimates based on quarterly data. The collateral measure for the *my*-CCAPM and *my*-HCAPM is $myfa$. We used the optimal lag length to estimate the spectral density matrix (Andrews, 1991).

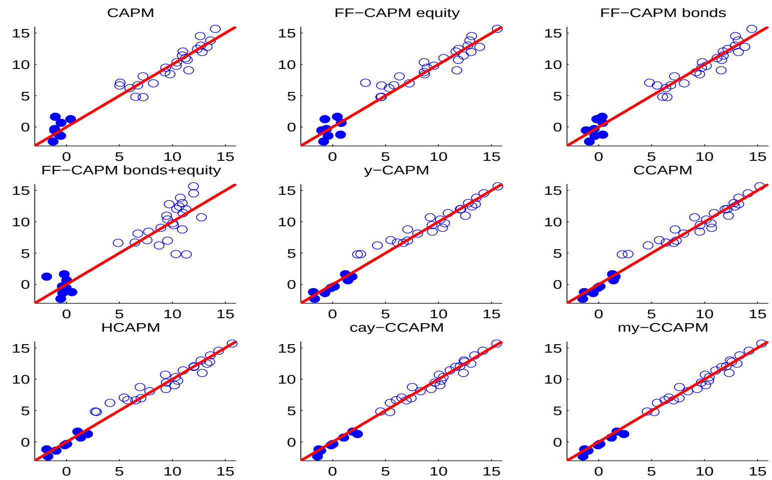
| *Model* | $\lambda_c$ | $\lambda_{c,x}$ | $\lambda_\rho$ | $\lambda_{\rho,x}$ | *p value* |
|---|---|---|---|---|---|
| **Panel A: Annual Data** | | | | | |
| | *1953-2002* | | | | |
| *CCAPM.* | 8 | | | | 1 |
| *s.e.* | (1.1) | | | | |
| *HCAPM.* | 5.5 | | 0.73 | | 1 |
| *s.e.* | (0.5) | | (0.07) | | |
| *cay-HCAPM.* | 8.1 | 1.7 | | | 1 |
| *s.e.* | (0.93) | (0.2) | | | |
| *my-CCAPM.* | 5.8 | 1 | | | 1 |
| *s.e.* | (0.84) | (0.13) | | | |
| *my-HCAPM.* | 5.6 | 0.9 | 0.43 | 0.077 | 1 |
| *s.e.* | (0.79) | (0.13) | (0.11) | (0.018) | |
| | *1971-2002* | | | | |
| *CCAPM.* | 5.8 | | | | 1 |
| *s.e.* | (0.15) | | | | |
| *HCAPM.* | 2.9 | | 0.76 | | 1 |
| *s.e.* | (0.023) | | (0.0098) | | |
| *cay-CCAPM.* | 5 | 1.1 | | | 1 |
| *s.e.* | (0.075) | (0.015) | | | |
| *my-CCAPM.* | 4.5 | 0.64 | | | 1 |
| *s.e.* | (0.22) | (0.038) | | | |
| *my-HCAPM.* | 3.3 | 0.29 | 0.51 | 0.077 | 1 |
| *s.e.* | (0.056) | (0.0014) | (0.011) | (0.0011) | |
| **Panel B: Quarterly Data** | | | | | |
| | *1953.1-2002.4* | | | | |
| *CCAPM* | 1.2 | | | | 0.99 |
| *s.e.* | (0.13) | | | | |
| *HCAPM* | 1.2 | | 0.14 | | 0.99 |
| *s.e.* | (0.14) | | (0.023) | | |
| *cay-CCAPM* | 1.2 | 0.3 | | | 0.99 |
| *s.e.* | (0.15) | (0.037) | | | |
| *my-CCAPM* | 1.6 | 0.11 | | | 0.99 |
| *s.e.* | (0.21) | (0.018) | | | |
| *my-HCAPM* | 1.7 | 0.12 | 0.24 | 0.016 | 0.99 |
| *s.e.* | (0.2) | (0.018) | (0.035) | (0.0041) | |
| | *1971.1-2002.4* | | | | |
| *CCAPM* | 1.1 | | | | 1 |
| *s.e.* | (0.084) | | | | |
| *HCAPM* | 1.2 | | 0.11 | | 1 |
| *s.e.* | (0.12) | | (0.027) | | |
| *cay-CCAPM* | 1.2 | 0.34 | | | 1 |
| *s.e.* | (0.099) | (0.028) | | | |
| *my-CCAPM* | 2 | 0.14 | | | 1 |
| *s.e.* | (0.23) | (0.022) | | | |
| *my-HCAPM* | 1.9 | 0.16 | 0.25 | 0.017 | 1 |
| *s.e.* | (0.25) | (0.032) | (0.045) | (0.005) | |

## Table VI
## CAPM Risk Price Estimates for Equity and Currency Portfolios

GMM estimates using 8 currency and 25 equity portfolios as test assets. The first column contains $\lambda_{R^{vw}}$, the second column $\lambda_{R^{SB}}$, the third column $\lambda_{R^{HL}}$, the fourth column $\lambda_{R^{long}}$, the fifth column $\lambda_{R^{corp}}$. For the $y$-CAPM, the second column contains $\lambda_{l/c,R^{vw}}$ We consider two samples: 1953-2003 and 1971-2002, at annual and quarterly frequency. We used the optimal lag length to estimate the spectral density matrix (Andrews, 1991).

| *Model* | $\lambda_1$ | $\lambda_2$ | $\lambda_3$ | $\lambda_4$ | $\lambda_5$ | *p value* |
|---|---|---|---|---|---|---|
| **Panel A: Annual Data** | | | | | | |
| | *1953-2002* | | | | | |
| *CAPM.* | 9.3 | | | | | 1 |
| *s.e.* | (0.73) | | | | | |
| *FF-CAPM e.* | 6.7 | 2.3 | 6.3 | | | 1 |
| *s.e.* | (0.84) | (0.8) | (0.5) | | | |
| *FF-CAPM b.* | | | | 6.1 | 4 | 1 |
| *s.e.* | | | | (0.89) | (0.42) | |
| *FF-CAPM b.+e.* | 7.1 | 2.4 | 6.4 | 6.2 | 1 | 1 |
| *s.e.* | (1) | (1) | (0.91) | (0.73) | (0.32) | |
| *y-CAPM.* | 6.5 | 7.6 | | | | 1 |
| *s.e.* | (0.9) | (0.98) | | | | |
| | *1971-2002* | | | | | |
| *CAPM* | 12 | | | | | 1 |
| *s.e.* | (0.71) | | | | | |
| *FF-CAPM e.* | 5.7 | 1.6 | 6 | | | 1 |
| *s.e.* | (0.31) | (0.28) | (0.34) | | | |
| *FF-CAPM b.* | | | | 8 | 2 | 1 |
| *s.e.* | | | | (0.053) | (0.03) | |
| *FF-CAPM e+b.* | 6.2 | 1.7 | 6.5 | 9.3 | 1.1 | 1 |
| *s.e.* | (0.25) | (0.24) | (0.26) | (0.31) | (0.069) | |
| *y-CAPM* | 4.5 | 5.3 | | | | 0.75 |
| *s.e.* | (1.1) | (1.1) | | | | |
| **Panel B: Quarterly Data** | | | | | | |
| | *1953.1-2002.4* | | | | | |
| *CAPM* | 2.1 | | | | | 0.99 |
| *s.e.* | (0.32) | | | | | |
| *FF-CAPM e.* | 1.7 | 0.47 | 1.6 | | | 0.99 |
| *s.e.* | (0.35) | (0.19) | (0.29) | | | |
| *FF-CAPM b.* | | | | 2.2 | 1.3 | 0.99 |
| *s.e.* | | | | (0.51) | (0.31) | |
| *FF-CAPM e.+b.* | 2.1 | 0.71 | 1.5 | 1.5 | 3.1 | 0.99 |
| *s.e.* | (0.61) | (0.29) | (0.38) | (0.67) | (0.5) | |
| *y-CAPM* | 1.7 | 3 | | | | 0.99 |
| *s.e.* | (0.62) | (0.91) | | | | |
| | *1971.1-2002.4* | | | | | |
| *CAPM* | 1.9 | | | | | 1 |
| *s.e.* | (0.34) | | | | | |
| *FF-CAPM e.* | 1.6 | 0.49 | 1.6 | | | 1 |
| *s.e.* | (0.45) | (0.23) | (0.33) | | | |
| *FF-CAPM b.* | | | | 2.4 | 0.081 | 1 |
| *s.e.* | | | | (0.35) | (0.16) | |
| *FF-CAPM e.+b.* | 1.5 | 0.73 | 1.7 | 1.4 | 0.95 | 1 |
| *s.e.* | (0.47) | (0.25) | (0.43) | (0.42) | (0.24) | |
| *y-CAPM* | 1.3 | 2.5 | | | | 1 |
| *s.e.* | (0.7) | (1) | | | | |

**Figure 9.** Predicted vs. Actual Excess Return for 8 Currency and 25 Stock Portfolios between 1953-2002. Predicted excess return on horizontal axis. GMM estimates using 25 equity and 8 currency portfolios as test assets. Annual Data. Each panel plots the results for one of the 9 linear factor models. The filled dots are the currency portfolios.

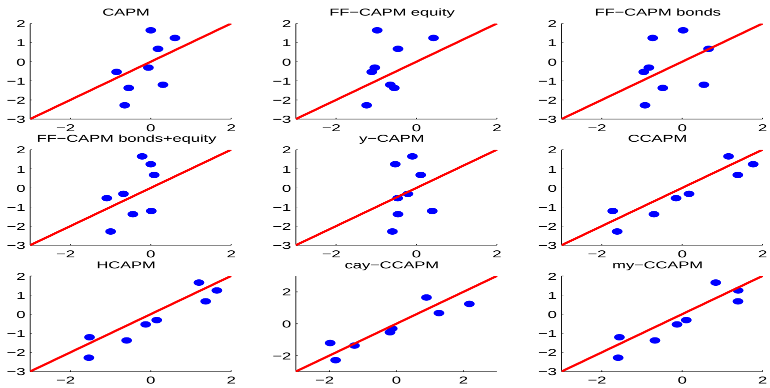



mean squared pricing error for the eight currency portfolios drops from 1.3 for the ad hoc factor models to around .6 for the consumption-based models. The factor models have very low explanatory power. More importantly, the consumption-based models also do well in explaining the excess returns on the 25 test assets. Pricing errors for these 25 portfolios are reported in a separate Appendix in Table XIII and Table XIV. In fact, the *my*-CCAPM produces an average pricing error for these 25 portfolios of .64 while the stock factors developed by Fama and French produce an average pricing error of 1.2. The coefficient estimates for consumption growth and for consumption growth interacted with the housing collateral ratio *my* and the consumption-wealth ratio *cay* are as predicted by the theory. The implied coefficient of risk aversion is around 50. The coefficients $b_{c,x}$ on consumption growth and the scaling variable are positive and significant, except for $b_{cay,c}$ (Results are reported in a separate Appendix in Table XV).

## C. Bonds as Test Assets

We have tried the same exercise using 6 bond portfolios of varying maturity as test assets. The results are quite similar. Only the consumption-based models can price bond risk and currency risk. The results are reported in Table XVI in the separate appendix.

**Figure 10.** Out of Sample: Predicted vs. Actual Excess Return for 8 Currency Portfolios between 1953-2002. Predicted excess return on horizontal axis. GMM estimates using 6 equity and 6 bond portfolios as test assets. Annual Data. Each panel plots the results for one of the 9 linear factor models. The filled dots are the currency portfolios.




## D. Out of Sample Test: Stocks and Bonds as Test Assets

Finally, this section performs an out-of-sample-test by estimating the model on the 6 bond portfolios and the 6 Fama-French benchmark portfolios. Table XVI confirms that only consumption-based models can truly price both equity, bond and currency risk. The first panel reports the results for annual data, the second panel reports the results for quarterly data using quarterly re-balanced currency portfolios.

Figure 10 confirms that only consumption-based models can truly price both equity, bond and currency risk, even though the pricing errors are quite large.

## E. Managed Portfolios as Test Assets

Instead of using these portfolios that change composition each period, the nominal interest rate differential itself can be used as an instrument $z_t$:

$$E_t\left[m_{t+1}\left(R^{i,\$}_{t+1}\frac{p_t}{p_{t+1}} - R^{US,\$}_{t,t+1}\frac{p_t}{p_{t+1}}\right)\right] z^i_t = 0, \tag{8}$$

where $z_t$ is the interest rate differential between the foreign T-bill and its US equivalent $R^{i,*}_{t,t+1} - R^{f,\$}_{t,t+1}$. We can think of this as applying GMM using the excess returns on managed portfolios as test assets, e.g. for country $i$ the return on its managed portfolio $\widetilde{R}^{e,i}_t$ is given by:

## Table VII
## GMM Pricing Errors for 8 Currency Portfolios

GMM estimates using 6 stock and 6 bond portfolios as test assets. Panel A reports the pricing errors for the factor models and Panel B reports the pricing errors for consumption-based models. The sample is 1953-2002, at annual frequency. The collateral measure for the $my$-CCAPM and $my$-HCAPM is $myfa$. We used 12 lags to estimate the spectral density matrix for the quarterly data and 4 lags for the annual data. The last column reports the square root of the average mean squared pricing error.

| *Model* | $R^2$ | $R^{2^2}_{adj}$ | *Average* |
|---|---|---|---|
| | **Panel I: Annual data 1953-2002** | | |
| | *Consumption-based Models* | | |
| *CCAPM.* | 0.804 | 0.804 | 0.606 |
| *HCAPM.* | 0.811 | 0.779 | 0.595 |
| *cay-CCAPM.* | 0.787 | 0.752 | 0.631 |
| *my-CCAPM.* | 0.8 | 0.766 | 0.612 |
| *my-HCAPM.* | 0.777 | 0.61 | 0.645 |
| | *Factor Models* | | |
| *CAPM .* | 0.299 | 0.299 | 1.14 |
| *FF-CAPM e.* | 0.0994 | −0.261 | 1.3 |
| *FF-CAPM b.* | 0.0308 | −0.131 | 1.35 |
| *FF-CAPM e.+b.* | 0.249 | −0.751 | 1.18 |
| *y-CAPM* | 0.0237 | −0.139 | 1.35 |
| | **Panel II: Quarterly data 1953.1-2002.4** | | |
| | *Consumption-based Models* | | |
| *CCAPM.* | 0.23 | 0.23 | 0.345 |
| *HCAPM.* | 0.572 | 0.5 | 0.257 |
| *cay-HCAPM.* | −3.95 | −4.78 | 0.875 |
| *my-CCAPM.* | 0.791 | 0.634 | 0.18 |
| *my-HCAPM.* | 0.269 | 0.147 | 0.336 |
| | *Factor Models* | | |
| *CAPM.* | −1.38 | −1.38 | 0.607 |
| *FF-CAPM e.* | −4.94 | −7.31 | 0.958 |
| *FF-CAPM b.* | −3.26 | −3.97 | 0.811 |
| *FF-CAPM e.+b.* | −0.175 | −1.74 | 0.426 |
| *y-CAPM* | −9.89 | −10 | 1.3 |

$$\widetilde{R}_t^{e,i} = \left[ R_{t+1}^{i,\$} \frac{p_t}{p_{t+1}} - R_{t,t+1}^{US,\$} \frac{p_t}{p_{t+1}} \right] \times z_t^i \tag{9}$$

These are the excess returns on portfolios that go long in a currency when its interest rate is high relative to the US, and short when its interest rate is low. The instrument variables GMM can be though of as applying GMM to the unconditional moment conditions for these managed returns.

The moment conditions are the sample analog of the population pricing errors:

$$g_T(b) = E_T(m_t \widetilde{R}_t^e) = E_T(\widetilde{R}_t^e) - E_T(\widetilde{R}_t^e f_t')b$$

where $R_t^e = [R_t^{1,e} \; R_t^{2,e} \; .. \; R_t^{Np,e}]'$ in the first stage of GMM we use the identity matrix as the weighting matrix, $W = I$, while in the second stage we use $W = S^{-1}$ where $S$ is the covariance matrix of the pricing errors in the first stage: $S = \sum_{-\infty}^{\infty} E[(m_t R_t^e)(m_{t-j} R_{t-j}^e)']$. Since we concentrate on linear factor models, the first stage is an OLS-cross-sectional regression of average prices on the second moment of payoff and factors, while the second stage is a GLS cross-sectional regression (Cochrane (2001)).

The advantage of this procedure is that we do not lose information by aggregating currencies into portfolios, as we did before. The disadvantage is that we need to restrict the study to countries with data over the whole sample. We construct these managed portfolios for 11 countries between 1971 and 2002. Table VIII reports the results. The explanatory power of the consumption-based models is lower than in the other pricing exercises. The risk premia that are due to movements in a currency's interest rate are harder to explain that the risk premia due to relative moments in one currency's interest rate, compared to all the others! Still, $\lambda_c$ and $\lambda_{c,x}$ are positive, significant and have the right sign. These estimates are in line with our previous estimates.

**Table VIII**
**CCAPM Risk Price Estimates for 11 Managed Currency Portfolios**

GMM estimates using 11 managed currency portfolios as test assets. Panel A reports the risk prices for the factor models and Panel B reports the results for consumption-based models. We consider one samples 1971-2002, at annual frequency. The collateral measure for the *my*-CCAPM and *my*-HCAPM is $myfa$. We used 4 lags to estimate the spectral density matrix.

| *Model* | $\lambda_c$ | $\lambda_{c,x}$ | $\lambda_\rho$ | $\lambda_{\rho,x}$ | *p value* | $R^2_{adj}$ | $R^2$ |
|---|---|---|---|---|---|---|---|
| *CCAPM.* | 4.4 | | | | 0.82 | −0.13 | −0.13 |
| *s.e.* | (0.87) | | | | | | |
| *HCAPM.* | 6.6 | | 0.25 | | 0.88 | −0.034 | 0.07 |
| *s.e.* | (3.7) | | (0.15) | | | | |
| *cay-CCAPM.* | 5 | 1.1 | | | 0.75 | −0.24 | −0.12 |
| *s.e.* | (1.4) | (0.29) | | | | | |
| *my-CCAPM.* | 3.7 | 0.55 | | | 0.8 | 0.27 | 0.34 |
| *s.e.* | (1.1) | (0.2) | | | | | |
| *my-HCAPM.* | 18 | 2.6 | 0.83 | 0.15 | 0.63 | 0.45 | 0.61 |
| *s.e.* | (26) | (3.8) | (1.1) | (0.21) | | | |

## V. Where do Exchange Rate Betas come from?

What gives rise to the monotonic relation between the consumption growth betas of exchange rates and interest rates in the data? The key is (1) negative correlation between the first and second moment of the foreign (nominal) stochastic discount factor, and/or (2) a higher correlation of the SDF between low interest rate currencies and the US. If the conditional volatility increases when the conditional mean of the foreign SDF decreases, then the currency risk premium switches signs with the foreign interest rate. To derive this result, we re-interpret a derivation by Backus et al. (2001).

## A. Exchange Rate Betas

First, we note that the log risk premium $log(crp^i_{t+1})$ can be decomposed into two parts:

$$-Cov_t\left(\log m_{t+1}, \Delta \log e^i_{t+1} - \Delta \log p_{t+1}\right) = -Cov_t\left(\log m_{t+1}, \Delta \log q^i_{t+1}\right) + Cov_t\left(\log m_{t+1}, \Delta \log p^i_{t+1}\right)$$

where $q^i = e^i \frac{p^i}{p}$ is the real exchange rate of country $i$. We focus on the first part and abstract from the inflation betas. To explain why the consumption growth betas of exchange rates switch signs, we assume that markets are complete to substitute the difference between the log stochastic discount factors for the change in the real exchange rate:

$$-Cov_t\left(\log m_{t+1}, \Delta \log q^i_{t+1}\right) = -Cov_t\left(\log m_{t+1}, \log m^i_{t+1} - \log m_{t+1}\right) \tag{10}$$

What do we learn from this? Under joint log-normality on the log SDF's, the sign of the covariance between the log SDF and the log change in the exchange rate, $-Cov_t\left(\log m_{t+1}, \Delta \log q^i_{t+1}\right)$, is determined by the standard deviation of the home SDF relative to the SDF of the foreign SDF, scaled by the correlation between the two:

$$sign\ \left[std_t \log m_{t+1} - Corr_t\left(\log m_{t+1}, \log m^i_{t+1}\right) std_t \log m^i_{t+1}\right]$$

**Heteroskedasticity** First, suppose the correlation between the SDFs is positive and constant. If countries characterized by a low conditional mean of the log stochastic discount factor, or a high interest rate, typically also have high conditional volatility of the nominal SDF, then the sign of the expression above will be negative. thus high interest countries will provide positive foreign risk premia. Through this mechanism, the betas switch sign between high and low nominal interest rate countries. This behavior is at the heart of the habits-based model of the exchange rate risk premium in Verdelhan (2003). The evidence from currency markets suggests that low interest rates signal an increase in the conditional market price of risk.

**Correlation** Second, suppose the conditional volatilities of the SDFs are constant. An increase in the conditional correlation of the SDFs with low interest rate currencies delivers negative conditional risk premia for those currencies, while a decrease in the conditional correlation with high interest rate currencies can deliver a positive conditional risk premia.

## B. Consumption Co-movements and Interest Rates

We start by considering the case of the Consumption CAPM and we assume all countries share the same coefficient of relative risk aversion. The sign of the conditional risk premium is determined by:

$$sign \ \left[std_t \log \Delta \log(c^{US}_{t+1}) - Corr_t \left(\Delta \log(c^{US}_{t+1}), \Delta \log(c^{i}_{t+1})\right) std_t \Delta \log(c^{i}_{t+1})\right]$$

An increase in the conditional correlation of foreign consumption growth with US consumption growth for low interest rate currencies can imply negative risk premia. The data seem to support the time-varying correlation mechanism.

Using a sample of twelve developed countries[8], we regressed a country's consumption growth on US consumption growth and US consumption growth interacted with the lagged interest rate differential:

$$\Delta \log(c_{t+1}) = \alpha_0 + \alpha_1 \Delta \log(c^{US}_{t+1}) + \alpha_2 \Delta \log(c^{US}_{t+1}) \left(R^{£}_{t,t+1} - R^{\$}_{t,t+1}\right) + \epsilon_{t+1}$$

These results are reported in Table IX. The coefficients on the interaction terms are negative for all countries, except for Japan and the Netherlands. The table also reports a ninety percent confidence interval for these interaction coefficients. The last row of each panel reports the pooled time series regression results. The ninety percent confidence interval includes only negative coefficients. These estimates are economically significant as well. The consumption growth sensitivity for the UK varies from 1 to -.2. This creates a lot of variation in the conditional risk premia. All else equal, the implied currency risk premium on the pound would be small or negative in the early eighties, when UK interest rates where low, and much larger in the mid-seventies, when UK interest rates were high relative to the US.

These results are reported in Table IX. The coefficients on the interaction terms are negative for all countries, except for Japan and the Netherlands. The table also reports a ninety percent

[8] We have used and updated the data set built by John Campbell and available on his web site.

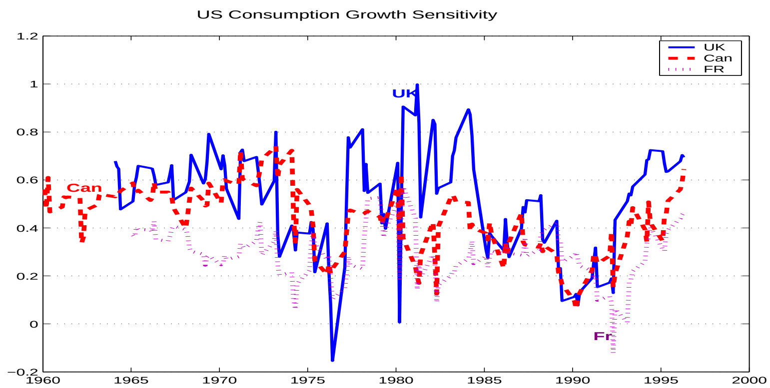
confidence interval for these interaction coefficients. The last row of each panel reports the pooled time series regression results. The ninety percent confidence interval includes only negative coefficients. These estimates are economically significant as well. The consumption growth sensitivity for the UK varies from 1 to -.2. This creates a lot of variation in the conditional risk premia. All else equal, the implied currency risk premium on the pound would be small or negative in the early eighties, when UK interest rates where low, and much larger in the mid-seventies, when UK interest rates were high relative to the US.

**Table IX**
**Consumption Growth Regressions**

Results for the following time-series regression: $\Delta \log(c_{t+1}) = \alpha_0 + \alpha_1 \Delta \log(c^{US}_{t+1}) + \alpha_2 \Delta \log(c^{US}_{t+1}) \left( R^{£}_{t,t+1} - R^{\$}_{t,t+1} \right) + \epsilon_{t+1}$. The last row reports the results from a pooled time series regression. The top panel reports the results for annual data. The bottom panel reports the quarterly results. We used the optimal lag length to estimate the spectral density matrix (Andrews, 1991). $\underline{\alpha_2}$ and $\overline{\alpha_2}$ correspond respectively to one standard error below and above the point estimate $\alpha_2$. Sample covers the post Bretton Woods period (or shorter periods when data are not available in 1971).

| *Country* | $\alpha_0$ | $\alpha_1$ | $\alpha_2$ | $\underline{\alpha_2}$ | $\overline{\alpha_2}$ | $R^2$ |
|---|---|---|---|---|---|---|
| | | | *Annual Data* | | | |
| *Australia, 1971-2002* | 0.02 | 0.071 | −0.06 | −0.086 | −0.033 | 0.13 |
| *Canada, 1971-2002* | 0.0094 | 0.58 | −0.095 | −0.15 | −0.039 | 0.26 |
| *France, 1971-2002* | 0.012 | 0.27 | −0.0058 | −0.092 | 0.081 | 0.056 |
| *Germany, 1971-2002* | 0.031 | −0.24 | −0.064 | −0.16 | 0.029 | 0.013 |
| *Italy, 1972-2002* | 0.029 | 0.26 | −0.06 | −0.098 | −0.022 | 0.072 |
| *Japan, 1971-2002* | 0.0095 | 0.71 | 0.072 | 0.003 | 0.14 | 0.26 |
| *Netherlands, 1978-2002* | 0.0096 | 0.21 | −0.11 | −0.17 | −0.057 | 0.15 |
| *Sweden, 1971-2002* | −0.044 | 0.59 | −0.24 | −0.39 | −0.089 | 0.18 |
| *Switzerland, 1981-2002* | 0.012 | −0.39 | −0.07 | −0.1 | −0.037 | 0.19 |
| *United Kingdom, 1971-2002* | 0.017 | 0.74 | −0.1 | −0.15 | −0.052 | 0.21 |
| *pooled, 1971-2002* | 0.0092 | 0.27 | −0.047 | −0.088 | −0.0073 | 0.038 |
| | | | *Quarterly Data* | | | |
| *Australia, 1971:1-2002:4* | 0.0049 | 0.086 | −0.064 | −0.087 | −0.04 | 0.058 |
| *Canada, 1971:1-2002:4* | 0.0028 | 0.44 | −0.067 | −0.12 | −0.011 | 0.094 |
| *France, 1971:1-2002:4* | 0.0031 | 0.35 | −0.052 | −0.1 | −0.0021 | 0.031 |
| *Germany, 1971:1-2002:4* | 0.0077 | −0.27 | −0.082 | −0.21 | 0.045 | 0.0087 |
| *Italy, 1971:1-2002:4* | 0.0075 | 0.093 | −0.046 | −0.073 | −0.018 | 0.036 |
| *Japan, 1971:1-2002:4* | 0.0028 | 0.53 | 0.0062 | −0.043 | 0.055 | 0.092 |
| *Netherlands, 1977:2-2002:4* | 0.0029 | 0.28 | 0.074 | 0.017 | 0.13 | 0.032 |
| *Sweden, 1971:1-2002:4* | −0.013 | 0.42 | −0.076 | −0.14 | −0.0087 | 0.026 |
| *Switzerland, 1980:2-2002:4* | 0.0015 | −0.0097 | −0.019 | −0.043 | 0.0042 | 0.0093 |
| *United Kingdom, 1971:1-2002:4* | 0.0033 | 0.91 | −0.1 | −0.15 | −0.056 | 0.12 |
| *pooled, 1971:1-2002:4* | 0.0024 | 0.23 | −0.035 | −0.06 | −0.0094 | 0.018 |

**Figure 11.** Consumption Growth Slope Coefficients. This figure plots the estimated consumption growth slope coefficients evaluated at the interest rate differential: $\alpha_c + \alpha_{c,\Delta R}\left(R^{£}_{t,t+1} - R^{\$}_{t,t+1}\right)$. The plot shows the UK (full line), Canada (dashed line) and France (dotted line)




**Towards A Model of Exchange Rates** This also delivers a linear factor model of exchange rates:

$$E_t(\log e^i_{t+1} - \Delta \log p_{t+1}) = -\sum_{j=1} b_j(x^t) Cov_t\left(\log F_{j,t+1}, \Delta \log e^i_{t+1}\right) + \left(\log R^f_{t,t+1} - \log R^{i,£}_{t,t+1}\right).$$

To the extent that we can price currency risk premia, we actually explain changes in the exchange rates!

## VI. Conclusion

Currency risk seems to be priced much like domestic equity risk. The pattern of average returns on currency portfolios is well explained by a class of linear macro-economic factor models. These models even perform well when the test assets also include domestic equity returns. The factor models constructed explicitly to price domestic equity risk break down when confronted with currency risk.

These results lead us to believe that macroeconomists need to understand how risk premia evolve in different countries in order to understand the behavior of (real) exchange rates. Exchange rates constantly adjust to eliminate arbitrage opportunities in asset markets created by the difference between the domestic and foreign SDF, and these exchange rate changes can be large. This might create other arbitrage opportunities in goods markets, but these are probably harder to exploit.

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## A. Proofs

### A.1. Pricing with Log-normality

**Log Currency Risk Premium** We assume the pricing kernel and portfolio returns are conditionally log-normal. Assume that the pricing kernel has the following form:

$$\log m_{t+1} = b_0 + \sum_{j=1}^{n} b_j(x_t) \log F_{j,t+1}.$$

Let $x_t$ be some vector of random variables. Assume that both $\log F_{i,t+1}$ and $\log R^i_{t+1}$ are normal so that $\log m_{t+1} + \log R^i_{t+1}$ is also normal. Returns are priced using the Euler equation:

$$E_t m_{t+1} R^i_{t+1} = 1$$

Hence,

$$\log E_t m_{t+1} R^i_{t+1} = 0$$

and, with log-normality

$$\log E_t m_{t+1} R^i_{t+1} = E_t \left(\log m_{t+1} + \log R^i_{t+1}\right) + \frac{1}{2} Var_t \left(\log m_{t+1} + \log R^i_{t+1}\right) = 0$$

This implies that the Euler equation can be restated as:

$$E_t \log m_{t+1} + E_t \log R^i_{t+1} + \frac{1}{2}\left[Var_t \log m_{t+1} + Var_t \log R^i_{t+1}\right] + Cov_t \left(\log m_{t+1}, \log R^i_{t+1}\right) = 0$$

Let $R^f_{t,t+1}$ be the risk free rate known at $t$, then $\log R^f_{t,t+1} = -\log E_t m_{t+1}$. Since $\log E_t m_{t+1} = E_t \log m_{t+1} + \frac{1}{2} Var_t \log m_{t+1}$ and likewise for $R^i_{t+1}$, we get:

$$\log E_t R^i_{t+1} - \log R^f_{t,t+1} = -Cov_t \left(\log m_{t+1}, \log R^i_{t+1}\right).$$

We know that:

$$\log R^i_{t+1} = \log R^{i,£}_{t,t+1} + \Delta \log e^i_{t+1} - \Delta \log p_{t+1},$$

where $e^i_t$ is the exchange rate between the currency of country $i$ and the dollar. The log currency risk premium is then equal to:

$$log(crp^i_{t+1}) = -Cov_t \left(\log m_{t+1}, \Delta \log e^i_{t+1}\right) + Cov_t \left(\log m_{t+1}, \Delta \log p_{t+1}\right)$$

or, abstracting from the inflation risk premium for now:

$$log(crp_{t+1}^{i}) = -\sum_{j=1} b_j(x^t) Cov_t \left( \log F_{j,t+1}, \Delta \log e_{t+1}^{i} \right).$$

**Scaled CCAPM** In the case of the scaled CCAPM, this equation becomes:

$$\log m_{t+1} = b_0 + (b_1 + b_2 x_t) \left( \Delta \log c_{t+1} \right),$$

which produces the following expression for the risk premium:

$$log(crp_{t+1}^{i}) = -\left(b_1 + b_2 x_t\right) Cov_t \left( \Delta \log c_{t+1}, \Delta \log e_{t+1}^{i} - \Delta \log p_{t+1} \right)$$

Notice the difficulty of matching the observation that currencies with high interest rates (at each date $t$) offer a high rate of return relative to currencies with low interest rates. Consider for example what would happen if

$$Cov_t \left( \Delta \log c_{t+1}, \Delta \log e_{t+1}^{i} \right)$$

were constant over time. In that case, there would be country specific risk premia that might fluctuate over time because $x_t$ changes, but there would be no tendency for these risk premia to be associated with currencies with temporarily high interest rates.

To get the observation that it is currencies with high interest rates that offer a high rate of return, it is necessary to show that

$$Cov_t \left( \Delta \log c_{t+1}, \Delta \log e_{t+1}^{i} \right)$$

and/or

$$Cov_t \left( \Delta \log p_{t+1}, \Delta \log e_{t+1}^{i} \right)$$

vary systematically with the interest rate in the foreign currency. Note that the assumption that the variability of the US pricing kernel (brought about by the introduction of $x_t$) does not seem like it should help that much in accounting for the observation here.

### A.2. Proof of Condition on Covariance of SDFs.

We assume that the pricing kernel is conditionally log-normal and we assume complete markets so that in each state of the world tomorrow the value of a dollar delivered tomorrow, in terms of dollars today, equals the value of a unit of foreign currency tomorrow delivered in the same state, in units of currency today:

$$\frac{e^i_{t+1}}{e^i_t} = \frac{m^i_{t+1}}{m^\$_{t+1}}.$$

The log risk premia on currencies given by

$$log(crp^i_{t+1}) = -Cov_t\left(\log m^\$_{t+1}, \log e^i_{t+1} - \log e^i_t\right).$$

Under the assumption of complete markets, this risk premium is given by

$$-Cov_t\left(\log m^\$_{t+1}, \log m^i_{t+1} - \log m^\$_{t+1}\right) =$$

$$Var_t \log m^\$_{t+1} - Cov_t\left(\log m^\$_{t+1}, \log m^i_{t+1}\right) =$$

$$Var_t \log m^\$_{t+1} - Corr_t\left(\log m^\$_{t+1}, \log m^i_{t+1}\right) std_t \log m^\$_{t+1} std_t \log m^i_{t+1} =$$

$$std_t \log m^\$_{t+1}\left[std_t \log m^\$_{t+1} - Corr_t\left(\log m^\$_{t+1}, \log m^i_{t+1}\right) std_t \log m^i_{t+1}\right].$$

Note that to get the observation that at date $t$, currencies with high nominal interest rates have a high expected rate of return, we should be thinking in terms of $std_t \log m^\$_{t+1}$ as being fixed and what is varying across currencies is either $Corr_t\left(\log m^\$_{t+1}, \log m^i_{t+1}\right)$ or $std_t \log m^i_{t+1}$.

## B. Data

### B.1. Panel

Our panel includes 81 countries. We include each of the following countries for the dates noted in parenthesis: Angola (2001-2002), Australia (1953-2002), Austria (1960-1991), Belgium (1953-2002), Bangladesh (1984-2001), Bulgaria (1992-2002), Bahrain (1987-2002), Bolivia (1994-2002), Brazil (1996-2002), Barbados (1966-2002), Botswana (1996-2002), Canada (1953-2002), Switzerland (1980-2002), Chile (1997-2002), China (2002-2002), Colombia (1998-2002), Costa-Rica (2000-2002), Cyprus (1975-2002), Czech Republic (1996-2000), Germany (1953-2002), Denmark (1976-2002), Egypt (1991-2002), Spain (1985-2002), France (1960-2002), United Kingdom (1953-2002), Ghana (1978-2002), Greece (1985-2002), Hong-Kong (1991-2002), Honduras (1998-2001), Croatia (2000-2002), Hungary (1988-2002), India (1993-2002), Ireland (1969-2002), Iceland (1987-2002), Israel (1995-2002), Italy (1953-2002), Jamaica (1953-2002), Japan (1960-2002), Kenya (1997-2002), Kuwait (1979-2002), Kazakhstan (1994-2002), Lebanon (1977-2002), Sri Lanka (1982-2002), Lithuania (1994-2001), Latvia (1994-2002), Mexico (1978-2002), Macedonia (1997-2002), Malta (1987-2002), Mauritius (1996-2002), Malaysia (1961-2002), Namibia (1991-2002), Nigeria (n.a), Netherlands (1953-2002), Norway (1984-2002), Nepal (1982-2002), New-Zealand (1978-2002), Pakistan (1997-2002), Philippines (1976-2002), Poland (1992-2002), Portugal (1985-2002), Rumania (1994-2002), Russian Federation (1994-2002), Singapore (1987-2002), El Salvador (2001-2002), Slovak Republic (1993-2002), Slovenia (1998-2002), Sweden (1955-2002), Swaziland (1981-2002), Thailand (1997-2002), Trinidad and Tobago (1964-2002), Tunisia (1990-2002), Turkey (1985-2002), Taiwan (1974-2002), Uruguay (1992-2002), United States (1953-2002), Venezuela (1996-2002), Vietnam (1997-2002), Serbia and Montenegro (2002-2002), South Africa (1988-2002), Zambia (1978-2002), Zimbabwe (1962-2002). The exchange and T-bill rates were downloaded from Global Financial Data. The maturity of the T-bill rates is 3 months, except for Costa-Rica and Poland (both 6 months). The time period for each country is determined by data availability and openness of the financial market (according to Quinn (1997)'s index, see below).

### B.2. Defaults

We have used the dataset compiled by Reinhardt et al. (2003) to identify defaults on rated and unrated sovereign debt which occurred after 1953: Angola (1985-99), Bulgaria (1990-94), Bolivia (1980-84, 1986-97), Brazil (1983-94), Chile (1983-85), Czech Republic (1959-60),Germany (1953), Egypt (1984), Ghana (1987), Greece (1953-64), Honduras (1981-1999), Croatia (1992-1996), Hungary (1953-1967), Jamaica (1978-79, 1981-85, 1987-93), Mexico (1982-90), Macedonia (1992-1997), Nigeria (1982-92), Pakistan (1998-99), Philippines (1983-1992), Poland (1981-1994), Rumania (1953-58, 1981-83, 1986), Slovenia (1992-96), Trinidad and Tobago (1988-89), Turkey (1978-79, 1982), Uruguay (1983-85, 1987, 1990-91), Venezuela (1983-88, 1990, 1995-97), Vietnam (1985-1998), Serbia and Montenegro (1983-1999), South Africa (1985-87, 1989, 1993), Zambia (1983-92), Zimbabwe (1965-80).

### B.3. Recovery Rates

First, Moody's research studies twenty-four defaulted sovereign bonds issued by seven countries. They compute the average of the face value thirty days after default. They obtain a recovery rate of thirty-four percent on an issue-based computation (and forty-one percent on an issuer-based one). These figures are biased downward as they do not include the Peruvian and Venezuelan cases. Second, Singh (2003) computes the recovery rate as the ratio of post-restructuring prices on average post-default prices. The sample considers seven debt restructuring events for four sovereigns (Ukraine, Ecuador, Russia and Ivory Coast). The author finds that the average debt work-out period is two years and the weighted average recovery rate is one hundred and fifteen percent. This figure might still be biased downwards as bond prices continued to rise after the two-year window. We have assumed a recovery rate of seventy percent.

### B.4. Capital Account Liberalization

The IMF distinguishes between Current Account Restrictions (on payments for goods and services) and Capital Account Restrictions. The IMF distinguishes further between Exchange Payments and Exchange Receipts. Quinn (1997) adhered to the IMF categories and used the following coding rule for capital payments and receipts: (1) if approval is rare and surrender of receipts is required: X=0, (2) if approval is required and sometimes granted: X=0.5, (3) if approval is required and frequently granted: X=1, (4) if approval is not required and receipts are heavily taxed: X=1, (5) if approval is not required and receipts are taxed: X=1.5 and (6) if approval is not required and receipts are not taxed: X=2.

This algorithm yields a 0-4 code for each country. The index is ten mapped onto a scale from zero to hundred. Quinn (1997)'s capital account liberalization index ranges from zero to one hundred. When working with annual data, we chose a cut-off value of 20: we eliminate countries where approval of both capital payments and receipts are rare, or when payments or receipts are at best only infrequently granted.

### B.5. Financial Data and Macroeconomic Factors

**Returns** We obtained the Fama-French factors and the 25 book-to-market portfolios for the US from Kenneth French's web site at *mba.tuck.dartmouth.edu/pages/faculty/ken.french*. The portfolios, which are constructed at the end of each June, are the intersections of 5 portfolios formed on size (market equity, ME) and 5 portfolios formed on the ratio of book equity to market equity (BE/ME). The size breakpoints for year t are the NYSE market equity quintiles at the end of June of t. BE/ME for June of year t is the book equity for the last fiscal year end in t-1 divided by ME for December of t-1. The BE/ME breakpoints are NYSE quintiles.

**Consumption Data** The consumption data were downloaded from John Campbell's web site at *http://kuznets.fas.harvard.edu/ campbell/data.html*. These data were used for "Asset Prices, Consumption, and the Business Cycle", Chapter 19 in Handbook of Macroeconomics, John Taylor and Michael Woodford eds., North-Holland, Amsterdam, 1999. We have updated the data set using Datastream and IFS series along John Campbell's guidelines. We use per capita consumption deflated by that country's CPI.

**Other Variables** $my$ is defined as the ratio of collateralizable housing wealth to non-collateralizable human wealth. We use three distinct measures of the housing collateral stock $HV$: the value of outstanding home mortgages ($mo$), the market value of residential real estate wealth ($rw$) and the net stock current cost value of owner-occupied and tenant occupied residential fixed assets ($fa$). The first two time series are from the Historical Statistics for the US (Bureau of the Census) for the period 1889-1945 and from the Flow of Funds data (Federal Board of Governors) for 1945-2001. The last series is from the Fixed Asset Tables (Bureau of Economic Analysis) for 1925-2001. To approximate the ratio of housing wealth to human wealth, deviations from a cointegrating relation between log labor income and log housing wealth (see Lustig & VanNieuwerburgh (2002)). The data are available from Stijn Van Nieuwerburgh's web site at *http://pages.stern.nyu.edu/ svnieuwe/*. $cay$, the consumption-wealth ratio, is computed as the residual from a cointegrating relation between log labor income and total wealth (see Lettau & Ludvigson (2001)). The data are available from Martin Lettau's web site at *pages.stern.nyu.edu/ mlettau/*.