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

Hanno Lustig and Adrien Verdelhan *
UCLA/NBER and Boston University

August 2005

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 currency returns into portfolios based on foreign interest rates, and we test the Euler equation of a domestic investor who invests in these currency portfolios. We find that domestic investors earn negative excess returns on low interest rate currency portfolios and positive excess returns on high interest rate currency portfolios. Because high interest rate currencies depreciate on average when domestic consumption growth is low and low interest rate currencies do not under the same conditions, low interest rate currencies provide domestic investors with a hedge against domestic aggregate consumption growth risk.

Keywords: Exchange Rates, Asset Pricing.

*First version November 2003. The authors especially thank Andy Atkeson, John Cochrane, Lars Hansen and Anil Kashyap for detailed comments, as well as Ravi Bansal, Hal Cole, Virgine Coudert, François Gourio, John Heaton, Patrick Kehoe, Lee Ohanian, Fabrizio Perri, Stijn Van Nieuwerburgh and the participants of the Chicago GSB macro lunch, the macro seminar at Chicago, Boston University, the Federal Reserve Board, Harvard University, HBS, HEI, LBS, the University of Maryland, UCLA, UCSD, the Fuqua School of Business at Duke, Stanford University, the Federal Reserve Bank of New York, the Federal Reserve Bank of Minneapolis, the Bank of France, the 2004 Annual European Meeting of the Econometric Society, and the 2004 Cleveland Fed conference on International Economics, for helpful suggestions. Dennis Quinn and Carmen Reinhardt generously shared their data with us. Stijn Van Nieuwerburgh shared his matlab code and econometric expertise with us. We would also like to thank John Campbell, Kenneth French, Martin Lettau and Motohiro Yogo for making data available on-line. This project was started when Lustig was at the University of Chicago.

When the foreign interest rate is higher than the US interest rate, risk-neutral and rational US investors should expect the foreign currency to depreciate against the dollar by the difference between the two interest rates. This way, borrowing at home and lending abroad or vice-versa produces a zero excess return. This is known as the uncovered interest rate parity (UIP) condition, and it is violated in the data[1]. What the data tell us, is that higher foreign interest rates almost always predict higher excess returns for a US investor in foreign currency markets.

We show that these excess returns compensate the US investor for taking on more US consumption growth risk. High foreign interest rate currencies on average depreciate against the dollar when US consumption growth is low, while low foreign interest rate currencies do not. The textbook logic we use for any other asset can be applied to exchange rates, and it works. If an asset offers low returns when the investor's consumption growth is low, it is risky, and the investor wants to be compensated through a positive excess return.

Currency Portfolios To uncover the link between exchange rates and consumption growth, we build portfolios of foreign currencies excess returns on the basis of the foreign interest rates, because investors know these predict excess returns.[2] Portfolios are re-balanced every period, so the first portfolio always contains the lowest interest rate currencies and the last portfolio always contains the highest interest rate currencies. This is the key innovation in our paper. Building these foreign currency portfolios serves three purposes. First, this method enables us to study the conditional correlation between consumption growth and exchange rate, where the conditioning information is here summarized by the interest rate differential. We find a significant link between US consumption growth and each portfolio's average exchange rate change, even though there is no similar relation between exchange rate changes for a particular currency and US consumption growth. Second, it allows us to keep the number of covariances that must be estimated low, while allowing us to continuously expand the number of countries studied as financial markets open up to international investors. This enables us to include data from the largest possible set of countries. Third, it isolates the source of variation that interests us, and it creates a large average spread of up to five hundred basis points between low and high

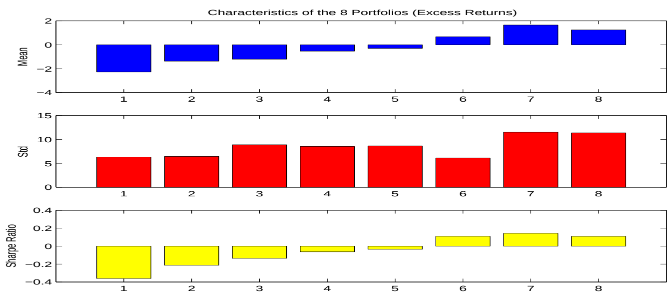
[1]The UIP condition implies that the slope in a regression of the change in the exchange rate on the interest rate differential is equal to one, and the data consistently produce coefficients less than one, and very often negative (Hansen and Hodrick (1980) and Fama (1984)). Hodrick (1987), Lewis (1995), and Verdelhan (2004b) provide extensive surveys and updated regression results.

[2]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. Engel and West (2005) argue this lack of predictability is consistent with a model in which the fundamentals are I(1) and the discount factor is close to 1. One exception is the work by Gourinchas and Rey (2003). By manipulating the country budget constraint, they argue that a measure of current account imbalance predicts returns on US assets held by foreigners, and hence exchange rates, but their predictor is the same across countries and can not be used to sort currencies into portfolios. Thus, we are less likely to miss important information in the investor's information set by conditioning only on interest rate differentials.

interest rate portfolios. This spread is an order of magnitude larger than the average for any two given countries.

These currency portfolios deliver a stable pattern in excess returns. We work with eight portfolios. As one would expect from the literature on the UIP condition, US investors earn on average 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. The same pattern is obtained when the same portfolio building exercise is repeated for 10 other developed countries.

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.



Factor Models To show that the excess returns on these portfolios are due to currency risk, we start from the US investor's Euler equation. Instead of committing to a single specification of the stochastic discount factor (or inter-temporal marginal rate of substitution) m, we run a horse race between a large cross section of models. We consider two large classes of pricing models. The first class uses returns as pricing factors. For this class, we draw on the Capital Asset Pricing Model (*CAPM*), the equity and bond factors proposed by Fama and French (1992) and the conditional *CAPM* derived from an equilibrium model by Santos and Veronesi (2005). The second class uses aggregate consumption growth as the main pricing factor. We consider different extensions of the Consumption-*CAPM* (*CCAPM*) developed by Yogo (2005), Piazzesi, Schneider and Tuzel (2002) and Parker and Julliard (2005). In addition, we bring in conditioning information, along the lines suggested by Lettau and Ludvigson (2001) and Lettau and Ludvigson (2005), and Lustig and Nieuwerburgh (2005a), to introduce potential time-variation in risk premia (see Cochrane (2005) for an overview of this literature).

We test the US investor's Euler equation in two ways. First, we minimize the pricing errors on eight 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. In this case, we use the nominal interest rate differential itself as an instrument.

This procedure is equivalent to the pricing of excess returns on managed portfolios that move in and out of a particular currency depending on the interest rate. In the paper, we report results obtained through the first method (GMM) on annual and quarterly data for the periods 1953-2002 and 1971-2002 and through the second method (managed portfolios) on annual data for the same two periods.

Main Results At annual frequencies, the Consumption-*CAPM* explains up to eighty percent of the variation in currency excess returns across these eight currency portfolios. At quarterly frequencies, Yogo's extension of the standard Consumption-*CAPM* to durables explains more than eighty percent of the variation in average returns. The estimated coefficient of risk aversion is around 50 for the Consumption-*CAPM*, and the estimated price of aggregate consumption growth risk is about five percent per annum. If we estimate the models using only US domestic stock portfolios (sorted by book-to-market and size) and US domestic bonds, we can still explain some of the variation in currency excess returns.

In addition, we test the Euler equation for an investor in each of 10 other developed economies. The standard Consumption-*CAPM* explains up to eighty percent of the variation in excess returns on these currency portfolios if we pool all the observations on developed countries.

Consumption-based models can explain the cross-section of currency excess returns if and only if high interest rate currencies typically depreciate when real US consumption growth is low, while low interest rate currencies appreciate, and that is exactly the pattern we find in the data. We can restate this result in standard finance language using the consumption growth beta of a currency. The consumption growth beta of a currency measures the sensitivity of the dollar return on cash holdings of foreign currency to changes in US consumption growth. These betas are negative for low interest rate currencies and positive for high interest rate currencies, and the spread between betas increases in bad times. All our results work off this basic finding.

Economic intuition From our vantage point, the UIP puzzle looks like a standard asset pricing puzzle. Now, where do these exchange rate betas come from and why are nominal interest rates correlated with betas? The key is time-variation in the conditional distribution of the foreign stochastic discount factor $m^{\star}$. We identify two potential mechanisms. Low foreign interest rates either signal (1) an increase in the volatility of the foreign stochastic discount factors or (2) an increase in the correlation of the foreign stochastic discount factor with the domestic one.

What is the economics behind the first mechanism? In our benchmark representative agent model with complete markets, the foreign currency appreciates when foreign consumption growth is lower than US aggregate consumption growth and depreciates when it is higher.[3] If the foreign stand-in agent's consumption growth is strongly correlated with

[3]When markets are complete, the value of a dollar delivered tomorrow in each state of the world, in

and more volatile than that of his US counterpart, his national currency provides a hedge for the US representative agent. For example, consider a representative agent with power utility preferences and risk aversion coefficient γ in a situation where foreign consumption growth is twice as volatile as US consumption growth, and perfectly correlated with US consumption growth. In this case, 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: it appreciates when US consumption growth is low.[4] Consequently, investing in this currency should provide a low excess return. Thus, for this mechanism to explain the pattern in currency excess returns, low interest rate currencies must have aggregate consumption growth processes that are conditionally more volatile than US aggregate consumption growth. An increase in the conditional volatility of aggregate consumption growth lowers the real risk-free rate in our benchmark model. If the real and nominal rates move in lockstep, that might account for part of the pattern in the consumption betas of exchange rates. We know interest rates are informative about risk, because interest rates predict stock returns and bond returns.

We identify time-variation in this correlation as the second mechanism. In the previous example, if the consumption growth of a high interest rate country is perfectly negatively correlated with US consumption growth, then a negative consumption shock of 2 percent in the US leads to a depreciation of the foreign currency by 2 percent. This currency depreciates when US consumption growth is low. Consequently, investing in this currency should provide a high excess return. Thus, for this mechanism to explain the pattern in currency excess returns, the correlation between domestic and foreign consumption growth should decrease with the interest rate differential. Empirically, we find strong evidence to support that mechanism: foreign consumption growth is less correlated with US consumption growth when the foreign interest rate is high. In a sample of 10 developed countries, the conditional correlation between foreign and US annual consumption growth decreases with the interest rate gap for all countries except Japan. We document the same pattern for Japanese and UK consumption growth processes.

Related Literature Our paper draws on at least two strands of the exchange rate literature. First, there is a large literature on the efficiency of foreign exchange markets. 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. This is known as the forward premium puzzle. Fama (1984) argues that time-varying-risk

terms of dollars today, equals the value of a unit of foreign currency tomorrow delivered in the same state, in units of currency today: $e_{t+1}/e_t = m^{\star}_{t+1}/m_{t+1}$, where the exchange rate e is in dollars per foreign currency and a star denotes a foreign variable. Thus, if investors are characterized by their constant relative risk aversion coefficient γ, then $e_{t+1}/e_t = (\Delta c_{t+1}/\Delta c^{\star}_{t+1})^{\gamma}$.

[4] Note that when consumption growth is +2 percent in the US, it is twice as high abroad (+4 percent), and the real exchange rate depreciates by γ times 2 percent. This currency is again a perfect hedge against US aggregate consumption growth risk: it depreciates when US consumption growth is high.

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 risk-based explanations[5]. Our paper is the first to show that the excess returns predicted by asset pricing's standard, real factor models that include aggregate consumption growth as a key factor, line up with the predictable excess returns in currency markets.

Other authors have pursued the risk premium explanation. Our paper is closest to Hollifield and Yaron (2001), Harvey, Solnik and Zhou (2002) and Sarkissian (2003). Hollifield and Yaron (2001) find some evidence that real factors, not nominal ones, 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 that is related to foreign exchange risk. Sarkissian (2003) finds that the cross-sectional variance of consumption growth across countries helps explain currency risk premia, but he focuses on unconditional moments of currency risk premia on a currency-by-currency basis, while we know that most of the variation depends on the level of the foreign interest rate. Finally, Backus, Foresi and Telmer (2001) show that, in a general class of affine models, explaining the forward premium puzzle requires the state variables to have asymmetric effects on the state prices in different currencies. We reinterpret their results in our framework, explaining the relation between interest rates and the consumption growth betas of exchange rates.

There is another literature that relates the volatility and persistence of real exchange rates to aggregate consumption. Standard, dynamic equilibrium models, imply a strong link between consumption ratios and the real exchange rate, but, as Backus and Smith (1993) point out, 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 real exchange rates 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.

Our results provide guidance for applied theoretical work in this area. A good theory of real exchange rates needs to explain why (nominal) interest rates line up with a currency's aggregate consumption growth betas. And it must explain why this relation breaks down

[5] Froot and Thaler (1990) conclude their survey of this literature as follows:

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

for very high interest rates. At least on the first count, our results provide empirical support for work by Verdelhan (2004a). He replicates the forward discount bias in a model with external habits and he provides estimates to support this mechanism.

Finally, we also contribute to the empirical asset pricing literature on the measurement of the marginal utility of wealth by testing a whole battery of pricing models on a completely different set of test assets. The results are unambiguous. Only consumption-based models price currency risk. This provides additional support to recent evidence that news about the demise of the *CCAPM* was premature.[6]

The first section outlines our empirical framework and defines the foreign currency excess returns and the potential pricing factors. The second section presents the asset pricing results obtained on our foreign currency portfolios, focusing on the US investor's perspective. The third section checks the robustness of our results in various ways and extends them to investors in other developed economies. The fourth section details the economic mechanisms at the core of our results. A separate appendix with auxiliary estimation results, data (including the composition of the currency portfolios) and programs is available on the authors' web sites.[7]

1 Framework

This section defines the excess returns on foreign T-bill investments and derives the Euler equation for a US investor. We describe our data set, we explain how we construct the currency portfolios and we present the potential pricing factors.

1.1 Definitions and data set

We first 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. $R^{i,\$}_{t+1}$ 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: $R^{i,\$}_{t+1} = R^{i,£}_{t,t+1}\frac{e^i_{t+1}}{e^i_t}$ where e^i_t is the exchange rate in dollar per unit of foreign currency and $R^{i,£}_{t,t+1}$ is the *risk-free* one-period return in units of foreign currency i. $R^{\$}_{t,t+1}$ is the nominal risk-free rate in US currency, while $R_{t,t+1}$ is the risk-free rate in units of US consumption.

[6] A standard *CCAPM* (using fourth quarter to fourth quarter non-durable consumption and services growth) explains the 25 Fama-French portfolios, according to Jagannathan and Wang (2005), while Parker and Julliard (2005) demonstrate that long-run measures of consumption risk do much better in explaining the cross section of stock returns. More recently, Lustig and Nieuwerburgh (2005b) back out a new measure of the return on the total market portfolio from aggregate consumption data, and they argue that the true market return is not correlated with stock market returns. Stock market risk is a poor measure of market risk. This could explain why Lewellen and Nagel (2003) conclude that there is not enough variation in conditional betas to explain stock returns.

[7] See

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:

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

where R^i_{t+1} denotes the *random* return in units of US consumption from investing in T-bills of currency i: $R^i_{t+1} = R^{i,\$}_{t+1}\frac{p_t}{p_{t+1}}$, and p_t is the dollar price of a unit of the US consumption basket.

Unconditional Pricing The conditional Euler equation $E_t\left[m_{t+1}R^i_{t+1}\right] = 1$ implies the following unconditional condition version:

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

where z_t contains the investor's entire information set. We can read the equation above as an unconditional pricing experiment of managed excess returns $z_tR^i_{t+1}$, where the currencies will be weighted according to the useful available information z_t. Fortunately, we know from Meese and Rogoff (1983) that our ability to forecast exchange rates is rather limited. Thus, by building our portfolios on the basis of the interest rate differential, we might have already used all the useful information available to the investor at the time of her decision. We focus on the currency portfolio returns.

Currency Portfolios To better analyze the risk-return trade-off for a US investor investing in foreign currency markets, we construct currency portfolios that zoom in on the effect we are after, the predictability of excess returns by foreign interest rates.

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. Portfolios are rebalanced every quarter when we work on quarterly data and every year when we use annual data. Low interest rate differential portfolios and high interest rate differential portfolios are ranked from 1 to N^p. We compute dollar returns of foreign T-bill investments $R^{j,\$}_{t+1}$ for each portfolio j by averaging across the different countries in each portfolio. 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 currencies $E_T\left[R^{i,e}_{t+1}\right], i = 1,\ldots,N^c$, because foreign interest rates fluctuate: the foreign excess return is positive (negative) when foreign interest rates are high (low), and periods of high excess returns are cancelled out by periods of low excess returns.

This US investor's currency portfolio Euler equation for excess returns is the focus of the rest of this paper:

$$E_t\left[m_{t+1}\left(R^{j,\$}_{t+1}\frac{p_t}{p_{t+1}}-R^{\$}_{t,t+1}\frac{p_t}{p_{t+1}}\right)\right]=0, j=1\dots N^p \quad (2)$$

Sample We always use a total number of eight portfolios. Given the limited number of countries, especially at the start of the sample, 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. 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 expected and actual default events, and the effects of financial liberalization.

Interest Rate The foreign interest rate is the interest rate on a 3-month government security (e.g. a US T-bill) or an equivalent instrument. When using annual data, we used the 3-month interest rate instead of the one-year rate, simply because fewer countries issue bills or trade equivalent instruments 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.

Default Defaults can impact our currency returns in two ways. First, expected defaults should lead rational investors to ask for a default premium, thus increasing the foreign interest rate and the foreign currency return. To check that our results are due to currency risk, we run and report all experiments for a sub-sample of developed countries.[8] None of these countries has ever defaulted, nor was ever considered likely to. Yet, we obtain very similar results.[9] Second, actual defaults modify the realized returns. To compute the actual returns on a T-bill investment after default, we used the data set of defaults compiled by Reinhardt, Rogoff and Savastano (2003). 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), presented in the Annex.

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

[8] Section .2.1 in the appendix provides a list of developed countries.

[9] Default risk tends to increase the spread between portfolios, thus making it harder for our factor models to produce small pricing errors, not easier.

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, and we eliminate countries below the cutoff. In these countries, approval of both capital payments and receipts are rare, or the payments and receipts are at best only infrequently granted.

1.2 Foreign Currency Excess Returns

US as the home country Interest rates predict excess returns, and that is why we build portfolios of currencies sorted on the current interest rate gap with the US. The first panel of table 1 lists the mean excess return, the standard deviation and the Sharpe ratio for each portfolio. The largest spread (between the first and the seventh portfolio) exceeds five percentage points for the 1971-2002 subsample. The average annual returns are almost monotonically increasing in the interest rate differential. The only exception is the last portfolio, which is comprised of high inflation currencies: the average interest rate difference for the eight portfolio is about 23 percent over the entire sample from 1953-2002. 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.

This pattern is not due to default risk. We find a similar pattern for developed countries in the second panel of table 1. Their spreads are only slightly smaller (between 3.5 and 4 percentage points between the first and the seventh portfolio for annual data, between 3 and 3.5 percentage points for quarterly data). And in this case UIP does not hold for the eight portfolio either.

Countries change portfolios frequently. In annual data, countries change portfolios 23 percent of the time, 14 percent in quarterly data. The changing composition of the portfolios is critical. If we allocate currencies into portfolios based on the average interest rate differential over the entire sample instead, then there is essentially no pattern in average excess returns, and the average excess return on the last portfolio is invariably below *minus* 5 percent.[10] Remarkably, the annualized returns a US investor earns from quarterly re-balanced portfolios are substantially more volatile than the returns from annually re-balanced portfolios. The sorting introduces mean-reversion in the average exchange rate of

[10]See table 13 in the separate appendix. Of course, this is not a feasible investment strategy.

each portfolio, even though there is little evidence of mean reversion in individual currency's exchange rates.

We generated standard errors on these moments by bootstrapping from actual returns.[11] The standard errors for the mean return are large[12], but the lowest excess return is generally more than one standard deviation below zero, while the highest standard spread is more than one standard deviation above zero, at least for the quarterly returns. Moreover, a very similar pattern obtains when looking at the excess returns from the perspective of foreign investors.

Cross-Country Comparison of Foreign Currency Excess Returns We repeat the same portfolio building exercise for 10 developed countries (those countries which have good consumption data). Take the case of the UK. We allocate all the currencies into portfolios based on the interest rate differential with the UK, and we compute the average excess return in £ for each portfolio. We find very similar patterns in every country.

Table 1 reports only the 11-country average (including the US) for the mean, standard deviation and the Sharpe ratio.[13] In annual data, the spread is 4.5 percentage points, 5.5 percentage points in quarterly data. If anything, the spreads are larger on average from the perspective of other foreign investors. Since the standard deviation of the returns on quarterly re-balanced portfolios is much higher, the Sharpe ratios on this investment strategy are smaller in absolute value. As before, the last portfolio is an exception: *very* high interest rate currencies do not yield excess returns on average.

Our currency portfolios create a stable set of excess returns, even across different countries. In order to explain these currency excess returns, we draw from a whole class linear factor models that have proven successful in pricing equity and bond returns.

1.3 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:

$$m_{t+1} = b_0 + \sum_{j=1}^{n} b_j f_{j,t+1}, \tag{3}$$

where $f_{j,t+1}, j = 1, \dots n$ are the n factors. This encompasses two large classes of pricing models.

The first class uses returns as pricing factors. In this group are the Capital Asset Pricing Model (*CAPM*), the factor models by Fama and French (1992) and the model by Santos and Veronesi (2005). Fama and French (1993) argue that these factors proxy for the underlying undiversifiable macroeconomic risk. Santos and Veronesi (2005) add a scaling

[11] Allowing for predictability by bootstrapping from the residuals of an AR-process does not change this.
[12] See table 11 and 12 in the separate appendix
[13] Table 15 and 16 in the separate appendix list the detailed results.

variable - labor income share - to the standard *CAPM*, based on an extension of Lucas (1978)'s equilibrium model to two trees; a labor income tree and a dividend tree.

The second class of models comprises the the Consumption-*CAPM* (*CCAPM*), its scaled versions (Piazzesi et al. (2002), Lettau and Ludvigson (2001), Lustig and Nieuwerburgh (2005a)) and other derivatives (Yogo (2005), Parker and Julliard (2005)).

Table 2 summarizes the factors we used. Cochrane (2005) presents an extensive survey of all these models and their foundations. They are all related to the two basic workhorses of the field, the *CAPM* and the *CCAPM*. The scaled versions of the *CAPM* and *CCAPM* introduce time-variation in the market price of risk and go a long way in resolving the equity premium and risk-free puzzles.[14] The relative success of the models proposed by Santos and Veronesi (2005), Lettau and Ludvigson (2001) and Lustig and Nieuwerburgh (2005a) in pricing domestic stock returns suggests that the Fama-French asset pricing factors do proxy for underlying macroeconomic risk. We will show that the consumption-based models can price both domestic equity risk and currency risk, which the Fama-French factors cannot.

We have set up a framework where linear factor models, some with time-varying market price of risk, can be tested on foreign currency portfolios through unconditional pricing of the investor's Euler equation, and we now turn to the estimation results.

2 Estimation

In this section, we test the Euler equation of a US investor for each of these currency portfolios, running a horse race between the pricing factors presented above. Following Hansen (1982), we estimate an unconditional version of the linear factor models using the general method of moments (GMM). We normalize the SDF to $m_{t+1} = 1 - b' f_{t+1}$.[15] We use $E_T(x_t)$ and $var_T(x_t x_t')$ to denote the sample moments of a random vector x_t.

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

$$g_T(b) = E_T(m_t R_t^e) = E_T(R_t^e) - E_T(R_t^e f_t')b,$$

where $R_t^e = [R_t^{1,e} \; R_t^{2,e} \; .. \; R_t^{N^p,e}]'$. In the first stage of the estimation procedure, we use

[14] For example, 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 in Santos and Veronesi (2005). In Lustig and Nieuwerburgh (2005a), 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 the housing collateral ratio my to keep it positive as follows: $my_{max} - my$. This makes the scaling variable an indicator of collateral scarcity. Lustig and Nieuwerburgh (2005a) explain how the ratio of collateralizable wealth is measured empirically as the residual from a co-integrating relationship between labor income and housing wealth, along the lines of the computation by Lettau and Ludvigson (2001) for the consumption-wealth ratio. Lettau and Ludvigson (2005) show that consumption, dividends from asset wealth, and dividends from human capital (labor income) are cointegrated. cdy is computed as the cointegration residual from a consumption-based present-value relation involving future dividend growth. Lettau and Ludvigson (2005) show that cdy summarizes expectations of future dividend growth and forecasts long-horizon excess returns on the US stock market.

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

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: $S = \sum_{j=-\infty}^{\infty} E[(m_t R_t^e)(m_{t-j} R_{t-j}^e)']$. The optimal number of lags in the estimation of the spectral density matrix above is determined using Andrews (1991). When pricing a large number of portfolios, this procedure is computationally intensive. So, we have used 4 lags on annual data and 12 lags on quarterly data when we use more than eight test assets. Since we focus on linear factor models, GMM is equivalent to a 2-stage linear regression of the average excess returns $Y = E_T(R_t^e)$ on the factor/return moments $X = E_T(Rf_t')$. Chapter 13 of Cochrane (2001) describes this estimation procedure and compares it to the one proposed by Fama and MacBeth (1973).

Market Price of Risk The Euler equation for excess returns can be rewritten as the product of the portfolio beta and the market price of risk:

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

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.

In the simplest case of the *CCAPM*, the only factor is consumption growth, $f_t = \Delta \log c_t$; the coefficient b equals the coefficient of risk aversion γ, and the market price of risk is given by $\lambda = \gamma \frac{1}{1/E(m)} var(\Delta \log c_t)$.

Moment Conditions We first test the pricing models on our eight foreign currency portfolios. This gives us eight sample moment conditions we can use to estimate the model:

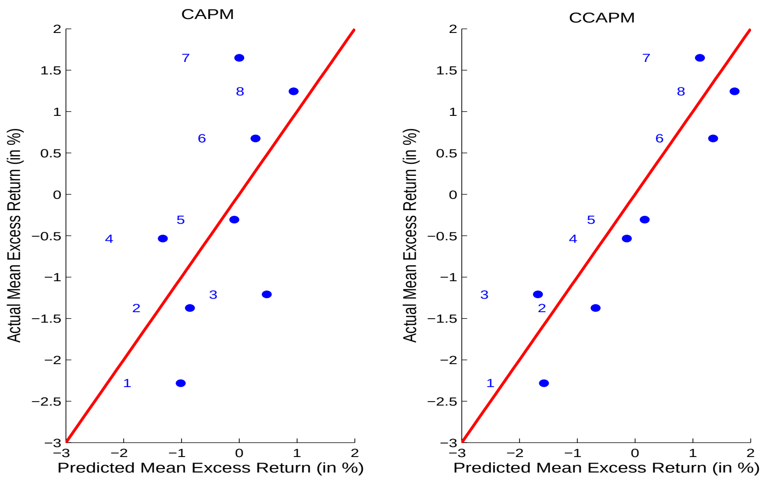
$$E_T \left[m_{t+1} R_{t+1}^{e,j} \right] = 0, j = 1, \ldots 8, \tag{4}$$

where E_T denotes the sample moment, and we examine each model's pricing errors $E_T(R_t^{e,j}) - \beta^j \lambda, j = 1, \ldots 8$, for each of the portfolios. Next, we introduce additional test assets to study whether 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.

2.1 Consumption-based Models

We start with the standard *CCAPM*, and its extensions. Then we switch to scaled versions of the *CCAPM*. The next section discusses the return-based factor models.

Figure 2: *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.



2.1.1 *CCAPM*

We use both annually and quarterly re-balanced currency portfolios as test assets. The *CCAPM* does very well on annual data.

Annual portfolio returns The standard *CCAPM* explains between sixty percent and eighty percent of the cross-sectional variation in average excess returns earned by a US investor on annually re-balanced currency portfolios: sixty percent for the 1971-2003 period and eighty percent for the 1953-2002 period. In contrast, the workhorse *CAPM* hardly explains any of the variation. This is apparent from figure 2: 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 $\beta^j_R\lambda_R$. This reflects the simple fact that there is very little variation in *CAPM* market betas across these eight portfolios, while there is a large difference of seventy-five basis points between the first (-.35) and the seventh portfolio (.3) in the *CCAPM* betas[16].

Table 3 reports the estimated market prices of risk and the p-value, in addition to the R^2, the R^2 adjusted for the number of estimated parameters, the mean squared pricing error (in percentage points), and the mean absolute pricing error (also in percentage points).

The estimated price of consumption growth risk λ_c is positive and large, around five. An asset with a consumption growth beta of one yields an average risk premium of five percent per annum. This number is similar for all of the consumption-based models, except

[16] shown in a separate appendix, figure 13.

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 (not reported in the table). This is in line with stock-based estimates of the coefficient of risk aversion found in the literature. The mean squared pricing error (*mspe*) on these eight currency portfolios is about 32 basis points over the entire sample, compared to 104 basis points for the *CAPM* (see table 4).

The coefficient estimates b, not reported in the table,[17] can easily be recovered from the risk price estimates. These reveal whether individual factors have explanatory power for currency risk premia, rather than wether the risk is priced. b_c is significant and positive across most models and sub-samples. For the *CCAPM* and the *HCAPM*, b_c is the estimated coefficient of relative risk aversion. It is around 50 (s.e. of 5) in annual data in the first two models. For the *DCAPM*, $b_c + b_x$ is the coefficient of risk aversion: it is between 30 and 40 in the annual data. The large positive coefficient b_x on $\Delta \log d_t$ reveals that the EIS $(1/\gamma)$ is much smaller than the intratemporal elasticity of substitution between durables and non-durables. As a result, the price of durable consumption growth risk is positive.

To give an overview, we plot the predicted against actual excess returns for all 4 factor models and 6 consumption-based models in figure 3. The single-factor *CCAPM* clearly does as well or better than the multi-factor models without consumption growth.

Quarterly portfolio returns 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 standard *CCAPM* explains only forty percent of the variation in returns on quarterly re-balanced portfolios, but the *mspe* is only half the *CAPM*'s (see table 4). Yogo (2005)'s model explains up to ninety 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 1.5 percent to 3.25 percent on an asset with a consumption growth beta of one, or between 6 and 13 percent annually. This is substantially higher than the estimated consumption growth risk premium from annual data.

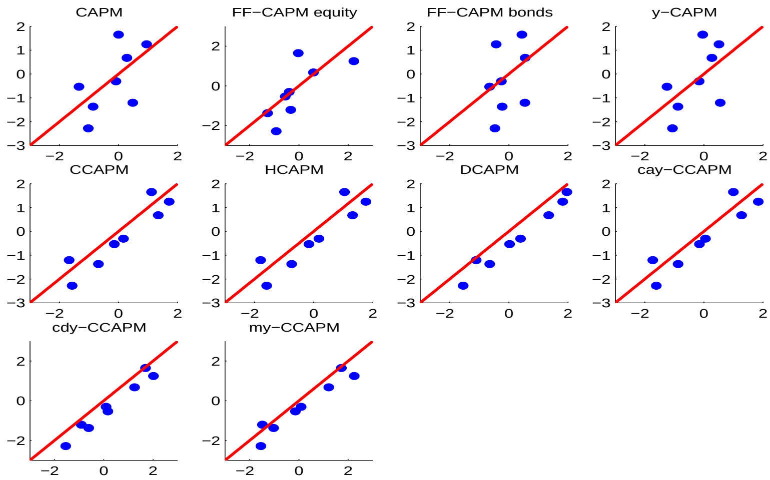
On quarterly data, the estimated coefficients of risk aversion are two to three times higher: in the *CCAPM* and the *HCAPM*, the estimates for γ are 120 and 160 respectively; the same number is around 100 in the *DCAPM*.

2.1.2 Scaled *CCAPM*

The scaled versions of the *CCAPM* capture the variation in currency risk premia, 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

[17] We reported the coefficient estimates b for all the models discussed in this paper in table 19 in the separate appendix.

Figure 3: 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 10 linear factor models. The filled dots are the currency portfolios.



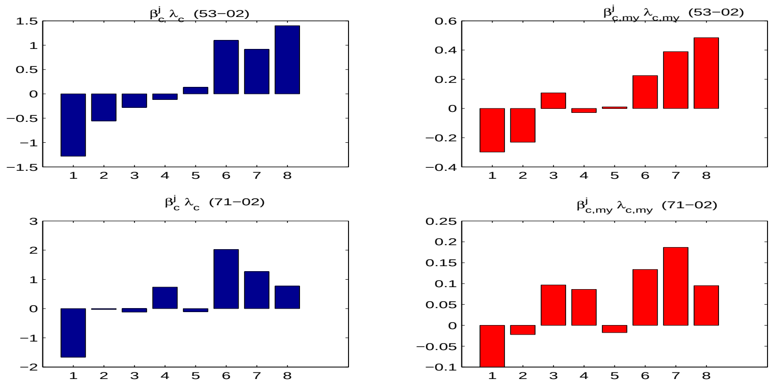
predicted by the model consists of two parts:

$$E[R^{e,j}] = \beta^j_c \lambda_c + \beta^j_{c,x} \lambda_{c,x}.$$

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. In other words, when the investor is more risk-averse, expected returns are higher. We take the example of the housing-collateral model to show the relative importance of the two parts in the equation above. Figure 4 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, -1.2 percentage points are due to consumption growth risk and about -.4 percentage points are due to consumption-growth-collateral risk. For high interest rate currencies, 1.5 percentage points are due to consumption growth risk and about .5 percentage points are due to consumption-growth-collateral risk.

In annual data, the estimated coefficients $b_{c,x}$ for the interaction term with the scaling variable are mostly positive and significant for the *my-CCAPM* and *cdy-CCAPM*, but not always for the *cay-CCAPM*. The scaling factors *cay* and *my* have low explanatory power for the quarterly returns. The implied coefficient of risk aversion cannot be recovered from these unconditional estimates of scaled *CCAPM* models (for recent work on estimating conditional factor models see Roussanov (2004)).

Figure 4: *my-CCAPM*: Risk Premia 1953-2002. GMM estimates using 8 currency portfolios as test assets. Annual Data.



2.1.3 Long-Run Consumption Risk

Parker and Julliard (2005) demonstrate that long-run measures of consumption growth risk outperform the standard *CCAPM* in explaining stock returns. In table 3 we only report the results for the optimal lead length. Two-year consumption growth outperforms one-year consumption growth in explaining currency risk for the annually rebalanced portfolios. Similarly, in quarterly data, the 5-quarter consumption growth rate explains 78 percent of the variation in quarterly returns over the entire sample, while the standard *CCAPM* explains only 39 percent. These long-run measures really outperform the benchmark *CCAPM*.

Next we compare the performance of the consumption-based models with the return-based models of m.

2.2 *CAPM* and Extensions

On annual data, the basic *CAPM* explains only 36 percent of the variation in 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. On annual returns the average pricing errors for the consumption-based models are only half the size of those for the return-based models. On quarterly data, only the Fama-French bond factors explain a large part of the variation in excess return over the whole sample 1953:1-2002:4, 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 28 monthly foreign excess returns over the 1976-1998 period. There is no relation between the stock market betas of currencies and the interest rates, or in other words, there is no spread in the stock market betas of the average exchange rate change in the currency portfolios. As a result, the *CAPM* and extensions of the *CAPM* cannot price currency risk.[18]

3 Robustness and extensions

In this section, we want to check the robustness of our results by changing (1) the sample of countries in the portfolios -only developed countries-, (2) the test assets -other assets like stocks and bonds-, (3) the construction of the currency portfolios themselves and (4) the nationality of the investor whose Euler equation we test.

3.1 Developed Countries

If we limit the sample to developed countries, the individual portfolio returns are less informative because the portfolios contain fewer countries. Still, we want to guard against the possibility that default risk is driving our results. The *CCAPM* explains between 46 and 38 percent of the variation in returns on the annually rebalanced portfolios.[19] The price of consumption growth risk is estimated quite precisely between 1.5 and 2 percentage points, about half of the number we found when we used the entire sample. For the *CCAPM*, the estimated coefficient of risk aversion is 30 (s.e. of 3) over the entire sample, and 52 (s.e. of 4) in the post-Bretton-Woods sample. The standard *CCAPM* breaks down in quarterly data, as do most of the other consumption-based models, except for the *DCAPM*. Only the *DCAPM* does quite well both in quarterly and annual returns: it explains about 60 percent of the variation. The market price of durable consumption growth risk is positive and significant.

Finally, as before, the *CAPM* and its extensions explain none of the variation in returns across these currency portfolios and all four factor models are rejected by the data in the longest, quarterly sample.[20]

This confirms that our results are not driven by default risk, but currency risk. A key question is then whether currency risk is priced differently from equity risk and bond risk, that is to say, whether the same m prices the returns in currency, bond and equity markets. To address this question, the next section adds domestic test assets to the currency portfolios.

3.2 Domestic Test Assets

First, we add stock portfolios as test assets. In a second step, we add bond portfolios as well.

[18]Note that the $y-CAPM$ does much better if we include the scaling variable (the labor income/consumption ratio) as a separate factor. This case is not reported in the tables, but it also argues in favor of the introduction of macroeconomic risk.

[19]See table 17 in the separate appendix.

[20]See table 18 in the separate appendix.

Stocks We examine whether the compensation for aggregate risk in currency markets differs from the one applied in domestic equity markets, as before, 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. This leaves us with 33 sample moment conditions:

$$E_T\left[m_{t+1}R^{e,j}_{t+1}\right] = 0, j = 1, \dots 8 + 25. \tag{5}$$

These Fama-French portfolios sort stocks according to size and book to market quintiles, because both size and book-to-market predict returns. We want to find out if these returns can be priced by the same stochastic discount factor m that prices currency risk.

Figure 5 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. As is apparent from the graphs, the first class of models, which uses financial returns as risk factors, cannot account for both the variation in equity and currency returns, while the consumption-based models can. The *my-CCAPM* does very well in annual data, while the *DCAPM* does better in quarterly data.

Figure 5: 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 10 linear factor models. The filled dots are the currency portfolios.

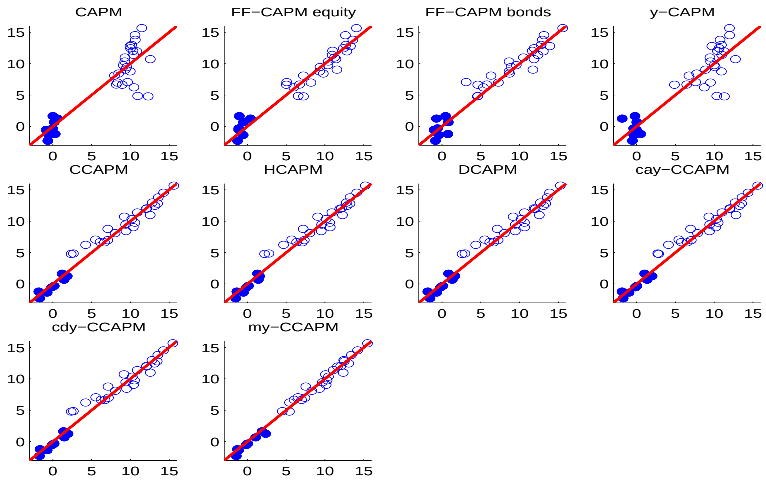


Table 5 reports the pricing error statistics for the eight currency portfolios only. The mean squared pricing error (*mspe*) for the eight currency portfolios are almost always lower for consumption-based models than for the return-based models. More importantly, the consumption-based models also do well in explaining the excess returns on the 25 test assets.[21] The standard *CCAPM* explains a substantial share of the total variation in

[21]Pricing errors for these 25 portfolios are not reported but available upon request.

annual stock and currency portfolio returns, although the *mspe* is more than twice as large as for the *my-CCAPM*. The *mspe* for the *CCAPM* is about 1 percentage point against 37 basis points for the *my-CCAPM*. On quarterly data, the *DCAPM* does better than the other consumption-based models.

There are clear gains in efficiency from combining the two set of test assets. The risk prices are estimated very precisely, as shown in table 6. The price of consumption growth risk estimated from annual data is 2 to 3 percentage points higher compared to the currency only estimates, but the consumption growth risk price estimates from quarterly data are very close to the currency-only estimates. The coefficient estimates for consumption growth and for consumption growth interacted with the housing collateral ratio my and the consumption-wealth ratio cdy have the right sign in accordance with the theory, and the estimates are very precise.[22] The implied coefficient of risk aversion in the *CCAPM* is around 54 (with a standard error of 1), 50 in the *HCAPM* and 60 in the *DCAPM*. The coefficients $b_{c,x}$ on consumption growth and the scaling variable are positive and significant. These estimates are very close to the ones we obtained using only the currency portfolios as test assets.

3.3 Out of Sample Test: Stocks and Bonds as Test Assets

We also ran an out-of-sample-test by estimating the model on the 6 bond portfolios and the 6 Fama-French benchmark portfolios, and then pricing the currency portfolio returns. Even though the pricing errors are quite large overall, the consumption-based models still produce much smaller pricing errors on the 8 currency portfolios.[23] This confirms that only consumption-based models can truly price both equity, bond and currency risk.

3.4 Managed Portfolios as Test Assets

The currency portfolios allocate currencies to bins based on their rank in the entire distribution of interest rates, but the absolute size of a currency's interest rate difference with the US predicts excess returns as well.

To address this, we also build portfolios that go long and short in a currency T-bills depending on that currency's interest rate, and we test the investor's Euler equation on each of these managed currency portfolios. This is equivalent to using the nominal interest rate differential itself as an instrument z_t:

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

where z^i_t is the interest rate differential between the foreign T-bill and its US counterpart $R^{i,*}_{t,t+1} - R^{f,\$}_{t,t+1}$ and n is the number of countries in the sample. This procedure is equivalent

[22] See table 22 in the separate appendix.

[23] Detailed statistics are reported in tables 23 and 24 and figure 14 in the separate appendix.

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:

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

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. This yields n moment conditions we can use to estimate the model:

$$E_T \left[m_{t+1} \widetilde{R}_{t+1}^{e,i} \right] = 0, i = 1, \ldots n, \tag{8}$$

Foreign Currency Excess Returns By taking this approach, we avoid the loss of information that results from aggregating currencies into portfolios, as we did before, but this comes at the cost of restricting our sample to those countries with data that span the entire time period.

We constructed these managed portfolios for 11 countries between 1971 and 2002, and for eight countries between 1956 and 2002. These highly leveraged trading strategies produce large excess returns in excess of twenty percent and Sharpe ratios in excess of .5, and there is substantial variation in average excess returns across countries as well.[24]

Estimation Results Table 7 reports the results. The explanatory power of the consumption-based models is much lower than in the other pricing exercises. The risk premia that are due to movements in a single 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, the estimated prices of aggregate consumption growth risk λ_c and the price of consumption growth risk in bad times $\lambda_{c,x}$ are positive, significant and have the right sign. These estimates are in line with our previous ones.

In the case of managed portfolios, there is an even starker contrast between the performance of these two classes of models. The *DCAPM* and the scaled *CCAPM* explain up to fifty percent of the cross-sectional variation in returns on these strategies. The conditioning information embedded in *cdy* and *my* plays a more important role here. On the other hand, the factor-based models do very poorly: they explain none of the variation and the price of market risk is significantly negative for most of the factors.

All these additional experiments show that US consumption growth risk is priced in currency risk premia, and this result is quite robust across frequencies, samples and different ways of building portfolios.

Of course, (foreign) consumption growth risk should be priced into the excess returns earned by investors in other countries as well. To check that, we test the Consumption-

[24]The characteristics of these managed portfolios are presented in figures 11 and 12 in the separate appendix.

CAPM on a cross-section of currency portfolios built from the perspective of domestic investors in other developed countries.

3.5 Other Countries

Sofar we have focused only on the Euler equation of a US investor. This sections looks at investors in 10 other developed countries. These are the countries for which reliable aggregate consumption data were available.[25] We built the currency portfolios for these other countries and we compute the excess returns for each of these portfolios in units of the foreign consumption.

Foreign Currency Excess Returns These portfolios share the same basic pattern we documented for the US returns: the returns increase from the first to the seventh bin and the spread is between four and five percentage points[26]. We report a summary for three of these in table 8.

Estimation results Since we do not have aggregate durable consumption growth, rental price growth and the scaling factors for each of these countries, we only test the standard *CCAPM*. We focus only on the annually re-balanced portfolios.[27]

Table 6 summarizes the estimation results. Except for France and Sweden, the *CCAPM* explains between 50 and 70 percent of the variation in excess returns in the longest sample, slightly less in the shorter sample, although the average pricing errors are still substantial.

In addition, we pooled all of these observations for 11 countries to test a single Euler equation. The implied coefficient of risk aversion is around 40, with a standard error of around 6. Consumption growth explains between seventy and eighty percent of the variation on these pooled portfolios, and the mean absolute pricing error is only 50 basis points.

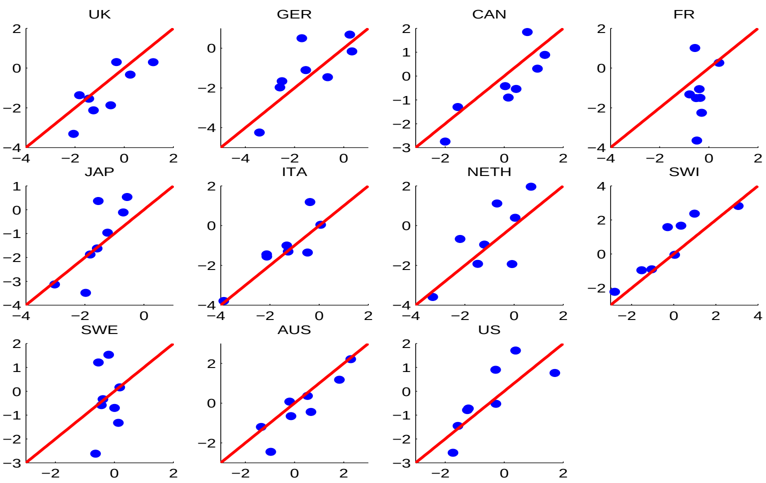
4 What drives these results?

We have shown that the predicted excess returns line up with the realized ones when the pricing factors take into account consumption growth risk. This is not mere luck on our part. The previous section provided many robustness checks.

This section sheds some light on the underlying economic mechanisms. The foreign currency excess return can be restated in terms of consumption growth betas. These betas differ from one portfolio to the other and they change in the right direction to explain the UIP puzzle.

[25] We have used and updated the data set built by John Campbell and available on his web site.
[26] See tables 15 and 16 in the appendix.
[27] The results for the quarterly data are available upon request.

Figure 6: Predicted vs. Actual Excess Return for 8 Currency Portfolios between 1962-2002. 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 11 countries. The filled dots are the currency portfolios.



Two mechanism can give rise to the monotonic relation between the consumption growth betas of exchange rates and interest rates in the data: (1) negative correlation between the interest rate and the second moment of the foreign stochastic discount factor (SDF), and/or (2) a higher correlation of the SDFs with their US counterpart in low interest rate currencies.

To obtain these results, we assume that markets are complete and that the SDFs are log-normal. Essentially, we re-interpret an existing derivation by Backus et al. (2001), and we explore its empirical implications.

4.1 Consumption Growth Betas and Currency Risk Premia

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 conditional 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):

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

where $\log R^f_{t,t+1}$ denotes the risk free rate, in units of US consumption. 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. It is important to note that these conditional risk premia do not depend on interest rates, while the unconditional ones we explained in the previous section do: we abstract from the part of the risk premium due to the unconditional covariance between interest rates and the SDF.

Consumption Growth and Exchange Rates The log 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 $x_t > 0$. This is not a surprising approximation. In the standard *CCAPM*, the log stochastic discount factor is equal to $\log m_{t+1} = \log \beta - \gamma \Delta \log c_{t+1}$ and b'_1 is (minus) the coefficient of relative risk aversion γ. As we have already seen, the scaling variable x_t introduces time-variation in the market price of consumption growth risk.

It immediately 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 (once again, what the finance literature calls consumption growth betas):

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

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:[28]

1. *the consumption growth betas of currencies need to be small when foreign interest rates are low and large 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

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

condition needs to be satisfied:

$$
\begin{aligned}
&Cov_t\left(\Delta \log c_{t+1}, \Delta \log e^i_{t+1}\right) \text{ \textit{small when} } R^{i,£}_{t,t+1} \text{ \textit{is low},}\\
&Cov_t\left(\Delta \log c_{t+1}, \Delta \log e^i_{t+1}\right) \text{ \textit{large when} } R^{i,£}_{t,t+1} \text{ \textit{is high}.}
\end{aligned}
$$

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 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.

Empirical evidence To check in a very simple way whether the necessary condition outlined above is satisfied in the data, we examine the unconditional and conditional consumption growth betas of exchange rates for the first and seventh portfolio, because the average returns of these two are farthest apart.

Unconditional Consumption Growth Betas Of course, we do not know the conditional expected return on each portfolio nor do we know the conditional covariance, but, by building the portfolios, we have conditioned on the interest rates. So, it seems informative to examine the consumption growth betas for the first six portfolios by regressing the deflated average change in the exchange rate on US consumption growth:

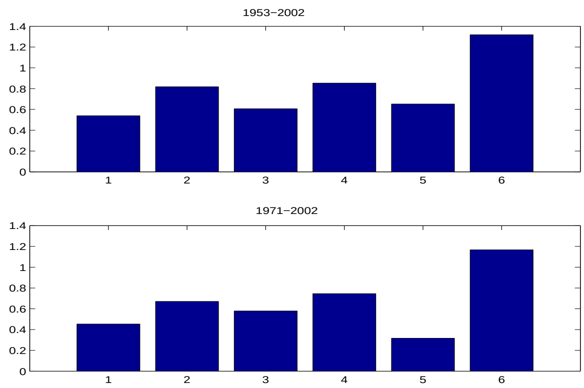
$$\Delta \log e^i_{t+1} - \Delta \log p_{t+1} = \alpha^i_0 + \alpha^i_1 \Delta \log c^{US}_{t+1} + \epsilon^i_{t+1}, i = 1, \ldots, 6$$

As we expected, the unconditional consumption growth betas of foreign cash holdings (here α^i_1) are close to zero for the currencies in the first portfolio and much higher for the currencies in the sixth portfolio. These betas are shown in figure 7, and they explain the pattern in excess return betas, to a large extent. Not completely, because there is interest rate risk too.

Most of our results can be understood through this basic finding. Foreign cash holdings in the high interest rate currencies expose US investors to more consumption growth risk, while foreign cash holdings in the low interest rate currencies do not. The beta on the first portfolio is .2, while the beta on the sixth is closer to 1.2 (t-stat of 2.12).

The betas for the very high interest rate currencies in the seventh and eight portfolio look radically different -they are large and negative, but not significant. These currencies are extremely volatile, but most of the currency risk is uncorrelated with US consumption growth risk and hence not priced.

Figure 7: Consumption Growth Betas of Exchange Rates Estimated slope coefficients in regression of percentage exchange rate changes on US consumption growth for the first and seventh currency portfolios. The figure shows α_1^i in $\Delta \log e_{t+1}^i - \Delta \log p_{t+1} = \alpha_0^i + \alpha_1^i \ \Delta \log c_{t+1}^{US} \ + \epsilon_{t+1}^i$. We used the entire sample (1953-2002) and the post-Bretton Woods sample (1971-2002) on annual data.

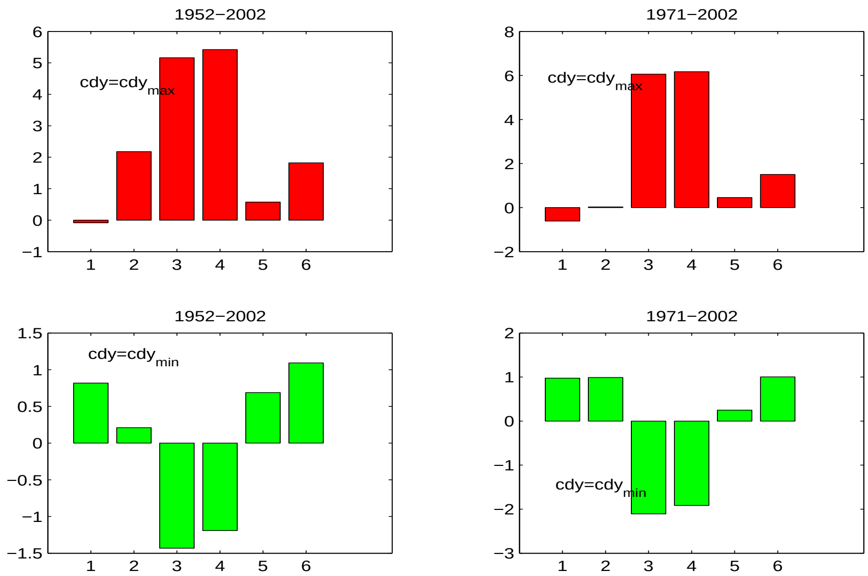


Conditional Consumption Growth Betas Just on the basis of consumption growth risk, the currencies in the third and fourth portfolio did not seem much riskier than those in the second portfolio, but higher interest rate currencies become riskier in bad times. To evaluate the conditional consumption growth betas, we run the following regression of exchange rates on US consumption growth and US consumption growth interacted with the scaling variable:

$$\Delta \log e_{t+1}^i - \Delta \log p_{t+1} = \alpha_0^i + \alpha_1^i \Delta \log c_{t+1}^{US} + \alpha_2^i x_t \Delta \log c_{t+1}^{US} + \epsilon_{t+1}^i, i = 1, \ldots, 6$$

We find that the sensitivity of higher interest rate currencies increases in bad times. We consider the consumption-wealth ratio *cdy* as the scaling variable x. Figure 8 plots the consumption growth betas evaluated at *cdy* equal to its maximum sample value in the top panel. The bottom panel plots the consumption growth betas evaluated at *cdy* equal to its minimum sample value. When the consumption-wealth ratio is high relative to its mean, the consumption growth betas for high interest rate currencies in the third and fourth portfolio are nearly five times the unconditional values, while the lowest interest rate currencies are essentially uncorrelated or even negatively correlated with US consumption growth risk in bad times.

Figure 8: Consumption Growth Betas of Exchange Rates in Bad and Good Times The upper panel shows the bad times consumption growth betas, the lower panel the good times betas. The figure shows α_1^i for exchange rate changes in: $\Delta \log e_{t+1}^i - \Delta \log p_{t+1} = \alpha_0 + \alpha_1 \ \Delta \log c_{t+1}^{US} \ + x_t \alpha_2 \ \Delta \log c_{t+1}^{US} \ + \epsilon_{t+1}^i$. The consumption growth betas of exchange rates in bad times are computed as $\alpha_1^i + \alpha_2^i x^b$ where x^b is either equal to the maximum sample value of x (bad times, upper panel) or the minimum sample value (good times, lower panel). The scaling variable x is cdy. The samples are 1953-2002 and 1971-2002. Annual Data.



4.2 Consumption Co-movements and Interest Rates

The observed pattern in excess returns can be linked back to the properties of the SDF. Assuming that the inflation betas are small enough and that markets are complete, the size 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 SDF's (see proof in Annex):

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

Heteroskedasticity First, suppose that the correlation between the SDFs is positive and constant. If countries characterized by a high interest rate $R_{t,t+1}^{i,\pounds}$, typically also have a low conditional volatility of the foreign SDF m_{t+1}^i 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. 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 or high marginal utility growth states!

In fact, suppose they are perfectly correlated, but the domestic SDF is only half as volatile, and 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 five percent. So, investing in this foreign currency provides a perfect hedge for the US investor.

In our benchmark model, an increase in the conditional volatility of aggregate consumption growth lowers the real interest rate.[29] If real and nominal interest rates move in sync, a low nominal interest rate should predict a higher conditional volatility of aggregate consumption growth. Of course, if inflation is very high and volatile, the nominal and the real interest rates effectively are detached, and this mechanism would disappear, as it seems to in the data.

Richer specifications rely on mechanisms other than changes in the conditional distribution of consumption growth to activate this heteroskedasticity mechanism. This behavior is for example 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 The second mechanism keeps the conditional volatilities of the SDFs constant, but allows for time-varying correlation. If the conditional correlation of the SDFs is positive for low-interest countries and negative for high-interest rate countries, then this can account for the cross-section of risk premia. We want to test wether the conditional correlation of the SDFs decreases with the interest rate differential.

We consider 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 size of the conditional risk premium is determined by:

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

A low correlation of foreign consumption growth with US consumption growth for high interest rate currencies implies large currency risk premia. What is the economic intuition behind this mechanism? If the consumption process of a high interest rate country is weakly correlated with US consumption, then a negative consumption shock in the US leads to

[29]This can be shown by starting from the Euler definition of the real risk-free rate and by assuming that aggregate consumption growth is log-normal.

an appreciation of the dollar and a lower foreign return. This currency depreciates in high marginal utility growth states for the US investor. The data support this time-varying correlation mechanism.

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

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

The results obtained over the post-Bretton Woods period are reported in table 10. 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 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.

5 Conclusion

Aggregate consumption growth risk explains a large fraction of the average changes in the exchange rate, conditional on foreign interest rates. On average, high interest rate currencies depreciate when US consumption growth is low and US investors want to be compensated for this risk. Thus, aggregate consumption growth risk is key to understanding exchange rates. So far real exchange rates appear unrelated to consumption data (e.g. Backus and Smith (1993) and Chari, Kehoe and McGrattan (2002)). But our results suggest that the correlation between the two is time-dependent because of time-varying risk aversion, captured here by the consumption-wealth ratio, or time-varying degrees of risk sharing within countries, captured by the housing collateral ratio. Future research should test whether these richer models of m can directly explain the behavior of real exchange rates.

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 home country at the end of period $t-1$. Portfolio 1 contains currencies with the smallest interest rate differential. The sample includes all countries in a given year which are assigned a Quinn capital account liberalization index that exceeds 20. Panel I reports the characteristics of the excess returns of an American investor investing in all countries. Panel II reports the characteristics of the excess returns of an American investor investing in developed countries. Panel III reports characteristics of the *average* excess returns of all investor investing in all countries.

	1	2	3	4	5	6	7	8
Panel I: All Countries								
Annual Returns								
	1953-2002							
Mean	−2.28	−1.37	−1.21	−0.53	−0.31	0.67	1.65	1.24
Std.	6.32	6.43	8.89	8.51	8.66	6.14	11.51	11.39
Sharpe Ratio	−0.36	−0.21	−0.14	−0.06	−0.04	0.11	0.14	0.11
	1971-2002							
Mean	−2.90	−0.76	−0.37	−0.15	−0.91	1.18	2.24	0.42
Std.	7.79	6.60	8.85	10.25	10.76	7.54	14.36	13.74
Sharpe Ratio	−0.37	−0.12	−0.04	−0.01	−0.08	0.16	0.16	0.03
Annualized Quarterly Returns								
	1953.1-2002.4							
Mean	−2.83	−0.40	−1.42	0.92	−0.23	0.01	2.46	0.55
Std.	12.76	9.40	15.79	14.19	13.62	12.21	13.13	16.36
Sharpe Ratio	−0.22	−0.04	−0.09	0.06	−0.02	0.00	0.19	0.03
	1971.1-2002.4							
Mean	−2.92	−0.29	−0.36	1.59	−0.60	−0.33	3.48	−0.32
Std.	13.08	11.69	14.18	16.80	16.70	15.20	15.50	19.50
Sharpe Ratio	−0.22	−0.03	−0.03	0.09	−0.04	−0.02	0.22	−0.02
Panel II: US Investor, Developed Countries								
Annual Returns								
	1955-2002							
Mean	−0.44	0.84	0.97	1.75	0.86	−0.05	2.95	2.69
Std.	10.34	10.35	11.34	8.92	10.92	9.26	10.30	10.13
Sharpe Ratio	−0.04	0.08	0.09	0.20	0.08	−0.01	0.29	0.27
	1971-2002							
Mean	−0.18	1.75	1.65	2.57	0.97	−0.34	3.96	3.46
Std.	12.63	12.62	13.87	10.83	13.37	11.37	12.54	12.32
Sharpe Ratio	−0.01	0.14	0.12	0.24	0.07	−0.03	0.32	0.28
Annualized Quarterly Returns								
	1955.1-2002.4							
Mean	−1.20	0.84	0.65	1.24	0.79	1.96	2.65	2.58
Std.	19.34	18.19	18.97	17.56	17.59	17.03	17.08	15.55
Sharpe Ratio	−0.06	0.05	0.03	0.07	0.04	0.12	0.16	0.17
	1971.1-2002.4							
Mean	−1.18	1.49	1.42	1.51	0.62	2.59	3.24	3.08
Std.	23.61	22.22	23.11	21.33	21.44	20.73	20.88	18.95
Sharpe Ratio	−0.05	0.07	0.06	0.07	0.03	0.12	0.16	0.16
Panel III: All Investors, All Countries								
Annual Data								
	1953-2002							
Mean	−2.49	−0.95	−0.66	−0.46	−0.40	0.73	2.11	1.64
Std.	12.85	13.53	14.33	13.27	12.85	13.09	14.90	20.27
Sharpe Ratio	−0.31	−0.14	−0.10	−0.10	−0.10	0.02	0.12	0.06
	1971-2002							
Mean	−3.41	−1.25	−0.74	−0.83	−1.17	0.13	1.76	0.46
Std.	9.81	10.17	10.42	10.32	10.00	10.32	12.56	17.00
Sharpe Ratio	−0.35	−0.12	−0.07	−0.09	−0.12	0.02	0.14	0.03
Annualized Quarterly Returns								
	1953.1-2002.4							
Mean	−3.33	−0.60	−0.83	0.11	−1.01	−0.22	2.17	0.01
Std.	17.53	17.42	16.47	15.65	15.68	16.20	15.33	22.05
Sharpe ratio	−0.17	−0.03	−0.05	0.01	−0.06	−0.01	0.13	−0.00
	1971.1-2002.4							
Mean	−2.68	−0.38	−0.66	0.37	−0.90	−0.33	2.29	−0.25
Std.	16.44	17.25	16.33	15.09	15.49	16.03	14.94	21.68
Sharpe ratio	−0.15	−0.02	−0.04	0.02	−0.05	−0.02	0.14	−0.01

Table 2: Linear Factor Models

The upper panel contains models with returns as factors. The Fama-French equity pricing factors ($FF - CAPMequity$) are the CRSP value-weighted excess return R^{vw}, the small-minus-big return R^{SB} and the high-minus-low return R^{HL}. The Fama-French bond pricing factors ($FF - CAPMbonds$) are the difference between the long term government bond return and the risk free rate (R^{long}) and the spread between the return on a long-term corporate bond index and a long term government bond (R^{corp}). Santos and Veronesi (2005)'s model (denoted here $y - CAPM$) uses the labor income share $\frac{l}{c}$ as scaling variable. The lower panel contains consumption-based models. The standard Consumption-CAPM ($CCAPM$) uses only aggregate consumption growth $\Delta(\log c_t)$. The Housing-$CAPM$ (henceforth $HCAPM$) proposed by Piazzesi et al. (2002) adds rental price growth $\Delta(\log \rho_t)$, scaled by the housing expenditure share A_{t-1}, as an additional pricing factor. Yogo (2005)'s model (denoted $DCAPM$) uses the growth rate of the stock of durables $\Delta \log d_{t+1}$. Lettau and Ludvigson (2001) introduce two measures of the consumption wealth ratio cay and cdy as scaling variables. Lustig and Nieuwerburgh (2005a) introduce the housing collateral ratio my. Parker and Julliard (2005) rely on long-run consumption growth risk (denoted $LR - CCAPM$).

	f_1	f_2	f_3
CAPM	R^{vw}		
FF-CAPM equity	R^{vw}	R^{SB}	R^{HL}
FF-CAPM bonds	R^{long}	R^{corp}	
y-CAPM	R^{vw}	$R^{vw}\,\frac{l}{c}$	
CCAPM	$\Delta(\log c_t)$		
HCAPM	$\Delta(\log c_t)$	$A_{t-1}\Delta(\log \rho_t)$	
DCAPM	$\Delta(\log c_t)$	$\Delta(\log d_t)$	
cay-CCAPM	$\Delta(\log c_t)$	$\Delta(\log c_t)cay_{t-1}$	
cdy-CCAPM	$\Delta(\log c_t)$	$\Delta(\log c_t)cdy_{t-1}$	
my-CCAPM	$\Delta(\log c_t)$	$\Delta(\log c_t)my_{t-1}$	
LR-CCAPM	$\Delta(\log c_{t+s})$		

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, for the *HCAPM* and *DCAPM* and $\lambda_{c,x}$, for the $cay - CCAPM$, $cdy - CCAPM$ and $my - CCAPM$, 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* is $myfa$ (see data app.). We used the optimal lag length to estimate the spectral density matrix (Andrews, 1991).

Model	λ_c	$\lambda_{c,x}$	*p value*	R^2_{adj}	R^2	*mspe*	*mape*
Panel A: Annual Data							
	1953-2002						
CCAPM	5.16		0.70	0.81	0.81	0.32	0.55
s.e.	[2.74]						
HCAPM	4.93	0.37	0.73	0.76	0.80	0.32	0.57
s.e.	[2.64]	[0.39]					
DCAPM	3.01	4.64	0.78	0.75	0.79	0.32	0.55
s.e.	[1.89]	[2.25]					
cay-CCAPM	5.35	1.14	0.61	0.79	0.82	0.30	0.53
s.e.	[3.73]	[0.80]					
cdy-CCAPM	3.25	0.80	0.79	0.75	0.79	0.32	0.54
s.e.	[1.55]	[0.37]					
my-CCAPM	7.68	1.20	0.74	0.79	0.82	0.29	0.47
s.e.	[7.42]	[1.12]					
lr-CCAPM s=2	8.52		0.62	0.91	0.91	0.15	0.34
s.e.	[4.99]						
	1971-2002						
CCAPM	4.87		0.87	0.68	0.68	0.66	0.76
s.e.	[1.78]						
HCAPM	4.74	0.44	0.84	0.60	0.66	0.67	0.77
s.e.	[1.95]	[0.46]					
DCAPM	4.33	4.53	0.94	0.54	0.61	0.76	0.84
s.e.	[1.84]	[1.89]					
cay-CCAPM	4.35	0.97	0.84	0.62	0.67	0.66	0.76
s.e.	[1.64]	[0.36]					
cdy-CCAPM	3.72	0.92	0.91	0.57	0.63	0.71	0.80
s.e.	[1.17]	[0.28]					
my-CCAPM	4.84	0.53	0.83	0.61	0.66	0.68	0.79
s.e.	[1.79]	[0.22]					
lr-CCAPM s=2	8.45		0.83	0.88	0.88	0.25	0.42
s.e.	[4.94]						
Panel B: Quarterly Data							
	1953.1-2002.4						
CCAPM	3.25		0.17	0.40	0.40	0.083	0.25
s.e.	[2.35]						
HCAPM	1.95	−0.08	0.51	0.64	0.69	0.041	0.17
s.e.	[1.27]	[0.10]					
DCAPM	2.05	3.32	0.46	0.92	0.93	0.008	0.08
s.e.	[3.79]	[3.81]					
cay-CCAPM	2.98	0.80	0.12	0.33	0.43	0.081	0.25
s.e.	[2.21]	[0.57]					
my-CCAPM	2.05	0.12	0.12	0.41	0.49	0.07	0.22
s.e.	[1.40]	[0.13]					
lr-CCAPM s=5	10.66		0.92	0.78	0.78	0.03	0.13
s.e.	[6.03]						
	1971.1-2002.4						
CCAPM	1.49		0.12	0.33	0.33	0.13	0.31
s.e.	[0.64]						
HCAPM	2.55	0.00	0.75	0.38	0.47	0.11	0.29
s.e.	[2.18]	[0.15]					
DCAPM	4.54	2.18	0.70	0.85	0.88	0.023	0.13
s.e.	[6.81]	[3.05]					
cay-CCAPM	2.07	0.56	0.15	−0.07	0.08	0.13	0.33
s.e.	[0.99]	[0.27]					
my-CCAPM	1.78	0.19	0.08	0.16	0.28	0.12	0.32
s.e.	[0.92]	[0.11]					
lr-CCAPM s=5	9.20		0.74	0.61	0.61	0.07	0.22
s.e.	[5.02]						

Table 4: *CAPM* Risk Price Estimates for Currency Portfolios

GMM estimates using 8 currency portfolios as test assets. For the $FF - CAPMe$, the first column contains $\lambda_{R^{vw}}$, the second column $\lambda_{R^{SB}}$, the third column $\lambda_{R^{HL}}$. For the $FF - CAPMb$, the first column contains $\lambda_{R^{long}}$, the second 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	*p value*	R^2_{adj}	R^2	*mspe*	*mape*
Panel A: Annual Data								
	1953-2002							
CAPM	16.13			0.42	0.36	0.36	1.04	0.85
s.e.	[4.97]							
FF-CAPM e.	22.46	−33.74	29.25	0.29	0.31	0.51	0.80	0.65
s.e.	[27.54]	[49.00]	[45.65]					
FF-CAPM b.	3.55	1.74		0.30	−0.06	0.09	1.49	0.98
s.e.	[2.73]	[1.90]						
y-CAPM	19.39	21.29		0.41	0.11	0.24	1.24	0.93
s.e.	[7.28]	[8.21]						
	1971-2002							
CAPM	9.08			0.69	0.25	0.25	1.53	1.00
s.e.	[3.54]							
FF-CAPM e.	6.21	14.05	−6.07	0.92	−0.88	−0.34	2.48	1.33
s.e.	[5.80]	[7.20]	[8.78]					
FF-CAPM b.	7.42	1.43		0.54	0.32	0.41	1.22	0.77
s.e.	[6.77]	[2.28]						
y-CAPM	9.17	9.84		0.58	0.13	0.25	1.53	0.98
s.e.	[3.66]	[3.93]						
Panel B: Quarterly Data								
	1953.1-2002.4							
CAPM	4.54			0.00	−0.00	−0.00	0.13	0.31
s.e.	[3.93]							
FF-CAPM e.	9.82	2.22	4.76	0.05	−0.46	−0.04	0.13	0.32
s.e.	[8.08]	[3.07]	[2.97]					
FF-CAPM b.	1.44	5.56		0.94	0.79	0.82	0.023	0.14
s.e.	[4.59]	[2.90]						
y-CAPM	2.91	4.00		0.00	−0.02	0.13	0.12	0.28
s.e.	[3.79]	[5.55]						
	1971.1-2002.4							
CAPM	6.64			0.02	−0.01	−0.01	0.19	0.40
s.e.	[4.10]							
FF-CAPM e.	15.04	2.96	6.42	0.58	−0.14	0.18	0.12	0.33
s.e.	[9.95]	[3.52]	[4.23]					
FF-CAPM b.	5.57	0.64		0.11	−0.29	−0.10	0.15	0.37
s.e.	[3.81]	[1.01]						
y-CAPM	6.66	9.61		0.01	−0.19	−0.02	0.19	0.41
	[4.15]	[6.03]						

Table 5: Pricing Errors for Currency Portfolios

GMM estimates using 8 currency and 25 equity portfolios as test assets. For the consumption-based models, the first column contains λ_c, the second column λ_x, for the *HCAPM* and *DCAPM* and $\lambda_{c,x}$, for the $cay - CCAPM$, $cdy - CCAPM$ and $my - CCAPM$, where x denotes the scaling factor. For the return-based models, 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	R^2_{adj}	R^2	*mspe*	*mape*
Panel A: Annual Data				
1953-2002				
CAPM	0.41	0.41	0.89	0.79
FF-CAPM e.	−0.40	−0.00	1.52	1.05
FF-CAPM b.	0.17	0.29	1.08	0.84
y-CAPM	−0.78	−0.52	2.32	1.22
CCAPM	0.81	0.81	0.29	0.50
HCAPM	0.76	0.79	0.32	0.52
DCAPM	0.76	0.79	0.32	0.51
cay-CCAPM	0.80	0.83	0.26	0.48
cdy-CCAPM	0.78	0.81	0.29	0.51
my-CCAPM	0.68	0.72	0.42	0.57
1971-2002				
CAPM	0.23	0.23	1.57	1.05
FF-CAPM e.	−1.30	−0.64	3.35	1.45
FF-CAPM b.	0.28	0.39	1.25	0.82
y-CAPM	−0.30	−0.12	2.28	1.22
CCAPM	0.67	0.67	0.66	0.77
HCAPM	0.38	0.47	1.08	0.88
DCAPM	0.46	0.54	0.94	0.77
cay-CCAPM	0.58	0.64	0.74	0.75
cdy-CCAPM	0.28	0.39	1.25	1.03
my-CCAPM	0.30	0.40	1.22	1.02
Panel B: Quarterly Data				
1953.1-2002.4				
CAPM	0.044	0.044	0.12	0.27
FF-CAPM e.	−0.4	−0.0021	0.12	0.26
FF-CAPM b.	0.5	0.57	0.052	0.18
y-CAPM	−1.3	−1	0.25	0.43
CCAPM	0.36	0.36	0.079	0.25
HCAPM	0.36	0.45	0.067	0.22
DCAPM	0.9	0.91	0.011	0.087
cay-CCAPM	0.21	0.32	0.083	0.25
my-CCAPM	0.29	0.39	0.075	0.25
1953.1-2002.4				
CAPM	0.017	0.017	0.19	0.38
FF-CAPM e.	−0.2	0.14	0.16	0.31
FF-CAPM b.	0.14	0.26	0.14	0.31
y-CAPM	−1.4	−1.1	0.39	0.55
CCAPM	0.34	0.34	0.12	0.3
HCAPM	0.45	0.53	0.089	0.24
DCAPM	0.87	0.89	0.021	0.11
cay-CCAPM	0.2	0.31	0.13	0.3
my-CCAPM	0.18	0.3	0.13	0.32

Table 6: *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, for the *HCAPM* and *DCAPM* and $\lambda_{c,x}$, for the $cay-CCAPM$, $cdy-CCAPM$ and $my-CCAPM$, 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* is $myfa$. We used 4 lags to estimate the spectral density matrix with annual data and 12 lags with quarterly data.

Model	λ_c	$\lambda_{c,x}$	*p value*	R^2_{adj}	R^2	*mspe*	*mape*
Panel A: Annual Data							
1952-2002							
CCAPM.	7.97		1.00	0.96	0.96	0.99	0.74
s.e.	[1.12]						
HCAPM.	5.45	0.73	1.00	0.96	0.96	0.94	0.74
s.e.	[0.50]	[0.07]					
DCAPM.	5.46	4.34	1.00	0.96	0.96	0.96	0.77
s.e.	[0.53]	[0.45]					
cay-CCAPM.	8.09	1.73	1.00	0.96	0.96	0.97	0.73
s.e.	[0.93]	[0.20]					
cdy-CCAPM.	7.76	1.87	1.00	0.96	0.96	0.99	0.74
s.e.	[1.11]	[0.27]					
my-CCAPM.	5.84	1.03	1.00	0.99	0.99	0.37	0.51
s.e.	[0.84]	[0.13]					
1971-2002							
CCAPM.	5.80		1.00	0.87	0.87	2.73	1.20
s.e.	[0.15]						
HCAPM.	2.91	0.76	1.00	0.93	0.93	1.56	0.96
s.e.	[0.02]	[0.01]					
DCAPM.	3.94	2.88	1.00	0.88	0.88	2.57	1.20
s.e.	[0.01]	[0.00]					
cay-CCAPM.	5.01	1.06	1.00	0.88	0.88	2.53	1.24
s.e.	[0.08]	[0.01]					
cdy-CCAPM.	5.97	1.49	1.00	0.88	0.89	2.47	1.22
s.e.	[0.22]	[0.06]					
my-CCAPM.	4.51	0.64	1.00	0.92	0.92	1.75	1.16
s.e.	[0.22]	[0.04]					
Panel B: Quarterly Data							
1953.1-2002.4							
CCAPM.	1.24		1.00	0.96	0.96	0.04	0.19
s.e.	[0.14]						
HCAPM.	1.27	0.16	0.99	0.96	0.96	0.04	0.17
s.e.	[0.15]	[0.02]					
DCAPM.	2.15	1.24	1.00	0.98	0.98	0.03	0.13
s.e.	[0.47]	[0.23]					
cay-CCAPM.	1.27	0.32	0.99	0.96	0.96	0.04	0.19
s.e.	[0.15]	[0.04]					
my-CCAPM.	1.64	0.12	1.00	0.97	0.97	0.04	0.18
s.e.	[0.20]	[0.02]					
1971.1-2002.4							
CCAPM.	1.22		1.00	0.92	0.92	0.064	0.19
s.e.	[0.09]						
HCAPM.	1.31	0.15	1.00	0.94	0.94	0.05	0.17
s.e.	[0.12]	[0.03]					
DCAPM.	2.38	0.96	1.00	0.99	0.99	0.024	0.13
s.e.	[0.43]	[0.16]					
cay-CCAPM.	1.26	0.36	1.00	0.92	0.92	0.064	0.19
s.e.	[0.09]	[0.03]					
my-CCAPM.	1.87	0.15	1.00	0.95	0.95	0.04	0.18
s.e.	[0.19]	[0.02]					

Table 7: Risk Price Estimates Managed Currency Portfolios

GMM estimates using Managed Currency portfolios as test assets. Panel A reports the risk prices for the factor models and Panel B reports the results for consumption-based models. We consider two samples with quarterly data. In the first sample we use 8 countries, in the second sample we use 11 countries. The collateral measure for the *my-CCAPM* is $myfa$. For the consumption-based models, the first column contains λ_c, the second column λ_x, for the *HCAPM* and *DCAPM* and $\lambda_{c,x}$, for the $cay - CCAPM$, $cdy - CCAPM$ and $my - CCAPM$, where x denotes the scaling factor. For the return-based models, 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 used 12 lags to estimate the spectral density matrix.

Model	λ_1	λ_2	λ_3	*p value*	R^2_{adj}	R^2	*mspe*	*mape*
			Panel A: Factor Models					
			1956-2002					
CAPM	−10.30			0.52	−4.64	−4.64	73.93	7.52
s.e.	[2.67]							
FF-CAPM e.	−10.18	−12.46	7.54	0.59	−7.88	−5.34	83.09	7.91
s.e.	[4.29]	[6.87]	[5.54]					
FF-CAPM b.	−8.28	30.45		0.71	−2.30	−1.83	37.11	4.94
s.e.	[8.38]	[16.17]						
y-CAPM	−12.98	−14.63		0.52	−5.18	−4.30	69.40	7.36
s.e.	[6.16]	[6.78]						
			1971-2002					
CAPM	−14.42			0.93	−3.66	−3.66	138.06	9.76
s.e.	[3.23]							
FF-CAPM e.	−16.40	−8.93	10.39	0.92	−5.25	−4.00	148.21	10.51
s.e.	[5.86]	[5.47]	[5.58]					
FF-CAPM b.	−6.08	22.83		0.90	−1.88	−1.59	76.77	7.41
s.e.	[7.49]	[13.66]						
y-CAPM	−15.14	−17.47		0.79	−2.53	−2.17	94.04	7.53
s.e.	[5.87]	[6.35]						
			Panel B: Consumption-based Models					
Model	λ_c	$\lambda_{c,x}$		*p value*	R^2_{adj}	R^2	*mspe*	*mape*
			1956-2002					
CCAPM	5.20			0.81	−0.13	−0.13	14.80	2.86
s.e.	[1.71]							
HCAPM	10.00	0.53		0.73	−0.26	−0.08	14.17	2.80
s.e.	[8.76]	[0.41]						
DCAPM	0.26	1.94		0.60	0.18	0.30	9.22	2.31
s.e.	[0.33]	[0.68]						
cay-CCAPM	3.16	0.72		0.75	−0.22	−0.05	13.75	2.66
s.e.	[1.33]	[0.29]						
cdy-CCAPM	1.55	0.41		0.74	0.37	0.46	7.10	1.93
s.e.	[0.87]	[0.21]						
my-CCAPM	10.24	2.35		0.44	0.47	0.55	5.91	2.13
s.e.	[15.36]	[3.44]						
			1971-2002					
CCAPM	4.40			0.82	−0.13	−0.13	33.38	4.59
s.e.	[0.87]							
HCAPM	6.56	0.25		0.88	−0.03	0.07	27.57	3.86
s.e.	[3.72]	[0.15]						
DCAPM	0.33	1.58		0.85	0.30	0.37	18.64	3.43
s.e.	[0.34]	[0.37]						
cay-CCAPM	4.98	1.10		0.75	−0.24	−0.12	33.15	4.58
s.e.	[1.35]	[0.29]						
cdy-CCAPM	2.23	0.60		0.81	0.49	0.54	13.72	2.84
s.e.	[0.46]	[0.12]						
my-CCAPM	3.65	0.55		0.80	0.27	0.34	19.53	3.30
s.e.	[1.14]	[0.20]						

Table 8: Cross-Country Comparison for 8 Currency Portfolios: Annual

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. The sample includes all countries in a given year which are assigned a Quinn capital account liberalization index that exceeds 20. Standard errors in brackets were generated by bootstrapping 10.000 observations. We use annual data. The sample starts in 1962-2002.

		1	2	3	4	5	6	7	8
		Panel A: Annual Data							
UK	*Mean*	−3.31	−2.12	−2.46	−1.87	−1.36	−0.33	0.29	0.30
	Std.	10.12	10.92	11.19	10.70	7.90	9.29	9.10	14.92
	Sharpe Ratio	−0.33	−0.19	−0.22	−0.18	−0.17	−0.04	0.03	0.02
GER	*Mean*	−4.24	−1.97	−2.02	−1.66	−1.46	−0.16	0.68	0.50
	Std.	8.66	9.33	10.39	7.68	8.30	9.25	9.63	15.56
	Sharpe Ratio	−0.49	−0.21	−0.19	−0.22	−0.18	−0.02	0.07	0.03
JAP	*Mean*	−3.48	−3.13	−1.87	−1.62	−1.88	−0.12	0.53	0.36
	Std.	10.03	11.95	12.13	12.37	11.48	10.46	13.43	17.97
	Sharpe Ratio	−0.35	−0.26	−0.15	−0.13	−0.16	−0.01	0.04	0.02
		Panel B: Quarterly Data							
UK	*Mean*	−3.05	−0.89	−1.07	−0.03	−0.90	−1.19	1.58	−0.97
	Std.	16.28	17.63	14.72	15.12	13.97	15.21	12.56	20.35
	Sharpe Ratio	−0.19	−0.05	−0.07	−0.00	−0.06	−0.08	0.13	−0.05
GER	*Mean*	−3.08	−0.82	−1.30	0.12	−1.56	−0.41	1.66	−0.77
	Std.	16.73	18.62	17.10	13.44	15.74	16.69	14.49	22.92
	Sharpe Ratio	−0.18	−0.04	−0.08	0.01	−0.10	−0.02	0.11	−0.03
JAP	*Mean*	−2.68	−1.16	−0.89	0.42	−1.24	−0.37	1.98	−0.62
	Std.	19.31	20.34	20.56	19.40	19.63	20.24	19.24	24.98
	Sharpe Ratio	−0.14	−0.06	−0.04	0.02	−0.06	−0.02	0.10	−0.02

Table 9: International $CCAPM$ Risk Price Estimates for Currency Portfolios

GMM estimates using 8 currency portfolios as test assets. The first column contains γ, the second column λ_c. The sample starts in 1962-2002, except for France(1966), Italy (1971), the Netherlands (1978), Switzerland (1981) and Sweden (1964).

Model	γ	λ_c	*p value*	R^2	*mspe*	*mape*
			1962-2002			
US	42.38	19.96	0.55	0.61	0.66	0.70
s.e.	[6.31]	[25.85]				
UK	30.15	7.45	0.70	0.49	0.71	0.76
s.e.	[6.84]	[6.32]				
GER	21.18	13.68	0.70	0.55	0.99	0.83
s.e.	[3.48]	[6.81]				
CAN	41.13	12.93	0.76	0.64	0.61	0.74
s.e.	[6.12]	[12.76]				
FR	−7.37	−11.63	0.90	−0.35	2.42	1.27
s.e.	[11.40]	[18.07]				
JAP	15.62	3.68	0.71	0.53	0.94	0.70
s.e.	[2.52]	[1.22]				
ITA	23.31	4.66	0.75	0.71	0.51	0.52
s.e.	[3.45]	[1.98]				
NET	20.89	1.68	0.69	0.52	1.37	0.98
s.e.	[2.63]	[0.31]				
SWI	50.27	1.03	0.76	0.66	0.99	0.78
s.e.	[7.69]	[0.25]				
SWE	−6.28	−1.76	0.70	−0.01	1.57	0.98
s.e.	[3.50]	[3.50]				
AUS	31.44	3.35	0.61	0.70	0.53	0.55
s.e.	[5.52]	[1.92]				
pooled	41.73	30.76	0.19	0.81	0.35	0.47
s.e.	[6.65]	[13.34]				
			1971-2002			
US	31.16	4.20	0.56	0.48	1.07	0.78
s.e.	[5.53]	[1.79]				
UK	28.01	7.91	0.59	0.64	0.55	0.66
s.e.	[3.83]	[3.96]				
GER	22.29	14.66	0.63	0.38	1.36	1.05
s.e.	[4.31]	[7.58]				
CAN	35.02	5.49	0.69	0.55	0.98	0.91
s.e.	[5.05]	[2.21]				
FR	22.20	2.37	0.52	−0.04	2.03	1.19
s.e.	[4.68]	[0.80]				
JAP	15.60	1.89	0.71	0.28	1.65	0.93
s.e.	[3.99]	[0.73]				
ITA	23.31	4.66	0.75	0.71	0.51	0.52
s.e.	[3.45]	[1.98]				
NET	20.89	1.68	0.69	0.52	1.37	0.98
s.e.	[2.63]	[0.31]				
SWI	50.27	1.03	0.76	0.66	0.99	0.78
s.e.	[7.69]	[0.25]				
SWE	−6.66	−2.00	0.76	−0.07	2.00	1.07
s.e.	[5.30]	[2.29]				
AUS	30.22	1.63	0.67	0.57	0.90	0.72
s.e.	[6.26]	[0.73]				
pooled	43.26	14.24	0.07	0.64	0.70	0.63
s.e.	[5.24]	[4.37]				

Table 10: Consumption Growth Regressions

Results for the following time-series regression: $\Delta \log c^i_{t+1} = \alpha_0 + \alpha_1 \Delta \log c^{US}_{t+1} + \alpha_2 \;\; R^{£}_{t,t+1} - R^{\$}_{t,t+1} \;\; \Delta \log c^{US}_{t+1} + \epsilon_{t+1}$. 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

References

Alvarez, Fernando, Andrew Atkeson, and Patrick J. Kehoe, "Money, Interest Rates and Exchange Rates with Endogenously Segmented Markets," *Journal of Political Economy*, 2002, *110(1)*, 73–112.

Andrews, Donald W.K., "Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimation," *Econometrica*, 1991, *59* (1), 817–858.

Backus, David and Gregor Smith, "Consumption and Real Exchange Rates in Dynamic Economies with Non-Traded Goods," *Journal of International Economics*, 1993, *35*, 297–316.

___ **, Silverio Foresi, and Chris Telmer**, "Affine Models of Currency Pricing: Accounting for the Forward Premium Anomaly," *Journal of Finance*, 2001, *56.*

Bansal, Ravi and Magnus Dahlquist, "The Forward Premium Puzzle: Different Tales from Developed and Emerging Economies," *Journal of International Economics*, 2000, *51*, 115–144.

Census, Bureau, *Historical Statistics of the United States, Colonial Times to 1970*, U.S. Department of Commerce, 1976.

Chari, V.V., Patrick Kehoe, and Ellen McGrattan, "Can Sticky Price Models Generate Volatile and Persistent Real Exchange Rates?," *The Review of Economic Studies*, 2002, *69*, 533–564.

Cochrane, John, "Financial Markets and the Real Economy," *NBER Working Paper*, 2005, *11193.*

Cochrane, John H., *Asset Pricing*, Princeton, N.J.: Princeton University Press, 2001.

Engel, Charles and Kenneth D. West, "Exchange Rates and Fundamentals," *Journal of Political Economy*, 2005, *forthcoming.*

Fama, Eugene F., "Forward and Spot Exchange Rates," *Journal of Monetary Economics*, 1984, *14*, 319–338.

___ **and James D. MacBeth**, "Risk, Return, and Equilibrium: Empirical Tests," *The Journal of Political Economy*, 3 1973, *81*, 607–636.

___ **and Kenneth R. French**, "The Cross-Section of Expected Stock Returns," *Journal of Finance*, 1992, *47* (2), 427–465.

___ **and** ___ , "Common Risk Factors in the Rteurns on Stocks and Bonds," *Journal of Financial Economics*, 1993, *33*, 3–56.

Froot, Kenneth and Richard Thaler, "Anomalies: Foreign Exchange," *The Journal of Economic Perspectives*, 1990, *4(3)*, 179–192.

Gourinchas, Pierre-Olivier and Helene Rey, "International Financial Adjustment," 2003. Mimeo Berkeley University and Princeton University.

Hansen, Lars Peter, "Large Sample Properties of Generalized Method of Moments Estimators," *Econometrica*, 1982, *50*, 1029–54.

___ and R. Jagannathan, "Implications of Security Markets Data for Models of Dynamic Economies," *Journal of Policitical Economy*, 1991, *99*, 252–262.

___ and Robert J. Hodrick, "Forward Exchange Rates as Optimal Predictors of Future Spot Rates: An Econometric Analysis," *Journal of Political Economy*, 1980, *88*, 829–853.

Harvey, Campbell R., Bruno Solnik, and Guofu Zhou, "What Determines Expected International Asset Returns?," *Annals of Economics and Finance*, 2002, *3*, 249–298.

Hodrick, Robert, *The Empirical Evidence on the Efficiency of Forward and Futures Foreign Exchange Markets*, Chur, Switzerland: Harwood Academic Publishers, 1987.

Hollifield, Burton and Amir Yaron, "The Foreign Exchange Risk Premium: Real and Nominal Factors," 2001. Mimeo GSIA Carnegie Mellon University and Wharton School University of Pennsylvania.

Jagannathan, Ravi and Yong Wang, "Consumption Risk and the Cost of Equity Capital," *NBER Working Paper 11026*, 2005.

Lettau, Martin and Sydney Ludvigson, "Resurrecting the (C)CAPM: A Cross-Sectional Test When Risk Premia Are Time-Varying," *The Journal of Political Economy*, 2001, *109* (6), 1238–1287.

___ and ___ , "Expected Returns and Dividend Growth," *Journal of Financial Economics*, 2005, *forthcoming.*

Lewellen, John and Stefan Nagel, "The Conditional CAPM Does Not Explain Asset Pricing Anomalies," *Working Paper MIT Sloan*, 2003.

Lewis, Karen K., "Puzzles in International Financial Markets," in G. Grossman and K. Rogoff, eds., *Handbook of International Economics*, Amsterdam: Elsevier Science B.V., 1995, pp. 1913–1971.

Lucas, Robert E., "Asset Prices in an Exchange Economy," *Econometrica*, 1978, *46* (6), 1429–54.

Lustig, Hanno and Stijn Van Nieuwerburgh, "Housing Collateral, Consumption Insurance and Risk Premia: an Empirical Perspective," *forthcoming Journal of Finance*, 2005.

——— **and** ——— , "The Returns on Human Capital: Good News on Wall Street is Bad News on Main Street," *Working Paper UCLA and NYU Stern*, 2005.

Maddison, Angus, *The World Economy: A Millennial Perspective* Economic History Services, Paris: OECD, 2001.

Meese, Richard A. and Kenneth Rogoff, "Empirical Exchange Rate Models of the Seventies: Do They Fit Out of Sample?," *Journal of International Economics*, 1983, *14*, 3–24.

Mehra and Edward Prescott, "The Equity Premium: A Puzzle.," *Journal of Monetary Economics*, 1985, *15* (2), 145–161.

Parker, Jonathan and Christian Julliard, "Consumption Risk and the Cross Section of Expected Returns," *Journal of Political Economy*, 2005, *113*, 185–222.

Piazzesi, Monika, Martin Schneider, and Selale Tuzel, "Housing, Consumption and Asset Pricing," December 2002. Mimeo GSB University of Chicago, NYU and Anderson UCLA.

Quinn, Dennis, "The Correlates of Changes in International Finance Regulation," *American Political Science Review*, 1997, *91(3).*

Reinhardt, Carmen, Kenneth Rogoff, and Miguel A. Savastano, "Debt Intolerance," *NBER Working Paper 9908*, 2003.

Roussanov, Nikolai, "The Composition of Wealth, Conditioning Information and the Cross Section of Stock Returns," *Working Paper*, 2004.

Santos, Tano and Pietro Veronesi, "Labor Income and Predictable Stock Returns," *Review of Financial Studies*, 2005, *forthcoming.*

Sarkissian, Sergei, "Incomplete Consumption Risk Sharing and Currency Risk Premiums," *Review Of Financial Studies*, 2003, *16*, 983–1005.

Service, Global Credit Research Moody's Investors, "Sovereign Bond Defaults, Rating Transitions, And Recoveries (1985-2002)," *Global Credit Research*, 2003, *February.*

Singh, Manmohan, "Recovery Rates from Distressed Debt - Empirical Evidence from Chapter 11 Filings, International Litigation, and Recent Sovereign Debt Restructurings," *IMF Working Paper WP/03/161*, 2003, *August.*

Verdelhan, Adrien, "A Habit Based Explanation of the Exchange Rate Risk Premium," 2004. Mimeo University of Chicago.

——, "Sharpe Ratios in Foreign Exchange," 2004. Mimeo University of Chicago.

Yogo, Motohiro, "A Consumption-Based Explanation of Expected Stock Returns," *Journal of Finance*, 2005, *forthcoming.*

.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:

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

which produces the following expression for the risk premium:

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

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).

Developed Countries Our panel of developed countries includes 20 countries. We include each of the following countries for the dates noted in parenthesis: Australia (1953-2002), Austria (1960-1991), Belgium (1953-2002), Canada (1953-2002), Switzerland (1980-2002), Germany (1953-2002), Denmark (1976-2002), Spain (1985-2002), France (1960-2002), United Kingdom (1953-2002), Greece (1985-2002), Ireland (1969-2002), Italy (1953-2002), Japan (1960-2002), the Netherlands (1953-2002), Norway (1984-2002), New-Zealand (1978-2002), Portugal (1985-2002), Sweden (1955-2002), United States (1953-2002). The exchange and T-bill rates were downloaded from Global Financial Data. The maturity of the T-bill rates is 3 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. 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.

.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.

International Consumption Data The international 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.

Labor Income plus Transfers 1929-2002: Bureau of economic Analysis, NIPA Table 2.1, Aggregate labor income is the sum of wage and salary disbursements (line 2), other labor income (line 9), and proprietors' income with inventory valuation and capital consumption adjustments (line 10). Transfers is transfer payments to persons (line 16) minus personal contributions for social insurance (line 23). Prior to 1929, labor income plus transfers is 0.65 times nominal GDP. Nominal GDP Between 1929 and 2002, the ratio of labor income plus transfers to nominal GDP stays between .65 and .70 and equals .65 in 1929 and 1930. Nominal GDP for 1889-1928 is from Maddison (2001).

Number of Households For 1889-1945: Census (1976), series A335, A2, and A7. Household data are for 1880, 1890 1900, 1910, 1920, 1930, and 1940, while the population data are annual. In constructing an annual series for the number of households, we assume that the number of persons per household declines linearly in between the decade observations. For 1945-2002: U.S. Bureau of the Census, table HH-1, Households by Type: 1940: Present.

Price Indices All Items (p^a) 1890-1912: Census (1976), Cost of Living Index (series L38). 1913-2002: CPI (BLS), base year is 1982-84. In parenthesis are the last letters of the BLS code. All codes start by CUUR0000S. Total price index (p^a): All items (code A0). Shelter (p^h): Item rent of primary residence (code EHA). Food (p^c) 1913-2002: Item food (code AF1). Apparel (p^{app}) 1913-2002: Item apparel (code AA).

Aggregate Consumption

Total Consumption Expenditures C 1909-1928: Census (1976), Total Consumption Expenditures (series G470). The observations are for 1909, 1914, 1919, 1921, 1923, 1925, and 1927. For 1929-2002: Bureau of economic Analysis, NIPA table 2.2. Total Consumption expenditures is personal consumption expenditures (line 1).

Housing Services Consumption C_{rent} 1909-1928: Census (1976), Rent and Imputed Rent (series G477). The observations are for 1909, 1914, 1919, 1921, 1923, 1925, and 1927. For 1929-2002: Bureau of economic Analysis, NIPA table 2.2. Housing services consumption H is nominal consumption on housing services (line 14).

Food Consumption C_{food} 1909-1928: Census (1976), Food (series G471 + G472 + G473). The observations are for 1909, 1914, 1919, 1921, 1923, 1925, and 1927. For 1929-2002: Bureau of economic Analysis, NIPA table 2.2. Nominal consumption of food (line 7).

Apparel Consumption C_{cloth} For 1909-1928: Census (1976), Apparel (series G474). The observations are for 1909, 1914, 1919, 1921, 1923, 1925, and 1927. For 1929-2002: Bureau of economic Analysis, NIPA table 2.2. Nominal consumption of clothing and shoes (line 8).

Housing Expenditure share A It is computed in two ways. The nondurable consumption share $\alpha = 1 - A$

First, for 1909-2002, the housing expenditure share is computed as rent and imputed rent divided by total consumption expenditures minus rent and imputed rent and minus apparel. The observations are for 1909, 1914, 1919, 1921, 1923, 1925, and 1927. The cell entries for 1920, 22, 24, 26, and 28 are the average of the adjacent cells. The corresponding measure for the nondurable consumption share is $\alpha_1 = 1 - A_1$.

Second, for 1929-2002: The housing expenditure share is A is nominal consumption on housing services (line 14) divided by nominal consumption of non-durables (line 6) and services (line 13) minus clothing and shoes (line 8). The corresponding measure for the nondurable consumption share is $\alpha_2 = 1 - A_2$.

Real Per Household Consumption Growth dc It is computed in two ways. First, for 1922-2002, we construct *real* nondurable consumption, as total consumption deflated by the all items CPI minus rent deflated by the rent component of the CPI minus clothes and shoes deflated by the apparel CPI component. Per household variables are obtained by dividing by the number of households. The missing data for 1924, 26, and 28 are interpolated using Mehra and Prescott (1985) real per capita consumption growth. The growth rate dc_1 is the log difference multiplied by 100. Second, for 1930-2002, we define *real* nondurable and services consumption (NDS), as nondurable consumption deflated by the NIPA nondurable price index plus services deflated by the NIPA services price index minus housing services deflated by the NIPA housing services price index minus clothes and shoes deflated by the NIPA clothes and shoes price index. The basis of all NIPA price deflators is 1996=100. They are not the same as the corresponding CPI components from the BLS. Per household variables are obtained by dividing by the number of households. The growth rate dc_2 is the log difference of NDS multiplied by 100.

Rental Price Growth $d\rho$ It is computed in two ways. First, for 1913-2002 we use the ratio of CPI rent component to the CPI food component: $\rho = \frac{p^h}{p^c}$. The growth rate $d\rho_1$ is the log difference multiplied by 100. Second, for 1930-2002, we construct nominal non-durable consumption (non-durables plus services excluding housing services and excluding clothes and shoes) and real non-durable consumption (where each item is separately deflated by its own NIPA price deflator, basis 1996=100). The deflator for nondurable consumption is then the ratio of the nominal to the real non-durable consumption series. The relative rental price is then the ratio of the price deflator for housing services to the price deflator for nondurable consumption. The growth rate $d\rho_2$ is the log difference multiplied by 100.

Durable Consumption growth Δd Our durable consumption growth series is the one used by Yogo (2005). it is available from his web site at *http://finance.wharton.upenn.edu/ yogo/.* We construct the quarterly series following the methodology outlined by Yogo (2005).

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 (2005a)). 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/.* cdy is also available from the same web site.