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

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Abstract

Aggregate consumption growth risk explains why low interest rate currencies do not appreciate as much as the interest rate differential and why high interest rate currencies do not depreciate as much as the interest rate differential. We sort foreign T-bills into portfolios based on the nominal interest rate differential with the US, and we test the Euler equation of a US investor who invests in these currency portfolios. US investors earn negative excess returns on low interest rate currency portfolios and positive excess returns on high interest rates currency portfolios. We find that low interest rate currencies provide US investors with a hedge against US aggregate consumption growth risk, because these currencies appreciate on average when US consumption growth is low, while high interest rate currencies depreciate when US consumption growth is low. As a result, the risk premia predicted by the Consumption-CAPM match the average excess returns on these currency portfolios.

Keywords: Exchange Rates, Asset Pricing, Risk Sharing.

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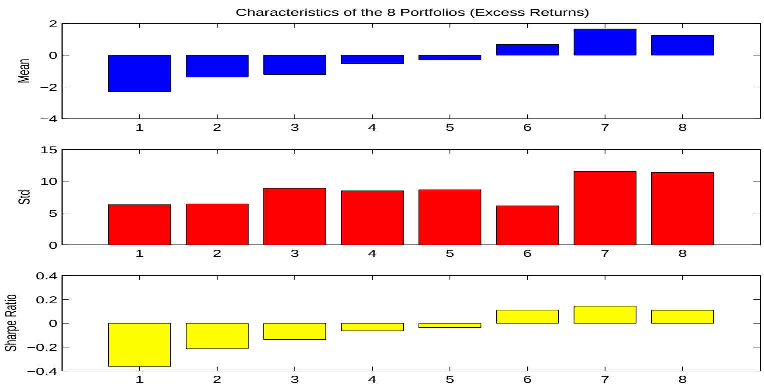
Aggregate consumption growth risk explains why low interest rate currencies do not appreciate as much as the interest rate differential and why high interest rate currencies do not depreciate as much as the interest rate differential. We sort foreign T-bills into portfolios based on the nominal interest rate differential with the US, and we test the Euler equation of a US investor who invests in these currency portfolios. 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: The figure presents means, standard deviations and Sharpe ratios of real ex-post excess returns on 8 currency portfolios. Currencies (listed in the Appendix) are allocated each year to portfolios on the basis of the interest rate differential with the US at the end of the previous year. The data are annual between 1953 and 2002.

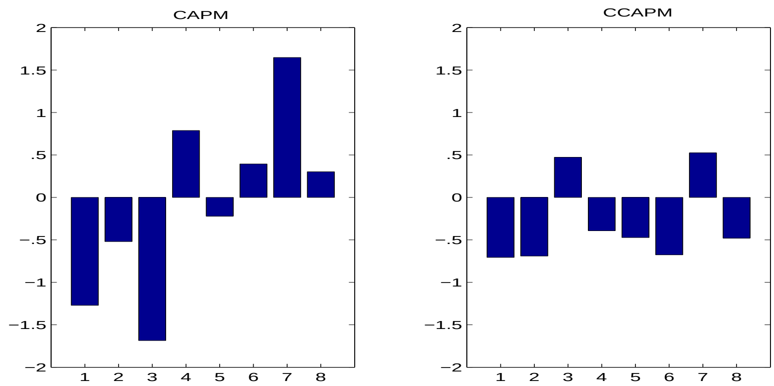


Why do interest rates predict real excess returns on currencies? We find that US consumption growth risk explains much of this variation in foreign real excess returns. Figure 2 plots the pricing errors on those same eight currency portfolios (see figure 1) for two workhorse models: the Capital Asset Pricing Model (CAPM) and the Consumption-CAPM. The CAPM pricing errors are almost identical to the average excess returns, but the pricing errors for the Consumption-CAPM are much smaller. Why does the Consumption-CAPM do so well?

High interest rate currencies typically depreciate when real 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. The consumption growth betas of exchange rates measure the sensitivity of the return on foreign currency holdings to changes in US consumption growth: these are negative for low interest rate currencies and positive for high interest rate currencies. All our results build on this basic finding. The market price of aggregate (consumption growth) risk is large, but comparable to that typically found when equity risk is studied. In this sense, the forward premium puzzle looks like a standard asset pricing puzzle.

We also find that ad hoc factors that price bond and equity risk do not price currency risk, but macroeconomic factor models that exploit conditioning information do. There is no relation between the US market beta of exchange rates and the foreign interest rate, while there is strong one for the US consumption growth betas of exchange rates. That is why the CAPM and extensions of the CAPM do not work.

Figure 2: Pricing errors for Currency Portfolios The left panels shows the CAPM pricing errors for 8 currency portfolios. The right panel shows the CCAPM pricing errors. GMM estimates using 8 currency portfolios as test assets for the CAPM and the Consumption-CAPM. The sample is 1953-2002 (annual data).



What gives rise to the monotonic relation between the consumption growth betas of exchange rates and interest rates? In complete markets, foreign currency appreciates when the state price of consumption is higher abroad than at home, and depreciates when the condition is reversed. If this were not the case, there would be an arbitrage opportunity. Thus, if the foreign state price of consumption is both more volatile than and highly correlated with its domestic counterpart, foreign currency provides a natural hedge to US investors.

What's the economics behind this? In the benchmark representative agent model, complete markets imply that foreign currency appreciates when foreign consumption growth is lower than US aggregate consumption growth and depreciates when it is higher. Now, if the foreign stand-in agent's consumption growth is strongly correlated with *and* more volatile than his US counterpart, his national currency provides a hedge for the US representative agent.

Take the benchmark representative agent model with power utility and risk aversion coefficient γ.If the foreign stand-in agent's consumption growth is strongly correlated with, and, more volatile than, its US counterpart, his national currency provides a hedge for the US representative agent. Why? Consider the case of perfectly correlated foreign consumption growth that is twice as volatile as US consumption growth. When consumption growth is -2 percent in the US, it is twice as low abroad (-4 percent) and the real exchange rate appreciates by γ times 2 percent. This currency is a perfect hedge against US aggregate consumption growth risk. So, in the benchmark representative agent model, in order to explain why foreign interest rates predict excess returns on foreign T-bills, low interest rate currencies must have aggregate consumption growth processes that are conditionally more volatile than and more correlated with US aggregate consumption growth.

We find strong evidence for increased correlation in the data. Foreign consumption growth is more correlated with US consumption growth when the foreign interest rate is low. In a sample of 10 developed countries, we find that the conditional correlation between foreign and US annual consumption growth decreases with the interest rate gap for all countries except Japan.

Currency Portfolios We test different pricing models on foreign currency portfolios. Building foreign currency portfolios, on the basis of what the investor knows at the time of her decision, the foreign and domestic interest rate, serves 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. This spread is an order of magnitude larger than the average 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.

Most traditional exchange rate models have proven largely unsuccessful in explaining and/or predicting exchange rates. Meese and 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 that we are less likely to miss important information in the investor's information set by focusing only on interest rate differentials.

We consider two large classes of pricing models. The first class uses returns as pricing factors (e.g. Fama and French (1992) and Santos and Veronesi (2001)). The second class introduces measures of macroeconomic undiversifiable risk. These models have proven successful in explaining the cross-sectional variation in US stock returns (e.g. Bansal, Dittmar and Lundblad (2005), Lettau and Ludvigson (2001), Lustig and Nieuwerburgh (2005), and Cochrane (2001) for an overview). These macroeconomic models suggest that conditioning information is important in order to characterize risk premia.

We test the US investor's Euler equation in two ways. First, we minimize the pricing errors on these currency portfolios using a GMM estimator. Second, we check the robustness of our results for a smaller set of countries by testing the investor's Euler equation on each currency. Instead of using portfolios that change composition each period, we use the nominal interest rate differential itself as an instrument. This procedure is equivalent to the pricing of excess returns on managed portfolios. In this paper we report results obtained through the first method (GMM) on annual and quarterly data for the periods 1953-2002 and 1971-2002. We also present results obtained through the second method (managed portfolios) on annual data for the same two periods.[1]

In this framework, we show that at annual frequencies the Consumption-CAPM (henceforth CCAPM) explains up to eighty percent of the variation in currency excess returns 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 and bond markets. If we estimate the models only on US domestic stock portfolios sorted by book-to-market and size and on US domestic bonds, we can still explain some of the variation in currency excess return.

Related Literature Our paper makes contact with at least three strands of the exchange rate literature. First, there is large literature on the forward premium puzzle. 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 and Hodrick (1980) and Fama (1984)).[2] 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. The main point of our paper is to show that currency excess returns predicted

[1] To save space, results obtained using managed portfolios on quarterly data are presented in a separate Appendix available on the authors web sites .

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

by real factor models, with aggregate consumption growth as a key factor, do line up with the predictable excess returns in currency markets.

Second, there is a large literature that relates predictable excess returns to currency risk premia. Our paper is closest to recent work by Hollifield and Yaron (2001), Harvey, Solnik and Zhou (2002) and Sarkissian (2003). Hollifield and Yaron (2001) find some evidence that real factors, not nominal factors, drive most of the predictable variation in currency risk premia. Using a latent factor technique on a sample of international bonds, Harvey et al. (2002) find empirical evidence of a factor premium which is related to foreign exchange risk. 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 premia, on a currency-by-currency basis, while most of the variation is conditional on the level of the foreign interest rate. Finally, Backus, Foresi and Telmer (2001) show that in a general class of affine models, the state variables need to have asymmetric effects on the state prices in different currencies to explain the forward premium puzzle. We reinterpret their results in our setting, to explain the relation between interest rates and the consumption growth betas of exchange rates.

There is a third literature that relates the volatility and persistence of real exchange rates to consumption. In any standard, dynamic equilibrium model, the theory implies a strong link between consumption ratios and the real exchange rate, but Backus and Smith (1993) pointed out that there is no obvious link in the data. This lack of correlation motivates the work by Alvarez, Atkeson and 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 (2004a). He replicates the forward discount bias in a model with external habits and provides estimates to support this mechanism.

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 mechanisms at the core of our results.

1 Framework

This section first defines the excess returns on foreign T-bill investments and it sets up the corresponding Euler equations for a US investor. We then present our data set, we explain how we construct the currency portfolios and, finally, we link the returns on these currency portfolios to the forward premium puzzle.

1.1 Definitions

We focus on a US investor who trades 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 or SDF, in the sense of Hansen and Jagannathan (1991). 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.

These Euler equations impose testable restrictions on the joint distribution of the US SDF, the exchange rate and interest rates. We do not make any assumptions about the span of markets, and we ignore the foreign SDF altogether, at least for now. Instead, we concentrate on the US investor's Euler equation, in terms of excess returns:

$$E_t \left[m_{t+1} \left(R_{t+1}^{i,\$} \frac{p_t}{p_{t+1}} - R_{t,t+1}^{\$} \frac{p_t}{p_{t+1}} \right) \right] = 0. \tag{2}$$

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

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.[3] 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 each portfolio.[4] 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 became available, new countries were added to these portfolios. As a result, the composition of the portfolio as well as the number of countries in a portfolio 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 $E_T\left[R^{i,e}_{t+1}\right], i = 1, \ldots, N^c$, because foreign interest rates fluctuate. The foreign excess return is positive (negative) when foreign interest rates are high (low), yet the average interest rate differential with the US tends to be rather small for most countries, thus period of high excess returns average out with periods of low excess returns. On the contrary, our portfolios exploit conditioning information and emphasize the distinction between high and low interest rate states.

Sample We always use a total number of eight portfolios. Given the limited number of countries, we did not want too many portfolios. With these eight portfolios, we consider two different time-horizons. First, we study the period 1953 to 2002, which spans a number of different exchange rate arrangements. This poses no inherent problem, because the Euler equation restrictions are valid regardless of the exchange rate regime. Second, we consider a shorter time period, 1971 to 2002, beginning with the demise of Bretton-Woods. For each time-horizon, we work successively with annual and quarterly data. Two additional problems arise: the existence of default events, and the effects of financial liberalization.

[3]Portfolios are rebalanced every quarter when we work on quarterly data and every year when we use annual data.

[4]Weights are computed from the capital flows available on the US Treasury website (using data on the gross purchases of foreign bonds by foreigners from US residents and data on the gross sales of foreign bonds by foreigners to US residents).

Default To compute the actual returns on a T-bill investment after default, we used the dataset of defaults compiled by Reinhardt, Rogoff and Savastano (2003).[5] 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 first quarter and we exclude these countries from the sample after the first quarter of the year in which they defaulted. In the annual data sample, we have assumed that, after the first default of a series of defaults, investors only lend when they expect their money back. Thus we have applied a 100% recovery rate for the next subsequent defaults of that series if capital still flows into that country. In the entire sample from 1953 to 2002, there are thirteen instances of default by a country whose currency is in one of our portfolios: Zimbabwe (1965), Jamaica (1978), Jamaica (1981), Mexico (1982), Brazil (1983), Philippines (1983), Zambia (1983), Ghana (1987), Jamaica (1987), Trinidad and Tobago (1988), South Africa (1989), South Africa (1993), Pakistan (1998). Of course, many more countries actually defaulted over this sample (see appendix), but those are not in our portfolios because they imposed capital controls, as explained in the next paragraph.

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 where payments or receipts are at best only infrequently granted.

1.2 Forward Premium Puzzle

UIP cross countries Uncovered interest rate parity (UIP) is a special case of the Euler equation in (2) when investors are risk-neutral and m is constant.[6] When the UIP condition is satisfied, the interest rate differential should be an unbiased predictor of changes in the exchange

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

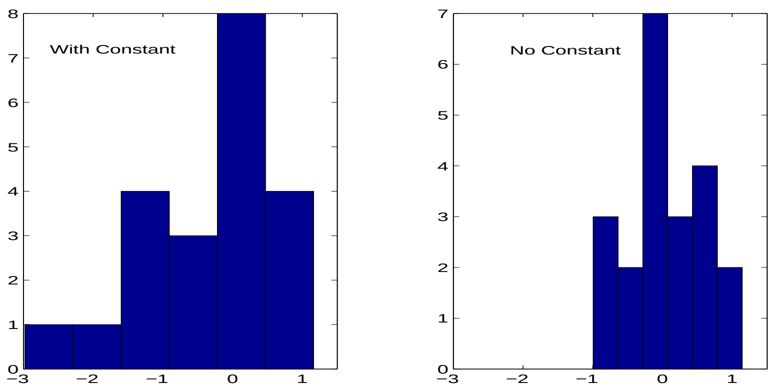
[6] When m is constant, equation (2) leads to: $E_t[(1+r^{i,£}_{t+1})(1+\frac{\Delta e^i}{e^i})(1-\pi_{t,t+1})-(1+r^{\$}_{t,t+1})(1-\pi_{t,t+1})]=0$, where π denotes the US inflation rate, $\frac{\Delta e}{e}$ is the change in the nominal exchange rate, $r^{i,£}$ and $r^{\$}$ are respectively the foreign and domestic interest rates. A first-order Taylor approximation gives: $E_t[\frac{\Delta e^i}{e^i}+r^{i,£}_{t+1}-r^{\$}_{t,t+1}]=0$. Thus, the constant in the UIP regression corresponds to second-order terms in the approximation above.

rate. This means that the slope coefficient α_1 in the projection of exchange rate changes on interest rate differentials should be equal to one:

$$\frac{\Delta e^i}{e^i} = \alpha_0 + \alpha_1 \left[R^{\$}_{t,t+1} - R^{i,£}_{t+1} \right] + \varepsilon^i_{t+1}, \tag{3}$$

But in the data this slope coefficient α_1 is usually not only smaller than one, but more often than not negative. Figure 3 plots the histogram of slope coefficients - with or without a constant - for 16 currencies between the first quarter of 1971 and the fourth quarter of 2002. All of the slope coefficients are smaller than one and most 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 approaches 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 know they can make money by chasing high interest rates.

Figure 3: 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, and Vietnam.



Foreign Currency Excess Returns What does the failure of the UIP condition imply for the excess returns on these currency portfolios? If the UIP slope coefficient is below one, foreign excess returns increase with the interest rate differential. Table 1 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. The only exception is the last portfolio, which is comprised of high inflation currencies. As Bansal and Dahlquist (2000) have documented, UIP tends to work best at high inflation levels. Most surprising, however, 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 9 in the Appendix on the authors' websites), although the spread in annual excess returns drops from five to two and a half percentage points.

Table 1: Statistics for 8 Currency Portfolios

Reports the mean, standard deviation and Sharpe ratio for the real excess return on investments in foreign T-Bills for each of the eight portfolio. These portfolios were constructed by sorting currencies into ranked groups at time t based on the nominal interest rate differential with the US at the end of period $t-1$. Portfolio 1 contains 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

UIP across portfolios We can check the UIP restictions on our portfolios. At first glance, UIP seems to fit the cross-section of portfolios (results are reported in the Appendix) rather well. The slope of the regression line between the average change in exchange rate and the average interest rate differential for each portfolio is .85 if we include the eight portfolios. But, for reasons noted above, the estimated slope coefficient drops to .4 if we exclude the highest interest rate portfolios. The estimated slope coefficient drops to .2 when we run the regression for the same period exclusively using developed countries. So, the UIP appears to work well in our cross-section only when developing countries with high interest rate differentials are included. The next section introduces the class of linear factor models we focus on and explores the relation between conditional risk premia and interest rates.

2 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 shows why some might account for the forward premia. In the next section, we run a horse race between these factor models.

2.1 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:

$$m_{t+1} = b_0 + \sum_{j=1}^{n} b_j f_{j,t+1}, \tag{4}$$

where $f_{j,t+1}$ are the pricing factors. We consider two large classes of pricing models. The first class uses returns as pricing factors. In this group are the CAPM, the factor models by Fama and French (1992) and the model by Santos and Veronesi (2001). The second class of models directly introduces measures of the undiversifiable, macroeconomic risk for which investors are compensated. The Consumption-CAPM (CCAPM) and its scaled versions belong to this class. We briefly present below these two classes of linear factor models. Table 2 summarizes the factors we used.

Table 2: 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
CAPM	R^{vw}				
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}	$R^{vw}\frac{l}{c}$			
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	$\Delta(\log c_t)$	$\Delta(\log c_t)my_{t-1}$	$A_{t-1}\Delta(\log \rho_t)$	$A_{t-1}\Delta(\log \rho_t)my_{t-1}$	

Return factors The Fama-French equity pricing factors are: 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 and French (1992)). We refer to this model as the FF-CAPM for equity. Fama and French (1992) also construct two bond pricing factors. The first takes the difference between the long term government bond return and the risk free rate (R^{long}) and the second uses the spread between the return on a long-term corporate bond index and a long term government bond (R^{corp}). We refer to this model as the FF-CAPM for bonds. Fama and French (1993) argue that these factors proxy for the underlying undiversifiable macroeconomic risk.

Santos and Veronesi (2001) add a scaling variable - labor income share - to the standard CAPM. Thus, in this model, the SDF is equal to:

$$m_{t+1} = b_0 + b_1 R^{vw}_{t+1} + b_2 x_t R^{vw}_{t+1},$$

where x_t is the labor income share. An increase in the labor income share reduces the stand-in investor's exposure to equity risk, which in turn reduces the market price of risk. Santos and Veronesi (2001) show that 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.

Scaled Consumption-CAPM We consider the standard Consumption-CAPM with only aggregate consumption growth risk and the Housing-CAPM (henceforth HCAPM) proposed by Piazzesi, Schneider and Tuzel (2002). The HCAPM introduces the rental price growth risk in addition to aggregate consumption growth risk.

We also consider two different scaled versions of the consumption CAPM. To allow for time-variation we follow Lettau and Ludvigson (2001) in proposing a linearized version of the standard Breeden-Lucas stochastic discount factor:

$$m_{t+1} = b_0 + b_1 \Delta \log c_{t+1} + b_2 x_t \Delta \log c_{t+1}. \tag{5}$$

Two scaling factors x_t are considered. Lettau and 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 and Ludvigson (2001) motivate this scaling by appealing to habit formation, which Campbell and Cochrane (2000) argue will outperform the CCAPM. Lettau and 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 and Nieuwerburgh (2005) derive (5) 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: $my_{max} - my$. This makes the scaling variable is an indicator of collateral scarcity.[7] 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 and Veronesi (2001), Lettau and Ludvigson (2001) and Lustig and Nieuwerburgh (2005) in pricing domestic stock returns suggests that the Fama-French asset pricing factors do proxy for underlying macroeconomic risk. We will show that the macroeconomic factor models can price both domestic equity risk and currency risk, which the Fama-French factors cannot.

Unconditional Pricing We use the factor models described above in unconditional pricing experiments. As explained by Hansen and Richard (1987), a simple conditional factor model, for example the CCAPM in which $m^c_{t+1} = b_0 + b_1 \Delta \log c_{t+1}$, can be turned into an unconditional factor model using all the variables z_t in the information set of the investor. The conditional Euler equation $E_t \left[m_{t+1} R^i_{t+1} \right] = 1$ is then equivalent to the following unconditional condition:

$$E \left[m^c_{t+1} z_t R^i_{t+1} \right] = 1,$$

when z_t contains all the investor's information set. If we assume that the scaling variable x_t summarizes this information set, then the unconditional factor model is:

$$m_{t+1} = b_0 + b_1 \Delta \log c_{t+1} + b_2 x_t + b_3 x_t \Delta \log c_{t+1}.$$

This would introduce the scaling variable x_t itself as a pricing factor. But, since there is no compelling economic reason to expect the scaling variable risk to be priced, we decided to only add the interaction term $x_t \Delta \log c_{t+1}$ to the CCAPM.

We have set up a framework where linear factor models with time-varying coefficients can

[7] Lustig and Nieuwerburgh (2005) propose three different measures of housing collateral: one based on residential wealth $myrw$, one based on fixed assets $myfa$, and one based on outstanding mortgages, $mymo$. In this paper we used $myfa$. Lustig and Nieuwerburgh (2005) 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 and Ludvigson (2001) for the consumption-wealth ratio.

be tested on foreign currency portfolios through unconditional pricing of the investor's Euler equation. Before we present extensively the estimation results, we would like to build intuition why some of these models might successfully explain the characteristics of foreign risk premia.

2.2 Theoretical Consumption Growth Betas

The risk premia on foreign currency increase in the foreign interest rates. We work out the log-normal version of scaled CCAPM models to explain which pattern in consumption growth betas gives rise to this pattern in risk premia.

Log-normality If we assume that $\log m_{t+1}$ and $\log R^i_{t+1}$ are jointly, conditionally normal, then the Euler equation can be restated in terms of the real currency risk premium (see proof in Annex):

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

We refer to this log currency premium as $\log(crp^i_{t+1})$. It is determined by the covariance between the log of the SDF m and the real returns on investment in the foreign T-bill. Substituting the definition of this return into this equation produces the following expression for the log currency risk premium:

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

The first term on the right-hand side of the equation represents pure currency risk compensation. The second term is inflation risk compensation. Using this equation, we examine what restrictions are implied on the joint distribution of consumption growth and exchange rates by the increasing pattern of currency risk premia in interest rates.

Consumption Growth and Exchange Rates In the standard CCAPM, the log stochastic discount factor is equal to $\log m_{t+1} = \log \beta - \gamma \Delta \log c_{t+1}$. Now, we assume that the stochastic discount factor in scaled versions of the CCAPM can be approximated by:

$$\log m_{t+1} \simeq b'_0 + b'_1 \Delta \log c_{t+1} + b'_2 x_t \Delta \log c_{t+1},$$

where $b'_1 < 0$ and $b'_2 < 0$ and $x_t > 0$. In the case of the simple C-CAPM, b'_1 is the coefficient of relative risk aversion γ. The scaling variable x_t introduces time-variation in the market price of consumption growth risk.

It follows that the log currency risk premium can be restated in terms of the conditional covariance between consumption growth and the change in the exchange rate (usually known in the finance literature as consumption growth betas):

$$\log\left(crp^i_{t+1}\right) \simeq -\left(b'_1 + b'_2 x_t\right)\left[Cov_t\left(\Delta \log c_{t+1}, \Delta \log e^i_{t+1}\right) - Cov_t\left(\Delta \log c_{t+1}, \Delta \log p_{t+1}\right)\right].$$

We can abstract from the inflation compensation term $\Delta \log p_{t+1}$ because it explains none of our cross-sectional variation: it matters for the levels of the risk premia but it depends only on US characteristics (recall that x is linked to the US economic stance and p refers to the US price level) and thus does not vary across countries. This equation uncovers two mechanisms that can explain the forward premium puzzle:[8]

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

   In the data, the risk premium $\left(crp^i_{t+1}\right)$ is positively correlated with foreign interest rates $R^{i,\pounds}_{t,t+1}$: 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:

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

   Currencies that appreciate on average when US consumption growth is high and depreciate when US consumption growth is low earn positive conditional risk premia. Since interest rates predict the risk premia on foreign currency, the covariance of changes in the exchange rate with US consumption growth term needs to switch signs over time for a given currency, depending on its interest rate!

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

   If the positive conditional covariance between US consumption growth and exchange rates for low interest currencies increases in bad times for the US investor, when she demands a

[8] We are grateful to Andy Atkeson for clarifying this to us.

high risk premium for consumption growth risk, this helps to explain currency risk premia. Changes in the conditioning variable cannot explain the switch in the sign of risk premia, depending on the interest rate, but these changes in x can and do help explain the level of risk premia.

2.3 Preliminary empirical betas

This section provides some preliminary, empirical evidence in support of both of these mechanism by examining the unconditional and conditional consumption growth betas of exchange rates for each portfolio. We find that (1) the consumption growth betas of currency (or foreign cash holdings) increase from low to high interest rate currency portfolios, and (2) these consumption growth betas of high interest rate portfolios increase in bad times.

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

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

In the quarterly data, Figure 4 shows negative or small positive betas for low interest rate currencies and large positive 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. The annual data in Figure 4 show a similar pattern for the post Bretton-Woods time period.

All our results build on this basic finding. Foreign cash holdings in high interest rate currencies expose US investors to more consumption growth risk, while foreign cash holdings in low interest rate currencies provide a hedge.

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:

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

We consider the housing collateral ratio my as the scaling variable x. The sensitivity of high

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: $\Delta \log e^i_{t+1} - \Delta \log p_{t+1} = \alpha_0 + \alpha_1 \left[\Delta \log c^{US}_{t+1}\right]$. We used the post-Bretton Woods sample (1971-2002). The left panel shows the results for annual data, the right panel for quarterly data.

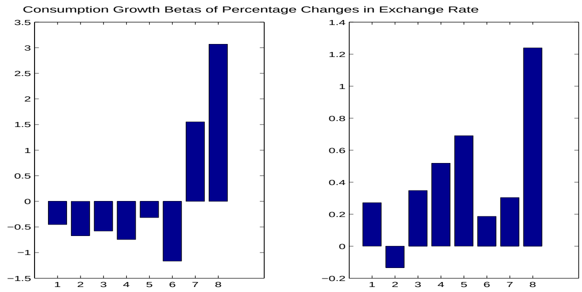
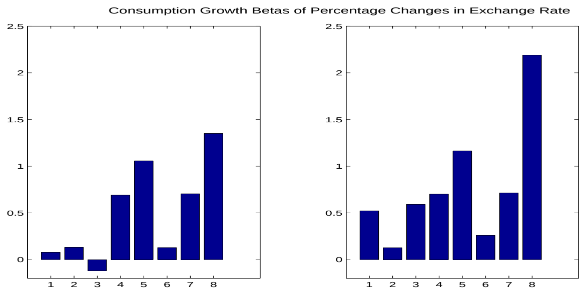


Figure 5: Consumption Growth Betas of Exchange Rates in Bad Times Estimates from regression of percentage exchange rate changes on consumption growth using 8 currency portfolios as test assets: $\Delta \log e^i_{t+1} - \Delta \log p_{t+1} = \alpha_0 + \alpha_1 \left[\Delta \log c^{US}_{t+1}\right] + x_t \alpha_2 \left[\Delta \log c^{US}_{t+1}\right]$. The consumption growth betas of exchange rates in bad times are computed as $\alpha_1 + \alpha_2 x^b$ where x^b is equal to one half standard deviation above the mean of x. The scaling variable x is the housing collateral ratio my. The left panel uses the 1953.1-2002.4 sample. The right panel uses the 1971.1-2002.4 sample. Quarterly Data.



interest rate currencies increases in bad times. Figure 16 plots the consumption growth betas evaluated at my equal to one half standard deviation above its sample mean. When housing collateral is scarce in the US, the consumption growth betas for high interest rate currencies are nearly double the unconditional values. High interest rate currencies become much riskier in bad times.

Towards A Model of Exchange Rates We have presented so far a theoretical framework that might account for some characteristics of the foreign excess returns. This framework is based on linear factor models with time-varying coefficients used in empirical finance. If we 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 the variable x_t known at date t, then it also delivers a linear factor model of exchange rates:

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

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

3 Estimation

In this section, we actually test the Euler equation of the US investor for each of these currency portfolios. Rather than commit ourselves to one model of m, we run a horse race between competing models of m. The stochastic discount factors we consider are linear combinations of return-based or consumption-based factors. Following Hansen (1982), the unconditional estimation of the linear factor models is based on the general method of moments (GMM). We normalize the SDF to $m_{t+1} = 1 - b' f_{t+1}$.[9] The moment conditions are the sample analog of the population pricing errors:

$$g_T(b) = E_T(m_t R^e_t) = E_T(R^e_t) - E_T(R^e_t f'_t) b,$$

where $R^e_t = [R^{1,e}_t \ R^{2,e}_t \ .. \ R^{N^p,e}_t]'$. In the first stage of the GMM estimation, 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: $S = \sum_{-\infty}^{\infty} E[(m_t R^e_t)(m_{t-j} R^e_{t-j})']$.[10] 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. The second stage is a GLS cross-sectional regression of average excess returns on the second moment of returns and

[9] These $b's$ have the opposite sign after this normalization.

[10] The optimal number of lags in the estimation of the spectral density matrix is determined using Andrews (1991). When pricing a large number of portfolios, this procedure is very time-consuming. Thus, we have used 4 lags on annual data and 12 lags on quarterly data when pricing 25 or 33 portfolios.

factors.[11]

The Euler equation can be rewritten as:

$$E(R^{j,e}) = -\frac{cov(m, R^{j,e})}{var(m)} \frac{var(m)}{E(m)} = \beta^j \lambda,$$

where λ is the market price of risk and β^j is the amount of risk that characterizes the excess return $R^{j,e}$. Essentially we gauge how much of the variation in average returns across portfolios can be explained by variation in the betas. If the predicted excess returns line up with the realized ones, 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.

We first test the pricing models on our foreign currency portfolios. We then introduce additional test assets to study wether currency risk is priced differently from equity and bond risk. Finally, to check that our results do not depend on the number and size of our portfolios, we test the Euler equation on each country.

3.1 Currency Portfolios as Test Assets

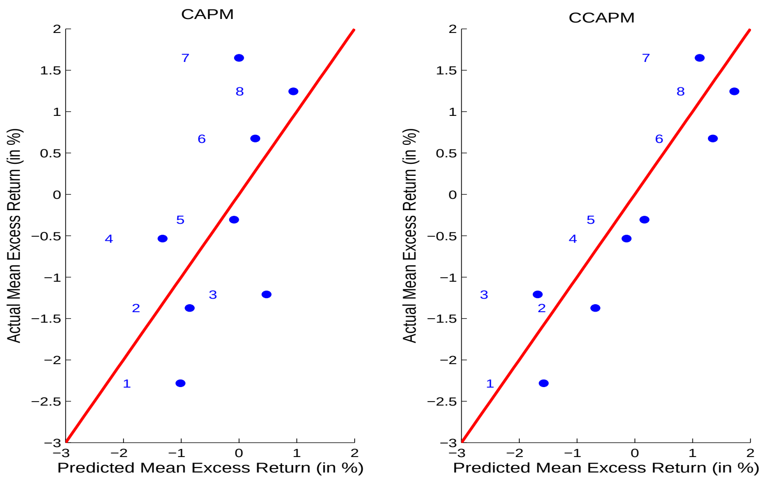
When it comes to pricing currency risk, the factors that directly measure macroeconomic, undiversifiable risk outperform the factors constructed only from asset returns. Consumption growth risk plays a key role in explaining currency risk premia.

CCAPM The standard CCAPM can explain between sixty percent and eighty percent of the cross-sectional variation in average excess returns that US investors earn on our constructed currency portfolios, sixty percent for the 1971-2003 period and eighty percent for the 1953-2002 period. The workhorse CAPM hardly explains any of the variation. This is apparent from Figure 6; it plots the actual sample average of the excess return $E_T\left[R^{j,e}_{t+1}\right]$ 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 $\beta^j_c\lambda_c$; the panel on the left plots the CAPM results with predicted excess return $\beta^j_R\lambda_R$. The variation in market betas hardly explains any of the variation in returns.

Table 3 reports the estimated market prices of risk and the p-value for the χ^2-test for the consumption-based models, while Table 4 reports the results for the factor models.

[11]Chapter 13 of Cochrane (2001) describes this estimation procedure and compares it to the one proposed by Fama and MacBeth (1973).

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.

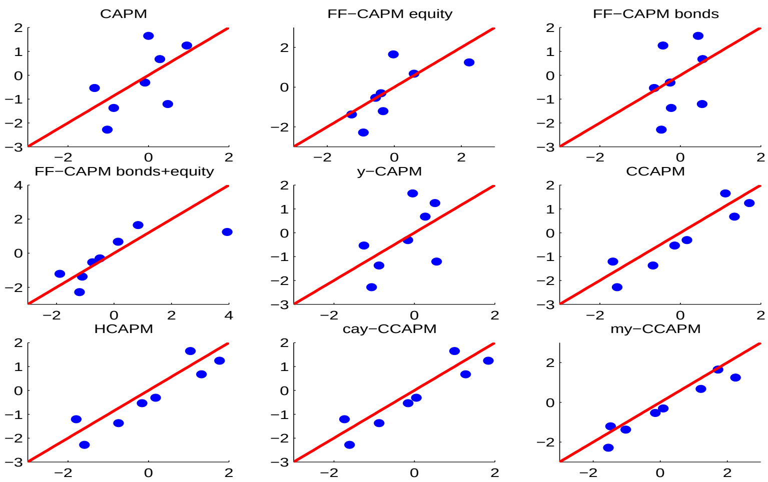


We discuss the results in Table 3 first. The estimated price of consumption growth risk λ_c is positive but not significant in all cases. λ_c is large, around five. An asset with a consumption growth beta of one yields an average risk premium of five percent. This number is similar for all of the consumption-based models, except the last one. This is a large number, but it is quite close to the market price of consumption growth risk we estimated on US equity portfolios. The implied coefficient of risk aversion in the CCAPM is around 56. This is line with stock-based estimates of the coefficient of risk aversion found in the literature.

The bottom panel of Table 3 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 each up to seventy percent. As before, the price of consumption growth risk λ_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, or 10 percent. This is much higher than the estimated consumption growth risk premium from annual data. The price of scaled consumption growth risk $\lambda_{c,x}$ is positive but not significant. [12]

[12]To assess whether individual factors have explanatory power, we also report the coefficient estimates b in a separate Appendix in Table 10.

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.



Finally, Figure 7 plots predicted against actual excess returns for 5 factor models and 4 consumption-based models. The single-factor CCAPM clearly does as well or better than the multi-factor models without consumption growth.

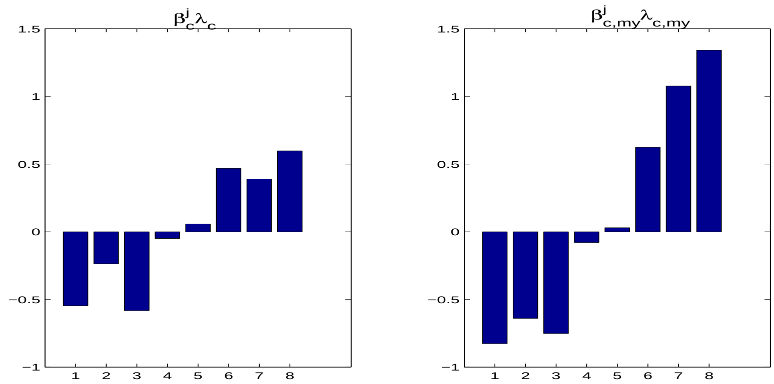
Scaled CCAPM The scaled versions of the CCAPM capture the variation in currency risk premia, especially at quarterly frequencies, because (1) the consumption growth betas of exchange rates switch signs between high and low interest rate episodes and (2) these betas increase in absolute value when the 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. In line with the theory, the estimated price of scaled consumption growth risk $\lambda_{c,x}$ is positive. This means that the price of consumption growth risk increases in bad times, when x is large.

We take the example of the housing-collateral model to show the relative importance of the two parts in the equation above. Figure 8 plots the consumption growth risk premium and the consumption-growth-collateral risk premium for each of the eight currency portfolios. For low interest rate currencies, -.60 percentage points are due to consumption growth risk and about -.75 percentage points are due to consumption-growth-collateral risk. For high interest rate currencies, .5 percentage points are due to consumption growth risk and about 1.5 percentage points are 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 10 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 mostly positive and significant for the *my*-CCAPM and *my*-HCAPM, but not always 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 (see Table 4). 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.

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 on the 1953:1-2002:4 period, but much less for the post-Bretton-Woods period. These results are in line with the ones reported in Bansal and Dahlquist (2000) who used a CAPM type of specification to price

Table 3: CCAPM Risk Price Estimates for Currency Portfolios

GMM estimates using 8 currency portfolios as test assets. The first column contains λ_c, the second column λ_x, the third column $\lambda_{c,x}$, the fourth column λ_ρ 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	λ_c	$\lambda_{c,x}$	λ_ρ	$\lambda_{\rho,x}$	*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	**4.9**				0.87	0.68	0.68
s.e.	(1.8)						
HCAPM	**4.7**	0.44			0.84	0.6	0.66
s.e.	(1.9)	(0.46)					
cay-CCAPM	**4.4**	0.97			0.84	0.62	0.67
s.e.	(1.6)	(0.36)					
my-CCAPM	**4.8**	0.53			0.83	0.61	0.66
s.e.	(1.8)	(0.22)					
my-HCAPM	**15**	1.6	1.2	0.062	0.62	0.68	0.82
s.e.	(42)	(4.3)	(4.1)	(0.39)			
			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)					
my-HCAPM	**7.4**	0.62	0.018	-0.012	0.66	0.13	0.52
s.e.	(16)	(1.4)	(0.77)	(0.11)			

28 monthly foreign excess returns over the 1976-1998 period.

Pricing Errors Clearly the consumption-based models do much better than the ad hoc factor models: on annual returns the average pricing errors for the consumption-based models are only half the size of those for the ad hoc factor models.[13] All models overpredict the risk premium on the first portfolio by at least seventy basis points, but the factor models misprice the first portfolio by at least 120 basis 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 thus our estimates.[14] To guard against this possibility, we focus only on the post Bretton-Woods sample. These results are also reported in Table 3 and Table 4. 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, which implies that the price of consumption growth risk increases in bad times.

3.2 Domestic Test Assets

Stocks 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.

[13] Table 11 in a separate Appendix reports the pricing errors for all the models we have tested.

[14] For example, Bekaert and Hodrick (1993) show that a test of the UIP condition that allows for endogenous shifts in the parameters leads to more severe deviations from unbiasedness in the 1980s.

Table 4: 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	λ_1	λ_2	λ_3	λ_4	λ_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)						

Table 5: 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 λ_c, the second column λ_x, the third column $\lambda_{c,x}$, the fourth column λ_ρ 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 4 lags to estimate the spectral density matrix with annual data and 12 lags with quarterly data.

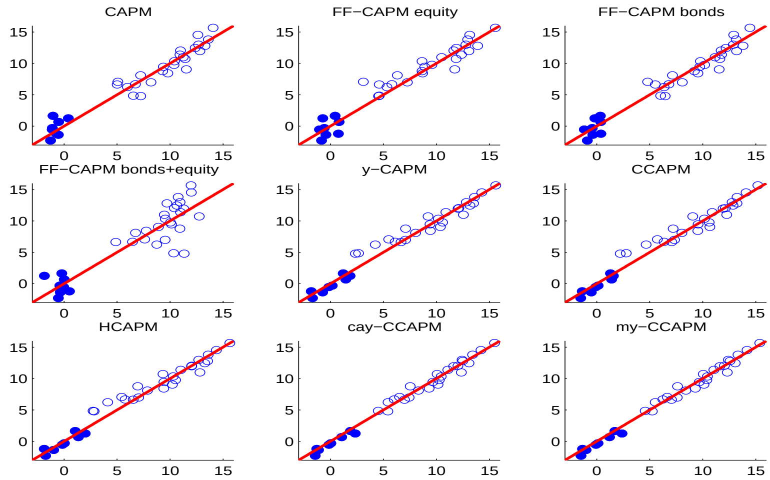
Model	λ_c	$\lambda_{c,x}$	λ_ρ	$\lambda_{\rho,x}$	*p value*	R^2_{adj}	R^2
Panel A: Annual Data							
1953-2002							
CCAPM.	**8**				1	0.96	0.96
s.e.	(1.1)						
HCAPM.	**5.5**		0.73		1	0.96	0.96
s.e.	(0.5)		(0.07)				
cay-HCAPM.	**8.1**	1.7			1	0.96	0.96
s.e.	(0.93)	(0.2)					
my-CCAPM.	**5.8**	1			1	0.99	0.99
s.e.	(0.84)	(0.13)					
my-HCAPM.	**5.6**	0.9	0.43	0.077	1	0.98	0.99
s.e.	(0.79)	(0.13)	(0.11)	(0.018)			
1971-2002							
CCAPM.	**5.8**				1	0.87	0.87
s.e.	(0.15)						
HCAPM.	**2.9**		0.76		1	0.93	0.93
s.e.	(0.023)		(0.0098)				
cay-CCAPM.	**5**	1.1			1	0.88	0.88
s.e.	(0.075)	(0.015)					
my-CCAPM.	**4.5**	0.64			1	0.92	0.92
s.e.	(0.22)	(0.038)					
my-HCAPM.	**3.3**	0.29	0.51	0.077	1	0.95	0.96
s.e.	(0.056)	(0.0014)	(0.011)	(0.0011)			
Panel B: Quarterly Data							
1953.1-2002.4							
CCAPM	**1.2**				0.99	0.94	0.94
s.e.	(0.13)						
HCAPM	**1.2**		0.14		0.99	0.94	0.95
s.e.	(0.14)		(0.023)				
cay-CCAPM	**1.2**	0.3			0.99	0.94	0.94
s.e.	(0.15)	(0.037)					
my-CCAPM	**1.6**	0.11			0.99	0.96	0.96
s.e.	(0.21)	(0.018)					
my-HCAPM	**1.7**	0.12	0.24	0.016	0.99	0.96	0.96
s.e.	(0.2)	(0.018)	(0.035)	(0.0041)			
1971.1-2002.4							
CCAPM	**1.1**				1	0.88	0.88
s.e.	(0.084)						
HCAPM	**1.2**		0.11		1	0.91	0.91
s.e.	(0.12)		(0.027)				
cay-CCAPM	**1.2**	0.34			1	0.88	0.88
s.e.	(0.099)	(0.028)					
my-CCAPM	**2**	0.14			1	0.93	0.93
s.e.	(0.23)	(0.022)					
my-HCAPM	**1.9**	0.16	0.25	0.017	1	0.96	0.96
s.e.	(0.25)	(0.032)	(0.045)	(0.005)			

Table 6: 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 4 lags to estimate the spectral density matrix with annual data and 12 lags with quarterly data.

Model	λ_1	λ_2	λ_3	λ_4	λ_5	*p value*	R^2_{adj}	R^2
Panel A: Annual Data								
1953-2002								
CAPM.	9.3					1	0.75	0.75
s.e.	(0.73)							
FF-CAPM e.	6.7	2.3	6.3			1	0.94	0.94
s.e.	(0.84)	(0.8)	(0.5)					
FF-CAPM b.				6.1	4	1	0.93	0.93
s.e.				(0.89)	(0.42)			
FF-CAPM b.+e.	7.1	2.4	6.4	6.2	1	1	0.95	0.95
s.e.	(1)	(1)	(0.91)	(0.73)	(0.32)			
y-CAPM.	6.5	7.6				1	0.79	0.8
s.e.	(0.9)	(0.98)						
1971-2002								
CAPM	12					1	0.25	0.25
s.e.	(0.71)							
FF-CAPM e.	5.7	1.6	6			1	0.9	0.91
s.e.	(0.31)	(0.28)	(0.34)					
FF-CAPM b.				8	2	1	0.95	0.95
s.e.				(0.053)	(0.03)			
FF-CAPM e+b.	6.2	1.7	6.5	9.3	1.1	1	0.96	0.96
s.e.	(0.25)	(0.24)	(0.26)	(0.31)	(0.069)			
y-CAPM	4.5	5.3				0.75	0.35	0.37
s.e.	(1.1)	(1.1)						
Panel B: Quarterly Data								
1953.1-2002.4								
CAPM	2.1					0.99	0.61	0.61
s.e.	(0.32)							
FF-CAPM e.	1.7	0.47	1.6			0.99	0.9	0.91
s.e.	(0.35)	(0.19)	(0.29)					
FF-CAPM b.				2.2	1.3	0.99	0.89	0.89
s.e.				(0.51)	(0.31)			
FF-CAPM e.+b.	2.1	0.71	1.5	1.5	3.1	0.99	0.9	0.91
s.e.	(0.61)	(0.29)	(0.38)	(0.67)	(0.5)			
y-CAPM	1.7	3				0.99	0.68	0.69
s.e.	(0.62)	(0.91)						
1971.1-2002.4								
CAPM	1.9					1	0.42	0.42
s.e.	(0.34)							
FF-CAPM e.	1.6	0.49	1.6			1	0.86	0.86
s.e.	(0.45)	(0.23)	(0.33)					
FF-CAPM b.				2.4	0.081	1	0.82	0.82
s.e.				(0.35)	(0.16)			
FF-CAPM e.+b.	1.5	0.73	1.7	1.4	0.95	1	0.9	0.91
s.e.	(0.47)	(0.25)	(0.43)	(0.42)	(0.24)			
y-CAPM	1.3	2.5				1	0.7	0.71
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.

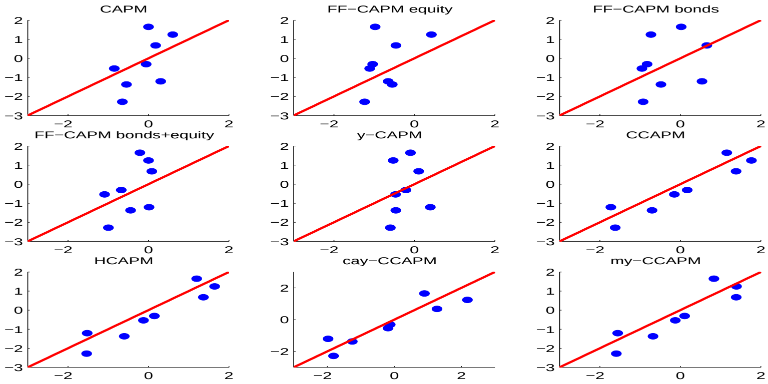


The three factor model for stock pricing developed by Fama and French (1993) fails to price both the equity risk and the currency risk. The two factor model for bonds does much better, but all five of these factors are needed to get a reasonably close match between predicted and actual excess returns. The models that explicitly introduce macroeconomic risk invariably do much better at pricing both currency risk and equity risk in a parsimonious way. The 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.[15]. 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 $b_{cay,c}$. [16]

[15]Pricing errors for these 25 portfolios are reported in a separate Appendix in Table 12 and Table 13

[16]Results are reported in a separate Appendix in Table 14.

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.



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. Pricing errors are reported in Table in the separate appendix.

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 15 shows 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 shows the predicted and realized foreign excess returns when the risks' prices are estimated on US domestic band and equity portfolios. The figure confirms the result above, even though the pricing errors are quite large.

3.3 Managed Portfolios as Test Assets

To check that our results do not depend on any undesirable feature of our portfolios, we also test the investor's Euler equation on each currency. Instead of using portfolios that change composition each period, we use the nominal interest rate differential itself as an instrument z_t:

$$E_t\left[m_{t+1}\left(R^{i,\$}_{t+1}\frac{p_t}{p_{t+1}} - R^{US,\$}_{t,t+1}\frac{p_t}{p_{t+1}}\right)\right]z^i_t = 0, \tag{6}$$

where z_t^i is the interest rate differential between the foreign T-bill and its US equivalent $R_{t,t+1}^{i,*} - R_{t,t+1}^{f,\$}$. This procedure is equivalent to the pricing of excess returns on managed portfolios. For country i the return on its managed portfolio $\widetilde{R}_t^{e,i}$ is given by:

$$\widetilde{R}_t^{e,i} = \left[R_{t+1}^{i,£} \frac{p_t}{p_{t+1}} - R_{t,t+1}^{US,\$} \frac{p_t}{p_{t+1}} \right] \times z_t^i. \tag{7}$$

These are 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. Using the instrumental variable z_t^i is equivalent to pricing unconditionally these managed returns.

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 7 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 than the risk premia due to relative movements in one currency's interest rate, compared to all the others. Still, λ_c and $\lambda_{c,x}$ are positive, significant and have the right sign. These estimates are in line with our previous ones.

4 Where do Exchange Rate Betas come from?

We have shown that the cross-section of currency risk premia is well explained by consumption growth risk. But what gives rise to the monotonic relation between the consumption growth betas of exchange rates and interest rates in the data? We show in this section that the key is (1) negative correlation between the interest rate and the second moment of the foreign stochastic discount factor, and/or (2) a higher correlation of the SDF in low interest rate currencies with their US counterpart. To obtain these results, we assume that markets are complete and that SDFs are log-normal. Essentially, we re-interpret an existing derivation by Backus et al. (2001), and we explore its empirical implications.

4.1 Exchange Rate Betas

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

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

Table 7: CCAPM Risk Price Estimates for Managed Currency Portfolios

GMM estimates using 7 managed currency portfolios as test assets in the first sample and 11 in the second sample. The data are annual. The collateral measure for the *my*-CCAPM and *my*-HCAPM is $myfa$. We used 4 lags to estimate the spectral density matrix.

Model	λ_c	$\lambda_{c,x}$	λ_ρ	$\lambda_{\rho,x}$		*p value*	R^2_{adj}	R^2
Panel A: Consumption-based Models								
1953-2002								
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)				
1971-2002								
CCAPM.	4.9					0.69	−0.16	−0.16
s.e.	(1.5)							
HCAPM.	9		0.32			0.64	−0.25	−0.043
s.e.	(13)		(0.82)					
cay-CCAPM.	3	0.7				0.6	−0.32	−0.1
s.e.	(1.3)	(0.28)						
my-CCAPM.	9.1	2.1				0.53	0.5	0.58
s.e.	(13)	(2.9)						
my-HCAPM.	2.2	0.54	0.093	0.078		0.37	0.21	0.61
s.e.	(2.6)	(0.6)	(0.33)	(0.056)				
Panel B: Factor Models								
1953-2002								
CAPM	−14					0.93	−3.7	−3.7
s.e.	(3.2)							
FF-CAPM e.	−16	−8.9	10			0.92	−5.3	−4
s.e.	(5.9)	(5.5)	(5.6)					
FF-CAPM b.				−6.1	23	0.9	−1.9	−1.6
s.e.				(7.5)	(14)			
FF-CAPM e.+b.	−15	−0.93	14	2.9	4.9	0.66	−1.8	−0.65
s.e.	(12)	(8.8)	(10)	(7.7)	(1.7)			
y-CAPM	−15	−17				0.79	−2.5	−2.2
s.e.	(5.9)	(6.4)						
1971-2002								
CAPM	−11					0.38	0	0
s.e.	(2.9)							
FF-CAPM e.	−9.8	−5.1	5.6			0.36	−4.2	−4.2
s.e.	(4.1)	(14)	(7.1)					
FF-CAPM b				−15	41	0.77	−0.93	−0.61
s.e.				(14)	(29)			
FF-CAPM e+b	40	36	−5.8	−23	26	0.52	−1.3	0.23
s.e.	(36)	(49)	(26)	(16)	(23)			
y-CAPM	−11	−12				0.31	−5.1	−4.1
s.e.	(9.5)	(11)						

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. Assuming that markets are complete, we substitute the difference between the log stochastic discount factors for the change in the real exchange rate:

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

What do we learn from this? Assuming that the inflation betas are small enough, the sign of the log risk premium is determined by the standard deviation of the home SDF relative to the one of the foreign SDF scaled by the correlation between the two SDFs:

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

Heteroskedasticity First, suppose that the correlation between the SDF's is positive and constant. If countries characterized by a high interest rate $R^{i,£}_{t,t+1}$, typically also have a low conditional volatility of the foreign SDF m^i_{t+1} relative to its domestic counterpart, then the sign of the expression above will be positive. Thus high interest countries will deliver positive currency risk premia. Conversely, if low interest rate currencies are characterized by a high conditional volatility of the SDF, then the sign of the expression above will be negative. This mechanism switches the sign of exchange rate betas between high and low nominal interest rate countries.

To understand this result, recall that the real exchange rate appreciates if the foreign SDF, or the state price of a unit consumption, is higher than the domestic state price of consumption. Now, if the foreign SDF is positively correlated with the domestic one, and if it is highly volatile, then it provides a hedge for the domestic investor against bad, high marginal utility growth, states! Suppose they are perfectly correlated, but the domestic SDF is only half as volatile. Consider the case in which the domestic SDF is 5 percent above its mean. The foreign SDF is 10 percent above its mean, and the real exchange rate appreciates by 5 percent. So, investing in this foreign currency provides a perfect hedge for a US investor.

This behavior is at the heart of the habit-based model of the exchange rate risk premium in Verdelhan (2004a). In this model, the domestic investor receives a positive exchange rate risk premium in times when he is more risk-averse than his foreign counterpart. Times of high risk-aversion correspond to low interest rates. Thus, the domestic investor receives a positive risk premium when interest rates are lower at home than abroad. A nonlinear estimation of the model using consumption data leads to reasonable parameters when pricing the foreign excess returns of an American investor. The evidence from currency markets suggests that low interest rates signal an increase in the conditional market price of risk.

Correlation We focus in this paper on a second mechanism, supposing that the conditional volatilities of the SDFs are constant. If the conditional correlation of the SDFs is positive for low-interest countries and negative for high-interest rate countries, then this conditional correlation might account for the cross-section of risk premia. Thus, we want to test wether the conditional correlation of the SDFs decreases with the interest rate differential.

4.2 Consumption Co-movements and Interest Rates

We start by considering the case of the Consumption-CAPM and we assume that all countries share the same coefficient of relative risk aversion. Abstracting from the inflation betas, the sign of the conditional risk premium is determined by:

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

A high correlation of foreign consumption growth with US consumption growth for low interest rate currencies can imply negative risk premia. What is the economic intuition behind this mechanism? If the consumption process of a high interest rate country presents a low correlation with US consumption, then a negative consumption shock in the US leads to an appreciation of the dollar and a lower foreign return. This currency depreciates in bad times for the US investor, who thus wants to be compensated for that risk. The data seem to support this time-varying correlation mechanism.

Using a sample of ten developed countries[17], we regressed a country's consumption growth on US consumption growth and US consumption growth interacted with the lagged interest rate differential:

$$\Delta \log c_{t+1}^{i} = \alpha_0 + \alpha_1 \Delta \log c_{t+1}^{US} + \alpha_2 \left(R_{t,t+1}^{i,\pounds} - R_{t,t+1}^{\$}\right) \Delta \log c_{t+1}^{US} + \epsilon_{t+1}.$$

The results obtained over the post-Bretton Woods period are reported in Table 8. On annual data, the coefficients on the interaction terms α_2 are negative for all countries, except for Japan. On quarterly data, the coefficients α_2 are negative for all countries, except for Japan and the Netherlands. The table also reports ninety percent confidence intervals for these interaction

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

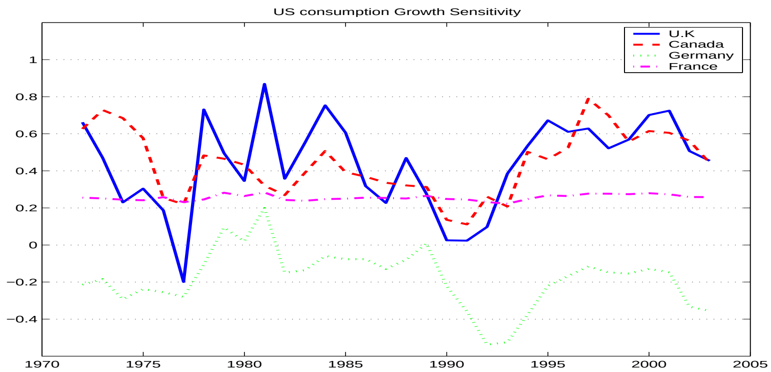
coefficients. They show that the α_2 coefficients are significatively negative for 7 countries of the annual sample (6 countries on quarterly data). The last row of each panel reports the pooled time series regression results. The ninety percent confidence interval includes only negative coefficients on both annual and quarterly samples. These estimates are economically significant as well. As shown in Figure 11, 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 8: Consumption Growth Regressions

Results for the following time-series regression: $\Delta \log c^i_{t+1} = \alpha_0 + \alpha_1 \Delta \log c^{US}_{t+1} + \alpha_2 \left(R^{£}_{t,t+1} - R^{\$}_{t,t+1} \right) \Delta \log c^{US}_{t+1} + \epsilon_{t+1}$. 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 α_2. Sample covers the post Bretton Woods period (or shorter periods when data are not available in 1971). Consumption growth rates are in percentage and interest rate differentials in basis points.

Country	α_0	α_1	α_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: $\alpha_1 + \alpha_2 \left(R^{£}_{t,t+1} - R^{\$}_{t,t+1} \right)$. The plot shows the UK (full line), Canada (dashed line), Germany (dotted line) and France (dash-dot line).



5 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 that includes US consumption growth as one of its factors. These models even perform well when the test assets also include domestic equity returns. The factor models constructed explicitly to price domestic equity and bond risk break down when confronted with currency risk. These results lead us to believe that macro-economists 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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.1 Proofs

.1.1 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:

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

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

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

which produces the following expression for the risk premium:

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

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 $Cov_t \left(\Delta \log c_{t+1}, \Delta \log e^i_{t+1}\right)$ 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:

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

and/or

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

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.

.1.2 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 $std_t \log m^{£,i}_{t+1}$.

We can derive a similar condition using the real SDF m_{t+1} or m^i_{t+1} instead of the nominal SDF $m^{£,i}_{t+1}$ or $m^{\$}_{t+1}$. The log currency risk premium is

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

This is equivalent to:

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

Assume $Cov_t\left(\log m_{t+1}, \Delta \log p^i_{t+1}\right) = 0$ and substitute the log difference in the SDF's for the change in the real exchange rate. This produces an equivalent condition in terms of the real SDF:

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

.2 Data

.2.1 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).

.2.2 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).

.2.3 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.

.2.4 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.

.2.5 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 and Nieuwerburgh (2005)). 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 and Ludvigson (2001)). The data are available from Martin Lettau's web site at *pages.stern.nyu.edu/∼mlettau/.*