Note on The Cross-Section of Foreign Currency Risk Premia and Consumption Growth Risk

Hanno Lustig and Adrien Verdelhan *
UCLA and Boston University

June 2007

1 Introduction

In our paper on "The Cross-Section of Currency Risk premia and Consumption Growth", we show that US consumption growth risk can explain predictable returns in currency markets. High interest rate currencies tend to appreciate, and hence US investors can earn positive excess returns by investing in these currencies, but we show this comes at the cost of bearing more US aggregate risk. To show this, we sort currencies into portfolios based on their interest rate, because this averages out changes in exchange rates that are purely idiosyncratic. On average, the high interest rate currency portfolio produces a return that is 5 percentage points larger per annum than the return on the low interest rate currency portfolio. This spread in returns is due to the different aggregate risk characteristics of these currency portfolios. In our paper, we interpret this common risk factor in currency risk markets using standard theory. We find that US aggregate consumption growth risk explains a large share of the variation in average returns on these currency portfolios, because the consumption betas for low interest rate currencies are smaller than the consumption betas for high interest rate currencies. In other words, high interest rate currencies do not depreciate as much as the interest gap on average, but these currencies tend to depreciate in bad times for a US

*We would like to thank Andy Atkeson, François Gourio, Christian Hellwig, Bob King, Nikolai Roussanov and Stijn VanNieuwerburgh for comments.

investor, who in turn receives a positive excess return in compensation for taking on this risk.

Our model is a standard-representative agent model that allows for non-separable utility from non-durable and durable consumption, and for non-separable utility over time. In Lustig and Verdelhan (2007), our analysis proceeds in two steps. First, as is standard in modern macro-economics, we calibrate the actual model, borrowing the structural parameters from Yogo (2006), who estimates these parameters on stock returns and macroeconomics data. We compute the pricing errors implied by the representative agent's Euler equation, evaluated over the sample for each of the eight currency portfolios. These results are shown in table 4 (section I.E) of the paper. When confronted with the post-war sample of foreign currency returns and US aggregate consumption growth, the representative agent demands a much higher risk premium on the high interest rate currency portfolio than on the low interest rate one The benchmark model explains 68 % of the variation in returns. This finding alone disproves the common claim that the forward premium puzzle *cannot* have a risk-based explanation (see Froot and Thaler (1990) for an earlier version of this argument and Burnside, Eichenbaum, Kleshchelski and Rebelo (2006) for a recent version).

Second, as is standard in empirical finance, we linearize the model (in section II of the paper), and we estimate the factor betas for this linearized model by regressing the currency portfolio returns on the three factors (non-durable, durable consumption growth and the market return). Then, we regress average returns on these betas to estimate the risk prices. This exercise confirms our earlier results. The risk prices of non-durable and durable consumption are large, and in-line with what we and others have found using different test assets (like stocks and bonds). Finally, our paper concludes by explaining why low interest rate currencies tend to appreciate when US consumption growth is lower than average.

Burnside's comments In his comment on our paper, Burnside (2007) (henceforth Burnside) replicates our point estimates for the risk prices in the linear model using only currency portfolios as test assets, and he agrees that the consumption betas line up with the returns on these currency portfolios. In other words, there is no question consumption risk is priced if you accept the consumption betas in our sample. Instead, Burnside questions how accurately these betas are measured.

As a result, the debate has shifted away from the claim that risk premia *cannot* explain the forward premium puzzle –we have shown that the sample moments of consumption growth and currency returns do support a risk-based explanation– to a debate about how accurately these sample moments are measured.

More specifically, Burnside questions the conclusion of our paper by claiming (1) that there is no statistical evidence that aggregate consumption growth risk is priced in currency markets and (2) that our definition of the measure of fit overstates our results. We briefly compare the evidence we report against his claims, starting with the first claim.

1. Burnside claims there is no statistical evidence that aggregate consumption growth risk is priced in currency markets.

    (a) In section IV.C of our paper, we show that the risk prices we obtain on currency excess returns are similar to those obtained when estimating the same model on other test assets like equity and bonds, even though these currency returns are not spanned by the usual factors like value and size.

    (b) Burnside argues that the price of consumption risk estimated on currency portfolios is not significantly different from zero once you correct for the fact that the betas are estimated in the first step of this procedure.[1] Burnside does not discuss the standard errors we obtained by bootstrapping samples from the observed consumption and return data that we report in section IV.C of our paper. These standard errors take into account the two steps and the small sample size. Using these bootstrapped standard errors, the price of durable consumption growth risk is significant at the 5 % level. In section 2 of this note, we briefly review the evidence reported in our paper and we also present some additional evidence from Generalized Least Squares (GLS) and Generalized Method of Moments estimates that were left out of the published version. All the evidence indicates that the price of consumption risk is statistically significant.

    (c) The difference between the consumption betas on the low and high interest rate portfolios reported in Table 6 of the paper is economically significant: there is at least a 100 basis points spread between the (univariate) non-durable and durable consumption betas on the first and the seventh portfolio over the entire sample; the spread increases to 150 basis points in the post-Bretton Woods sample. This is large, because we estimate that an asset with a non-durable consumption beta of one (or 100 basis points) earns a risk premium of 2% per annum; 4% per annum for durable consumption. In addition, the spread is statistically significant. The durable consumption beta on the seventh portfolio is about 2 standard errors removed from the one on the first portfolio in both sub-samples.

[1] These market prices of risk are estimated using a standard two-step procedure. In the first stage, we run a time-series regression of currency excess returns on the pricing factors (consumption growth in non durables and services, consumption growth in durables and stock market return) in order to estimate the betas. In the second stage, we run a cross-sectional regression of average currency excess returns on the betas, to estimate the market prices of risk for all the factors.

(d) In section II.D of our paper, we use the average interest rate gap with US for each portfolio as conditioning variables to estimate the conditional consumption betas, because this delivers more precise estimates if the consumption betas vary over time. These interest rate gaps predict currency returns, and hence these are natural variables to condition on. We show that the spread between the conditional consumption betas on low and high interest rate portfolios is large, and statistically significant. The low interest rate portfolios have negative consumption betas, because the exchange rates of low interest rate currencies depreciate in US recessions, and they depreciate by more as the foreign interest rate decreases. Burnside does not discuss these results.

(e) Burnside objects when we plug the sample moments of consumption growth and returns into the representative agent's Euler equation, or equivalently in the linearized model, when we plug in the sample estimates for the consumption betas, because these are not estimated accurately. Consumption betas are not estimated as precisely as return-based betas (e.g. CAPM betas). This is a well-known fact in finance, and we show in this note that is also true when one uses equity portfolios portfolios to test Yogo (2006)'s model. This is hardly surprising; there are few recessions in post-war data, and consumption is not measured accurately.

Consider the case of US stock returns. In our sample, the consumption beta of the return on the US stock market (the return on the value-weighted CRSP index) is 1.78. To explain the average annual stock market return of 6.95 % in the standard consumption-CAPM, the price of consumption risk has to be 3.90. This implies a very high coefficient of risk aversion. That is the equity premium puzzle as we know it. However, the t-stat on this consumption beta is only 1.04 in this sample. Do we conclude that the consumption beta of stock returns is really zero, and that the equity premium puzzle really is that the average excess return is positive?[2]

2. Burnside points out that the constant in the second stage of our regression is large and negative, and he argues that a risk-based explanation can be discounted because our model over-predicts the returns on the eight currency portfolios.

(a) The constant is large (about 300 basis points), but is not precisely estimated and it is not significantly different from zero. Since Burnside's comment is mainly about

[2]We disagree. Mehra and Prescott (1985) also took a different view. They never claimed that the equity premium is a puzzle because one cannot absolutely be sure that the correlation of consumption growth and stock returns is really positive.

estimation uncertainty, we are puzzled by the emphasis on the point estimate for the constant.

(b) This constant is difficult to estimate precisely because these currency returns (in units of US consumption) are all driven largely by the same swings in the dollar exchange rate. These swings can generate large across-the-board pricing errors for all test assets in small samples by driving a gap between investor's expected depreciation of the dollar and the actual sample average. If instead we use test assets that go long in high interest rate portfolios and short in low interest rate portfolios, we eliminate the effect of the dollar on returns. In section 3 of this note, we show that in this case the constant is much smaller and insignificant, as is to be expected, and that the model does even better on these test assets. Figure 1 plots the benchmark model's predicted excess returns (horizontal axis) against the realized excess returns for these seven test assets. The model's predicted excess returns on the vertical axis are a linear combination of the factor betas. On the left panel, we include a constant; on the right panel, we do not, and there is hardly any difference in the fit. The consumption-CAPM model explains 80 % of the variation in currency returns, regardless of whether we include a constant. Even though we agree that the model over-predicts the average (dollar) excess return on foreign currency investments, the model has no trouble explaining the spread between high and low interest currency returns and this what the forward premium puzzle is about. We could have written our entire paper about these zero cost investment strategies that go long in high and short in low interest rate currencies without changing a single line in the conclusion.

[Figure 1 about here.]

Outline The rest of this note is structured as follows. Section 2 of the paper addresses Burnside's first claim in detail by going over all the evidence in our paper. In section 3, we address the second claim.

Finally, we conclude our note by offering a preview of results presented in Lustig, Roussanov and Verdelhan (2007) in section 4. We show that a single risk factor, the spread between high and low interest currency returns, explains up to 80 % of the variation in these average returns. This is an aggregate risk factor that cannot be diversified away by US investors, and, according to standard APT (Arbitrage Pricing Theory) investors will be compensated for bearing this risk. That is what we find. We show that this factor's risk price is equal to its sample mean, as imposed by APT. Not surprisingly, these factor betas are estimated more precisely than the consumption betas, but the message is the same: we find that this risk factor is highly pro-cyclical.

The evidence presented in our paper, and in this note, presents a serious challenge to the view that risk is not priced in currency markets (see e.g. Burnside et al. (2006)). All the data used in Lustig and Verdelhan (2007) and in this note are available on-line.[3] As a result, all tables in the paper and in this note can be easily replicated. The figures and tables are in the appendix.

2 Estimating the Price of Consumption Risk and the Consumption Betas

Starting from the Euler equation and following Yogo (2006), we derive a linear factor model whose factors are non-durable US consumption growth Δc_t, durable US consumption growth Δd_t and the log of the US market return r_t^m. The US investor's unconditional Euler equation (approximately) implies a linear three-factor model for the expected excess return on portfolio j:

$$E[R^{j,e}] = b_1 cov\left(\Delta c_t, R_t^{j,e}\right) + b_2 cov\left(\Delta d_t, R_t^{j,e}\right) + b_3 cov\left(r_t^w, R_{t+1}^{j,e}\right). \tag{1}$$

Our benchmark asset pricing model, denoted *EZ-DCAPM*, is described by equation (1). This specification however nests the *CCAPM* with Δc_t as the only factor, the *DCAPM* with Δc_t and Δd_t as factors, the *EZ-CCAPM*, with Δc_t and r_t^m, and, finally the *CAPM* as special cases. This linear factor model can be restated as a beta pricing model, where the expected excess return $E[R^{j,e}]$ of portfolio $j$is equal to the factor price λ times the amount of risk of portfolio β^j:

$$E[R^{j,e}] = \lambda' \beta^j, \tag{2}$$

where $\lambda = \Sigma_{ff} b$ and $\Sigma_{ff} = E(f_t - \mu_f)(f_t - \mu_f)'$ is the variance-covariance matrix of the factors. The estimation procedure proceeds in two stages. In the first stage, we run a time-series regression of returns on the factors, to estimate the betas (β^j). In the second stage, we run a cross-sectional regression of average returns on the betas, to estimate the market prices of risk for all the factors (λ). Burnside argues that the estimated market prices of risk are not significant once one considers the sampling uncertainty introduced by the first-stage estimation of the betas. This is wrong.

[3] Data sets are available at http://www.econ.ucla.edu/people/faculty/Lustig.html, and at http://people.bu.edu/av/Research.html.

2.1 Prices of risk

We start by comparing the evidence in our paper on risk price estimates against this claim; in the paper, we report bootstrapped standard errors, Shanken-corrected standard errors, and Generalized Method of Moments (GMM) standard errors. In this note, we add Generalized Least Squares (GLS) standard errors.

Bootstrap In table 14, panel B (page 112 of the paper), we report the standard errors in brackets {} obtained by bootstrapping the whole estimation. We reproduce these results here in table 1 for the reader's convenience. These standard errors take into account the uncertainty in the first-stage of the estimation and the small sample size. They were generated by running the estimation procedure on 10.000 samples constructed by drawing both from the observed returns and factors with replacement under the assumption that returns and factors are not predictable. The first column reports the results with only currency portfolios as test assets. The market price of risk associated with consumption growth in durables is highly significant on currency portfolios. The point estimate is 4.7 and the standard error is 1.7 (Panel B, first column). If currency returns and consumption growth are independent, as Burnside claims, this bootstrapping exercise would have revealed this. Instead, it confirms that our results are significant.

[Table 1 about here.]

Shanken-correction Table 1 also reports the Shanken (1992)-corrected standard errors in parenthesis () –also in the paper. The Shanken correction, which is only valid asymptotically, produces substantially larger standard errors than the ones we generated by bootstrapping. Jagannathan and Wang (1998) actually show that the uncorrected Fama-MacBeth standard errors do not necessarily overstate the precision of the factor price estimates in the presence of conditional heteroskedasticity. We show in section III of the paper that conditional heteroskedasticity is the key to understanding these currency betas.

GMM In addition, panel A of Table 1 reports the 2-stage linear GMM estimates obtained on the same test assets. These standard errors also reflect the estimation uncertainty for these betas. Again, the price of non-durable consumption risk is significant (3.2 with s.e. of .9); likewise, the price of durable consumption risk is positive and significant (3.4 with a s.e. of 1.2). Burnside discards the GMM evidence as well, because he insists on estimating the mean of the factors, adding 3 separate moments. He obtains different point estimates. This means his GMM estimates of the factor means differ from the sample means, which is not a very appealing outcome. Yogo (2006) encounters a similar problem and he adjusts the weighting matrix to deal with this, as he explains in the appendix (p. 575). Because of these issues, our approach of not estimating the mean of the factors is actually more standard. For example, in table 8, page 1279, Lettau and Ludvigson (2001) report results from a GMM estimation of their linear factor model, and they also decide not to estimate the mean of the factors.

GLS Finally, in Table 2 of this document, we report the Generalized Least Squares (GLS) estimates that we left out of the published version of the paper. GLS estimators are more efficient than OLS estimators because they put more weight on the more informative moment conditions.[4] Clearly, for the *D-CAPM* and the *EZ-DCAPM*, the market price durable consumption risk is significant at the 5 % level, even when we use the asymptotic Shanken-correction that Burnside insists on. The price of non-durable consumption risk is around 3.2, with a Shanken-corrected s.e. of 1.8 and bootstrapped errors around 1.2. The price of durable consumption risk is around 5.15, with a Shanken-corrected s.e. of about 2.3 and bootstrapped errors around 1.7. The measures of fit are lower because GLS does not simply minimize the squared pricing errors; it minimizes the weighted sum. Table 3 reports similar results for the post-Bretton-Woods sub-sample. Burnside's claim that the risk prices are not statistically different from zero is not correct.

[Table 2 about here.]

[Table 3 about here.]

2.2 Consumption Betas

In his comment, Burnside stresses that the consumption betas are much less precisely estimated than return-based betas. We report univariate consumption betas and standard errors in table 6 of the published paper, reproduced here in Table 10. The difference between the consumption betas on the low and high interest rate portfolios reported in Table 6 of the paper is economically significant: there is at least a 100 basis points spread between the (univariate) non-durable and durable consumption betas on the first and the seventh portfolio over the entire sample; the spread increases to 150 basis points in the post-Bretton Woods sample. This is large, because we estimate that the price of non-durable consumption risk is 2% per annum; 4% per annum for durable consumption. In addition, the spread is statistically significant. The durable consumption beta on the seventh portfolio is about 2 standard errors removed from the one on the first portfolio in both sub-samples.

[4] For a comparison of estimators for beta pricing models, see Shanken and Zhou (2007).

[Table 4 about here.]

The (non-durable and durable) consumption betas for the seventh currency portfolios are significantly different from zero, but most of the others are not. We obviously agree that consumption betas are not estimated as precisely as return-based betas, but this is well known in finance, and certainly not a reason to give up on economic theory. To give an example, we estimated the factor betas on the Fama-French 25 equity portfolios sorted on size-and-book-to-market (see Table 12 in the appendix). Most of the consumption betas are not significantly different from zero. However, that does not mean Yogo (2006) reached the wrong conclusion in his paper. Asset pricing models are not tested by checking the t-stats on different betas. Should all of our currency portfolios have significant betas, even when they produce small and insignificant excess returns? In fact, in the example of the Fama-French 25 stock portfolios, the statistically significant market betas explain almost none of the variation in stock returns, while the durable consumption betas do. That is the whole point of Yogo (2006)'s paper, and we obtain similar results on currency portfolios.

Conditioning Information Lettau and Ludvigson (2001) have shown that bringing conditioning information to bear on the estimation produces more precise estimates of these consumption betas. This is why we condition on the portfolio's interest rate gap. It is a natural conditioning variable, because we know from the forward premium puzzle literature that interest rate gaps predict currency excess returns. The average interest rate gap with the US varies over time for each currency portfolio. We report conditional consumption betas in Table 7 and Figure 3 in the paper. Burnside does not discuss this evidence. We reproduce it in table 5 for the reader's convenience.

Note that we report conditional betas for changes in exchange rates. These are equivalent to conditional betas of log currency returns, because interest rates are known at the start of the period. We compute these betas by first running the standard uncovered interest rate parity regression for each portfolio, and then regressing the residuals on the factor and the factor interacted with interest rate gaps. The first panel reports the nondurable consumption betas, the second panel the durable consumption betas, the third panel reports the market betas. When the interest rate difference with the US hits the lowest point, the currencies in the first portfolio *appreciate* on average by 287 basis points when US non-durable consumption growth drops 100 basis points below its mean, while the currencies in the seventh portfolio *depreciate* on average by 96 basis points. Similarly, when US durable consumption growth drops 100 basis points below its mean, the currencies in the first portfolio appreciate by 174 basis points, while the currencies in the seventh portfolio depreciate by 105 basis points. Low interest rate currencies provide consumption insurance to US investors, while high interest rate currencies expose US investors to more consumption risk. As the interest rate gap closes on the currencies in the first portfolio, the low interest rate currencies provide less consumption insurance. For every 4 percentage points reduction in the interest rate gap, the non-durable consumption betas decrease by about 100 basis points.[5] These differences are not only economically significant, but statistically significant as well. The non-durable consumption betas on these two portfolios (1 and 7) are 4 standard errors apart.

An additional robustness check for the betas and market prices of risk comes from estimating the model on different classes of assets. We report the results of these asset pricing experiments in section IV.C of the published paper: our benchmark model can jointly account for the variation in currency and equity returns (as we show in figure 4 on page 109). We obtain similar market prices of risk on currency portfolios and on stock portfolios. [6]

[Table 5 about here.]

Finally, Burnside also claims the preference parameters implied by our estimates are nonsensical. We address this claim in a separate appendix (section A).

3 Estimating the intercept

We now turn to Burnside's second claim. Burnside stresses that the constant in the second stage of our regression is large and negative. He then argues that a risk-based explanation can be discounted because our model over-predicts the returns on all eight currency portfolios and that our R^2 overstates the fit of the model because it includes this constant. We first review the evidence and then turn to its implications. It turns out that the constant is not significantly different from zero; it is difficult to estimate because of large swings in the dollar, which affect all portfolios. However, the dollar does not affect the spread between portfolios, and when we estimate the model on spreads we obtain similar prices of risk and even higher R^2, with or without the constant.

3.1 Swings in Dollar

The constant in the second stage of our regression (λ_0) is negative (-2.9%) for the benchmark *EZ-DCAPM* model. This implies that a zero beta asset gets a negative excess return

[5] This table also shows our asset pricing results are entirely driven by how exchange rates respond to consumption growth shocks in the US, not by sovereign risk.

[6] Adding currency portfolios actually addresses one of the main criticism of the empirical finance literature: Daniel and Titman (2005) and Lewellen, Nagel and Shanken (2006) show that the Fama-French portfolios are highly correlated and thus do not put the bar high enough when testing models. Currency returns are not spanned by the usual size and value factors and thus constitute an additional challenge.

of 290 basis points. In other words, the model overpredicts the returns on all eight currency portfolios by 290 basis points. The uncorrected standard error on the intercept is 80 basis points. The Shanken-corrected standard error is 220 basis points, but in this case, Burnside only highlights the uncorrected standard errors. In the bootstrapping exercise, we find a standard error of 175 basis points. This clearly shows that the intercept is not significantly different from zero. Is this non-zero intercept a sufficient reason to reject a risk-based explanation of these currency returns?[7]

No, especially because the large swings in the dollar make it hard to accurately estimate the constant. The difference between the sample mean and the investor's expected rate of depreciation directly shows up in the intercept. The uncertainty that results from the dollar's fluctuations affects our estimates of the average excess return on *all* portfolios, but obviously not the spread between high and low interest rate portfolios. The latter is what we are interested in. We show that the intercept all but disappears when we look at the spreads. All these currency returns on the different portfolios have a large common component: the dollar's exchange rate vs. other currencies. When the dollar depreciates, this raises the returns on all portfolios, and when the dollar appreciates this lowers the returns on all portfolios, by the same amount for all portfolios. This makes it very hard to estimate the intercept accurately. Let E^i_{t+1} denote the exchange rate of currency i in dollars and let P_t denote the US price level. Lowercase letters denote logs. We use $\Delta e_{t+1} = (1/I)\sum_i \Delta e^i_{t+1}$ to denote the un-weighted average depreciation of the dollar at $t+1$. Estimating the intercept essentially amounts to estimating the expected rate of depreciation for the dollar: $E(\Delta e_{t+1} - \Delta p_{t+1})$. If the dollar appreciates more than expected in the sample, then the intercept λ_0 is negative, and the model over-predicts the returns on all foreign currency portfolios. Now, the standard deviation of changes in the deflated dollar exchange rate $(\Delta e_{t+1} - \Delta p_{t+1})$ is around 15 % per annum in our sample. Since we only have 50 observations, this means the standard error on the estimate of the expected rate of depreciation is about 2.12 % $(.15/\sqrt{50})$. So, the estimated intercept is only 1.36 standard errors (for the deflated dollar exchange rate) away from zero. A one standard error additional (average) depreciation of the dollar (by 2.12 percent) reduces the intercept to minus 78 basis points.[8]

[7]It is simply not the case that models with non-zero constants are rejected in the literature, as Burnside seems to imply. For example, in the *cay-CCAPM*, the constant λ_0 reported on page 1260 of Lettau and Ludvigson (2001) is positive and highly significant. For the three-factor Fama-French model, Shanken and Zhou (2007) find that the constant is positive and significant (see table 12 page 73).

[8]This problem does not arise when one uses stock returns as test assets. Stock returns do have a large common component (the market return), but different stocks or portfolios of stocks have different betas. There is no one-to-one mapping from the gap between the expected return on the market and its sample mean over the sample to changes in the intercept when estimating a model on a cross-section of stock portfolios.

The best way to estimate the intercept accurately is by considering different "home currencies" and the respective Euler equations of the "home" investors, all at the same time. This eliminates the common "dollar" component of course, but it requires more data (durable consumption series for other countries in our case). This also implies that forcing the intercept to be zero in the estimation only makes sense if you want to test whether the model can explain the average foreign currency return for US investors. That is not what our paper or the forward premium puzzle is about.

3.2 Long in High and Short in Low Interest Rate Currencies

Using the data we have posted on-line, we can simply test the model's performance on test assets that go long in the high interest rate currency portfolios and short in the first low interest rate portfolio. The returns on this strategy are given by the return on the high interest rate currency portfolio less the return on the lowest interest rate portfolio: $R_t^j - R_t^1$. The Euler equations should be satisfied as well for these zero cost strategies, but these returns are not affected by the dollar's fluctuations. This sidesteps the dollar issue altogether. If we are right, we should observe a smaller intercept λ_0. Table 6 reports the results for the Fama-Macbeth estimation of the linear factor models on these test assets. In the benchmark *EZ-DCAPM* (column 5), it drops from 290 basis points to -60 basis points, and it is not significantly different from zero. The R^2 is 81 %.[9] The risk prices of consumption are estimated precisely. The *DCAPM* in column 3 also has a small intercept (λ_0) of about 60 basis points. This model accounts for 60 % of the variation in the returns across these portfolios. We find similar results over the second sub-sample. Once you eliminate the effect of swings in the dollar by going long in high and short in low interest rate currencies, the intercept is essentially zero.

[Table 6 about here.]

No Constant Finally, Burnside argues that our definition of the cross-sectional regression's R^2 overstates the fit of the model, because we include the constant, even though this is the standard measure reported in this literature.[10] So, let us turn again to test assets that go long in high interest rate currency portfolios and short in the first portfolio. We redo the estimation *without a constant*, and, hence, we use Burnside's preferred measure of fit. Table 7 reports the results. The price of non-durable and durable consumption risk are

[9]This measure is based on the regression *with* a constant. The next paragraph considers the case *without* a constant. The R^2 drops to 79 %.

[10]For example, Lettau and Ludvigson (2001) report the standard R^2 as a measure of fit; we use the same measure.

significantly different from zero, and the model accounts for 79 % of the variation in these returns. Figure 1 compares the models estimated with and without the constant. It plots the benchmark model's predicted excess return (horizontal axis) against the realized excess return for these seven test assets. On the left panel, we include a constant; on the right panel, we do not. There is hardly any difference in the fit. The pricing error on the first and seventh portfolios is close to zero in both cases.

Another way to avoid this 'dollar problem' is to include the average excess return on all eight portfolios as a separate factor and estimate the model on all eight portfolios. This additional factor RX_{FX} absorbs the effect of the dollar variation in returns; there is no variation in the betas of this factor across portfolios, because all have the same dollar exposure. In this case, the model can be estimated on all eight test assets without a constant, and the risk price estimates are very similar to the ones we obtained on the same test assets without this additional factor, but including a constant.[11]

[Table 7 about here.]

GMM In table 8, we also report the GMM estimates obtained on these 7 test assets as well. The factors are demeaned. The consumption risk prices are 3.8 and 4.8 respectively. These are statistically significant. Again, the benchmark *EZ-DCAPM* model explains about 80% of the variation.

[Table 8 about here.]

The *EZ-DCAPM* model over-predicts the average (dollar) excess return on foreign currency investments by 290 basis points in our sample, but it has no trouble explaining the spread between high and low interest currency returns. This what the forward premium puzzle and our paper is about. Finally, the R^2 is not the only measure of fit we consider. The tables in the paper also report other measures of fit, like the mean absolute pricing error, and the p-value for a χ^2 test of the model.

4 Common Risk Factor in Currency Markets

We conclude by offering a preview of Lustig et al. (2007). In that paper, we show that one common risk factor explains most of the variation in currency excess returns. We apply the basic Fama-French technology by constructing two risk factors: the mean excess return on all the portfolios RX_{FX} and the spread between the returns on the high and low interest

[11] These results are reported in Table 14 in the appendix.

rate portfolio HML_{FX} –the equivalent of the HML factor from Fama and French (1993)). These two factors are the first two principal components of the foreign currency portfolio returns. As, we are about to show, the second risk factor HML_{FX} can account for most of the cross-sectional variation in average excess returns.

Table 9 shows the results of a cross-sectional regression of average returns on this risk factor's betas using annual data for two samples (1953-2002 and 1971-2002). This single risk factor explains between 57 % and 78 % of the cross-sectional variation in currency returns, depending on the sample. RX_{FX} adds little explanatory power. Table 10 shows the betas. Not surprisingly, these are estimated more precisely than the consumption betas, and they have the right pattern to explain the variation in average returns. Moreover, APT (arbitrage pricing theory) implies that the risk prices λ for these factors should be equal to their sample means ($\lambda = E[f_t]$), because these factors are returns. In particular, HML_{FX} is the return on a zero-cost strategy and hence should satisfy

$$E\left[HML_{t+1,FX}m_{t+1}\right] = 0.$$

This in turn implies that $E\left[HML_{FX}\right] = \lambda_{HML}$. In the first sample, the point estimate for the risk price is 4.45 %, compared to a sample mean of 5.32 %. In the second sample, the point estimate is 6.25 %, compared to a sample mean of 6.92 %. In both of these cases, the sample mean of HML_{FX} is within a one standard error band around the risk price. Note that the models with the second factor (column 2 and 4) have large negative constants offset by a large positive risk price for the second factor, because there is no variation in the RX_{FX} betas. This factor captures the effects of variations in the dollar and these portfolios all have the some dollar exposure. Table 15 in the appendix reports the results we obtained when no constant is included in the regression. These results are very similar when we include the second factor, because this second factor absorbs the effects of the fluctuations in the dollar.

This establishes two points:

1. the differences in average foreign currency returns (and hence changes in exchange rates) are driven mostly by a single risk factor, and hence are mostly about risk.

2. understanding the properties of this factor HML_{FX} is critical to understanding returns, and hence changes in exchange rates and interest rates.

So, let us examine the time series properties of this factor. Figure 2 plots HML_{FX} against durable consumption growth and the NBER recession dates. The risk factor is clearly procyclical. Table 11 reports the consumption betas in univariate regressions of HML_{FX} on non-durable and durable consumption growth. The consumption beta estimates vary between 1 over the entire sample and 1.5 in the post-Bretton woods sample; all are statistically significant at the 5 % confidence level.

[Table 9 about here.]

[Table 10 about here.]

[Table 11 about here.]

[Figure 2 about here.]

5 Conclusion

Our paper on "The Cross-Section of Currency Risk premia and Consumption Growth" demonstrated that consumption growth risk is priced in currency markets. To make this point, we use currency portfolios sorted on interest rates. These portfolios average out the idiosyncratic risk in exchange rate changes, and this produces a sharper picture of the relation between exchange rates, interest rates and risk factors. In our sample, low interest-rates currency portfolios have low consumption growth betas, high interest-rates currency portfolios have high consumption growth betas. This implies that the forward premium puzzle has a risk-based explanation. Verdelhan (2005) proposes a fully developed model that is consistent with these facts. In the last section of this note, we also showed that a single risk factor explains the spread in foreign currency returns. The betas of this risk factor are measured much more precisely than the consumption betas, and the estimated risk prices satisfy the restrictions implied by APT, but the final analysis is the same: this risk factor is highly pro-cyclical.

Burnside et al. (2006) argue that the predictable excess returns in currency markets are orthogonal to risk factors, but instead can be attributed to market frictions. To strengthen their case against a risk-based explanation, Burnside initiates a statistical debate in his note about the accuracy with which the sample moments of consumption and currency returns are measured. He argues the data are not informative about the relation between consumption growth and foreign currency returns. We disagree, and we have pointed out the parts of our paper that Burnside overlooked. We have also provided additional evidence in favor of a risk-based explanation based on factor betas that are measured very accurately.

Burnside is right in pointing out that the model seems to over-predict the average foreign currency return for US investor, but that is not what our paper is about, and it is not what the forward premium puzzle is about. Our paper is about the spread between high and low interest rate currency returns, and we have shown that the model explains about 80 % of the variation in these returns.

References

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

Burnside, Craig, "The Cross-Section of Foreign Currency Risk Premia and US Consumption Growth: A Comment," 2007. NBER Working Paper.

—— **, Martin Eichenbaum, Isaac Kleshchelski, and Sergio Rebelo**, "The Returns to Currency Speculation," 2006. NBER Working Paper 12489.

Daniel, Kent and Sheridan Titman, "Testing Factor Market Explanations of Market Anomalies," July 2005. Northwestern University and UT Austin.

Fama, Eugene F. and Kenneth R. French, "Common Risk Factors in the Returns 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.

Jagannathan, Ravi and Zhenyu Wang, "An Asymptotic Theory for Estimating Beta-Pricing Models Using Cross-Sectional Regressions," *Journal of Finance*, 1998, *53*, 1285–1309.

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

Lewellen, Jonathan, Stefan Nagel, and Jay Shanken, "A Skeptical Appraisal of Asset-Pricing Tests," *NBER Working Paper*, 2006, *12360.*

Lustig, Hanno and Adrien Verdelhan, "The Cross-section of Foreign Currency Risk Premia and US Consumption Growth Risk," *American Economic Review*, 2007, *97(1)*, 89–117.

—— **, Nick Roussanov, and Adrien Verdelhan**, "Risk Factors in Currency Markets," 2007. Working Paper Boston University.

Mehra, Rajnish and Edward C. Prescott, "The Equity Premium Puzzle," *Journal of Monetary Economics*, 1985, *15*, 145–161.

Shanken, Jay and Guofo Zhou, "Estimating and Testing Beta Pricing Models: Alternative Methods and their Performance in Simulations," *Journal of Financial Economics*, 2007, *84*, 40–86.

Verdelhan, Adrien, "A Habit Based Explanation of the Exchange Rate Risk Premium," 2005. Working Paper Boston University.

Yogo, Motohiro, "A Consumption-Based Explanation of Expected Stock Returns," *Journal of Finance*, 2006, *61(2).*

Table 1: Estimation of Linear Factor Models with 8 Currency Portfolios sorted on Interest Rates, 6 Equity Portfolios sorted on Size and Book to Market and 5 Bond Portfolios

	C	E	E/C	E/B	E/B/C
Factor Price					
Panel A: GMM					
Nondurables	2.372	2.732	2.537	0.822	2.006
	[0.846]	[1.192]	[0.723]	[0.877]	[0.486]
Durables	3.476	2.573	2.699	-0.562	1.386
	[1.204]	[1.942]	[0.985]	[1.418]	[0.662]
Market	10.204	12.216	13.238	8.380	9.566
	[7.868]	[5.869]	[4.075]	[6.072]	[3.472]
Stats					
MAE	1.170	1.384	1.400	1.128	1.286
$p-value$	0.068	0.629	0.781	0.795	0.409
Panel B: FMB					
Nondurables	2.194	4.276	3.757	2.467	2.445
	[0.830]	[0.945]	[0.567]	[0.786]	[0.507]
	(2.154)	(3.059)	(1.656)	(1.574)	(1.025)
	{1.343}	{3.725}	{1.143}	{1.496}	{0.926}
Durables	4.696	3.788	4.294	1.889	2.047
	[0.968]	[1.227]	[0.785]	[1.300]	[0.875]
	(2.518)	(3.973)	(2.292)	(2.595)	(1.756)
	{1.716}	{4.449}	{1.758}	{2.579}	{1.445}
Market	3.331	23.292	13.992	9.730	10.787
	[7.586]	[8.658]	[2.846]	[2.667]	[2.804]
	(19.754)	(28.057)	(8.613)	(5.857)	(6.092)
	{11.182}	{27.202}	{3.395}	{3.300}	{2.998}
Stats					
MAE	0.325	1.263	1.657	1.283	1.992
$p-value$	0.628	0.353	0.002	0.000	0.000

Notes: Panel A reports the 2-stage GMM estimates of the factor prices (in percentage points) using 8 annually re-balanced currency portfolios, 6 Fama-French benchmark portfolios sorted on size and book-to-market and 5 Fama bond portfolios (CRSP) as test assets. The sample is 1953-2002 (annual data). In the first stage, we use the identity matrix as the weighting matrix. In the second stage we use the optimal weighting matrix (no lags). The sample is 1953-2002 (annual data). The standard errors are reported between brackets. The factors are demeaned. The pricing errors correspond to the first stage estimates. Panel B reports the Fama-MacBeth estimates of the factor prices (in percentage points) using 8 annually re-balanced currency portfolios, 6 Fama-French benchmark portfolios sorted on size and book-to-market and 5 Fama bond portfolios (CRSP) as test assets. The sample is 1953-2002 (annual data). The standard errors are reported between brackets. The standard errors in parentheses include the Shanken correction. The standard errors in {} are generated by bootstrapping 10.000 times. The factors are demeaned. The last two rows report the mean absolute pricing error (in percentage points) and the p-value for a χ^2 test.

Table 2: GLS Estimation of Linear Factor Models with 8 Currency Portfolios sorted on Interest Rates

	CCAPM	DCAPM	EZ-CCAPM	EZ-DCAPM
Factor Prices				
Constant	-2.765	-3.414	-2.939	-3.390
	[0.784]	[0.805]	[0.797]	[0.809]
	(1.850)	(2.215)	(1.990)	(2.212)
	{1.521}	{1.656}	{1.691}	{1.996|]
Non-durables	3.134	3.004	3.290	2.953
	[0.659]	[0.660]	[0.672]	[0.680]
	(1.570)	(1.829)	(1.691)	(1.871)
	{1.237}	{1.236}	{1.334}	{1.348}
Durables		5.153		5.125
		[0.860]		[0.864]
		(2.384)		(2.382)
		{1.557}		{1.783}
Market			-1.817	-3.650
			[5.907]	[5.933]
			(14.958)	(16.421)
			{11.420}	{11.480}
Stats				
MAE	4.657	0.855	4.449	0.732
R^2	0.110	0.678	-0.033	0.728
$p-value$	0.561	0.996	0.559	0.991

Notes: This table reports the **GLS** estimates of the risk prices (in percentage points) using 8 annually re-balanced currency portfolios as test assets. The sample is 1953-2002 (annual data). The factors are demeaned. The OLS standard errors are reported between brackets. The Shanken-corrected standard errors are reported in (). The boostrapped errors are in {}. The last three rows report the mean absolute pricing error (in percentage points), the R^2 and the p-value for a χ^2 test.

Table 3: GLS Estimation of Linear Factor Models with 8 Currency Portfolios sorted on Interest Rates

	CCAPM	DCAPM	EZ-CCAPM	EZ-DCAPM
Factor Prices				
Constant	-2.853	-3.251	-2.833	-3.167
	[1.089]	[1.111]	[1.103]	[1.117]
	(2.295)	(2.430)	(2.339)	(2.535)
	{1.852}	{2.016}	{2.108}	{2.336}
Nondurables	3.060	3.043	3.081	3.191
	[0.682]	[0.682]	[0.708]	[0.710]
	(1.467)	(1.520)	(1.529)	(1.638)
	{1.182}	{1.276}	{1.248}	{1.383}
Durables		3.431		3.517
		[0.703]		[0.712]
		(1.576)		(1.653)
		{1.250}		{1.339}
Market			6.895	5.975
			[6.154]	[6.173]
			(13.448)	(14.383)
			{10.182}	{11.045}
Stats				
MAE	5.689	2.452	5.666	1.902
R^2	0.095	0.337	0.117	0.482
$p-value$	0.782	0.931	0.893	0.947

Notes: This table reports the **GLS** estimates of the risk prices (in percentage points) using 8 annually re-balanced currency portfolios as test assets. The sample is 1971-2002 (annual data). The factors are demeaned. The OLS standard errors are reported between brackets. The Shanken-corrected standard errors are reported in (). The bootstrapped errors are in {}. The last three rows report the mean absolute pricing error (in percentage points), the R^2 and the p-value for a χ^2 test.

Table 4: Estimation of Factor Betas for 8 Currency Portfolios sorted on Interest Rates

Portfolios	1	2	3	4	5	6	7	8
				Panel A: 1953-2002				
Non-durables	0.105	0.762	0.263	0.182	0.634	0.260	1.100	0.085
	[0.550]	[0.368]	[0.620]	[1.163]	[0.628]	[0.845]	[0.790]	[1.060]
Durables	0.240	0.489	0.636	0.892	0.550	0.695	1.298*	0.675
	[0.492]	[0.341]	[0.396]	[0.617]	[0.584]	[0.601]	[0.562]	[0.618]
Market	−0.066*	−0.027	−0.012	−0.119*	−0.000	−0.012	−0.056	0.028
	[0.037]	[0.058]	[0.037]	[0.056]	[0.054]	[0.054]	[0.060]	[0.118]
				Panel B: 1971-2002				
Non-durables	0.005	0.896	0.359	0.665	0.698	0.319	1.546	−0.461
	[0.679]	[0.512]	[0.805]	[1.445]	[0.746]	[1.060]	[1.020]	[1.287]
Durables	0.537	0.786	1.288*	2.032*	1.225*	1.359	2.183*	0.845
	[0.741]	[0.571]	[0.568]	[0.761]	[0.842]	[0.949]	[0.826]	[0.889]
Market	−0.106*	−0.099*	−0.026	−0.171*	−0.017	−0.007	−0.083	0.052
	[0.046]	[0.055]	[0.052]	[0.063]	[0.077]	[0.076]	[0.084]	[0.177]

Notes: Each column of this table reports OLS estimates of β^j in the following time-series regression of excess returns on the factor for each portfolio j: $R^{j,e}_{t+1} = \beta^j_0 + \beta^j_1 f_t + \epsilon^j_{t+1}$. The estimates are based on annual data. Panel A reports results for 1953-2002 and Panel B reports results for 1971-2002. We use 8 annually re-balanced currency portfolios sorted on interest rates as test assets. $*$ indicates significance at 5 percent level. We use Newey-West heteroskedasticity-consistent standard errors (reported in brackets) ; we use an optimal number of lags to estimate the spectral density matrix following Andrews (1991).

Table 5: Estimation of Conditional Consumption Betas for Changes in Exchange Rates on Currency Portfolios Sorted on Interest Rates

	1	2	3	4	5	6	7	8
				Panel A: Non-durables				
$\theta_1^{j,c}$	-2.87	-0.90	-0.94	1.17	0.83	0.58	0.96	-0.08
	[0.73]	[1.20]	[1.28]	[1.99]	[0.91]	[1.00]	[0.75]	[0.90]
$\theta_2^{j,c}$	0.27	0.18	0.10	-0.22	-0.16	-0.13	-0.04	-0.02
	[0.10]	[0.19]	[0.17]	[0.30]	[0.17]	[0.14]	[0.07]	[0.03]
				Panel B: Durables				
$\theta_1^{j,d}$	-1.74	-1.05	-0.68	0.99	0.36	0.55	1.05	-0.00
	[1.01]	[1.47]	[1.39]	[1.44]	[0.92]	[0.67]	[0.51]	[0.53]
$\theta_2^{j,d}$	0.18	0.18	0.15	-0.03	-0.03	-0.02	-0.00	-0.00
	[0.10]	[0.17]	[0.17]	[0.19]	[0.14]	[0.08]	[0.06]	[0.01]
				Panel C: Market				
$\theta_1^{j,m}$	-0.04	0.18	0.37	0.15	0.12	0.05	0.04	-0.06
	[0.13]	[0.19]	[0.14]	[0.24]	[0.10]	[0.09]	[0.06]	[0.08]
$\theta_2^{j,m}$	-0.01	-0.03	-0.05	-0.04	-0.03	-0.02	-0.02	0.00
	[0.02]	[0.02]	[0.02]	[0.03]	[0.02]	[0.01]	[0.01]	[0.00]

Notes: Each column of this table reports OLS estimates of $\theta^{j,k}$ in the following time-series regression of innovations to returns for each portfolio j (ϵ_{t+1}^j) on the factor f^k and the interest rate difference interacted with the factor: $\epsilon_{t+1}^j = \theta_0^{j,k} + \theta_1^{j,k} f_{t+1}^k + \theta_2^{j,k} \Delta\widetilde{R}_t^j f_{t+1}^k + \eta_{t+1}^{j,k}$. We normalized the interest rate difference $\Delta\widetilde{R}_t^j$ to be zero when the interest rate difference ΔR_t^j is at a minimum and hence positive in the entire sample. ϵ_{t+1}^j are the residuals from the time series regression of changes in the exchange rate on the interest rate difference (UIP regression): $E_{t+1}^j/E_t^j = \phi_0^j + \phi_1^j \Delta R_t^j + \epsilon_{t+1}^j$. The estimates are based on annual data and the sample is 1953-2002. We use 8 annually re-balanced currency portfolios sorted on interest rates as test assets. The pricing factors are consumption growth rates in non-durables (c) and durables (d) and the market return (w). The Newey-West heteroskedasticity-consistent standard errors computed with an optimal number of lags to estimate the spectral density matrix following Donald W. K. Andrews (1991) are reported in brackets.

Table 6: Long in High and Short in Low Interest Rate Currency Portfolios

	CCAPM	DCAPM	EZ-CCAPM	EZ-DCAPM
Factor Prices				
Constant	2.406	0.694	2.417	−0.641
	[0.901]	[0.869]	[0.845]	[0.848]
	(1.135)	(1.946)	(1.062)	(2.382)
	{0.999}	{1.213}	{1.263}	{1.691}
Nondurables	1.123	1.735	1.116	2.450
	[1.074]	[1.065]	[0.949]	[0.818]
	(1.369)	(2.394)	(1.211)	(2.307)
	{1.305}	{1.398}	{1.434}	{1.542}
Durables		4.129		5.144
		[1.225]		[1.042]
		(2.758)		(2.941)
		{1.819}		{2.217}
Market			1.757	4.699
			[7.978]	[8.190]
			(10.336)	(23.144)
			{12.598}	{12.751}
Parameters				
γ	52.274	90.704	44.392	123.622
	[50.004]	[55.429]	[46.192]	[38.382]
	(90.065)	(121.554)	(57.576)	(104.774)
σ			0.167	−0.035
			[0.887]	[0.035]
			(1.106)	(0.096)
α		1.140		1.124
		[0.613]		[0.487]
		(1.344)		(1.334)
Stats				
MAE	1.699	0.703	1.698	0.348
R^2	0.081	0.620	0.081	0.812
$p-value$	0.038	0.620	0.023	0.510

Notes: This table reports the **Fama-MacBeth** estimates of the risk prices (in percentage points) using 7 annually re-balanced currency portfolios as test assets. These test assets go long in the n-th currency portfolio and short in the first portfolio. The sample is 1953-2002 (annual data). The factors are demeaned. The OLS standard errors are reported between brackets. The Shanken-corrected standard errors are reported in (). The boostrapped errors are in {}. The last three rows report the mean absolute pricing error (in percentage points), the R^2 and the p-value for a χ^2 test.

Table 7: Long in High and Short in Low Interest Rate Currency Portfolios: No Constant

	CCAPM	DCAPM	EZ-CCAPM	EZ-DCAPM
Factor Prices				
Nondurables	4.617	2.302	4.021	2.016
	[1.060]	[0.848]	[1.005]	[0.915]
	(3.509)	(2.325)	(3.103)	(2.233)
	{1.881}	{1.617}	{1.905}	{1.524}
Durables		5.244		4.385
		[1.175]		[1.117]
		(3.221)		(2.729)
		{2.097}		{2.093}
Market			24.470	2.383
			[10.191]	[7.401]
			(31.500)	(18.151)
			{17.883}	{12.965}
Stats				
MAE	1.654	0.672	1.538	0.451
R^2	-0.700	0.578	-0.602	0.792
$p - value$	0.018	0.613	0.012	0.483

Notes: This table reports the Fama-MacBeth estimates of the risk prices (in percentage points) using 7 annually re-balanced currency portfolios as test assets. These test assets go long in the n-th currency portfolio and short in the first portfolio. The sample is 1953-2002 (annual data). The factors are demeaned. The OLS standard errors are reported between brackets. The Shanken-corrected standard errors are reported in (). The boostrapped errors are in {}. The last three rows report the mean absolute pricing error (in percentage points), the R^2 and the p-value for a χ^2 test.

Table 8: Long in High and Short in Low Interest Rate Currency Portfolios: GMM

	CCAPM	DCAPM	EZ-CCAPM	EZ-DCAPM
Factor Prices				
Nondurables	4.073	2.917	3.839	2.757
	[1.785]	[1.363]	[2.031]	[1.306]
Durables		4.886		4.864
		[2.128]		[1.866]
Market			0.171	0.261
			[0.141]	[10.834]
Parameters				
γ	193.44	147.45	514.39	139.53
	[84.77]	[67.01]	[452.25]	[63.22]
σ			-1.912	-0.009
			[2.839]	[0.026]
α		0.626		0.767
		[0.522]		[0.420]
Stats				
MAE	1.654	0.672	1.538	0.451
R^2	-1.392	0.568	-0.916	0.790
$p-value$	0.962	0.968	0.818	0.674

Notes: This table reports the 2-stage GMM estimates of the factor prices (in percentage points) using 7 annually re-balanced currency portfolios as test assets. These test assets go long in the n-th currency portfolio and short in the first portfolio. The sample is 1953-2002 (annual data). In the first stage, we use the identity matrix as the weighting matrix. In the second stage we use the optimal weighting matrix (no lags). The sample is 1953-2002 (annual data). The standard errors are reported between brackets. The factors are demeaned. The pricing errors correspond to the first stage estimates. The factors are demeaned. The last two rows report the mean absolute pricing error (in percentage points) and the p-value for a χ^2 test.

Table 9: High Minus Low

	1953-2002		1971-2002	
Factor Prices				
Constant	-0.856	-4.043	-1.257	-1.873
	[0.796]	[2.084]	[1.198]	[2.693]
	(0.934)	(2.696)	(1.480)	(3.325)
HML_{FX}	**4.454**	**4.328**	**6.252**	**6.181**
	[1.107]	[1.124]	[1.591]	[1.643]
	(1.655)	(1.780)	(2.486)	(2.535)
RX_{FX}		4.172		2.128
		[2.242]		[2.980]
		(3.020)		(3.901)
Stats				
MAE	1.052	0.722	0.751	0.739
R^2	0.577	0.709	0.790	0.793
$p-value$	0.241	0.197	0.788	0.696
Factor Mean				
HML_{FX}	**5.323**		**6.924**	
RX_{FX}		0.128		0.255

Notes: This table reports the Fama-MacBeth estimates of the risk prices (in percentage points) using 8 annually re-balanced currency portfolios as test assets. The factors are demeaned. The OLS standard errors are reported in brackets []. The Shanken-corrected standard errors are reported in (). The last three rows report the mean absolute pricing error (in percentage points), the R^2 and the p-value for a χ^2 test.

Table 10: Estimation of HML_{FX} Betas for 8 Currency Portfolios sorted on Interest Rates

Portfolios	1	2	3	4	5	6	7	8
				Panel A: 1953-2002				
	$-0.26*$	0.18	0.16	$0.36*$	0.09	$0.38*$	$0.74*$	0.11
	[0.11]	[0.13]	[0.15]	[0.14]	[0.15]	[0.14]	[0.11]	[0.14]
				Panel B: 1971-2002				
	$-0.23*$	0.11	0.20	$0.35*$	0.15	$0.41*$	$0.77*$	0.17
	[0.13]	[0.12]	[0.19]	[0.17]	[0.17]	[0.18]	[0.13]	[0.19]

Notes: Each entry reports OLS estimates of β^j in the following time-series regression of excess returns on the factor for each portfolio j: $R_{t+1}^{j,e} = \beta_0^j + \beta_1^j HML_{t+1,FX} + \epsilon_{t+1}^j$. The estimates are based on annual data. Panel A reports results for 1953-2002 and Panel B reports results for 1971-2002. We use 8 annually re-balanced currency portfolios sorted on interest rates as test assets. $*$ indicates significance at 5 percent level. We use Newey-West heteroskedasticity-consistent standard errors with an optimal number of lags to estimate the spectral density matrix following Andrews (1991).

Table 11: Estimation of Consumption Betas for HML_{FX}

	1953-2002	1971-2002
Nondurables	1.00	1.54
	[0.44]	[0.52]
Durables	1.06	1.65
	[0.40]	[0.60]

Notes: Each entry of this table reports OLS estimates of β_1 in the following time-series regression of the spread on the factor: $HML_{FX,t+1} = \beta_0 + \beta_1 f_t + \epsilon^j_{t+1}$. The estimates are based on annual data. The standard errors are reported in brackets. We use Newey-West heteroskedasticity-consistent standard errors with an optimal number of lags to estimate the spectral density matrix following Andrews (1991).

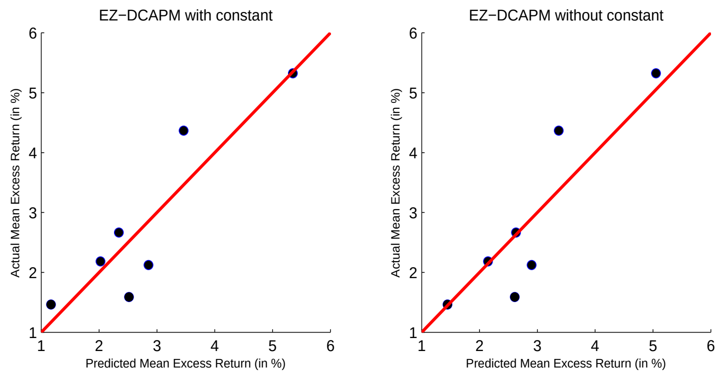

Figure 1: Short in Low and Long in High Interest Rate Currencies

This figure plots actual vs. predicted excess returns for 7 test assets. Currencies are sorted into 8 portfolios according to their interest rates. The 7 test assets are obtained by subtracting the returns on the first portfolio from the returns on the other portfolios. These test assets correspond to the following investment strategy: long in the high interest rate currency portfolios and short in the first currency portfolio. The data are annual and the sample is 1953-2002.

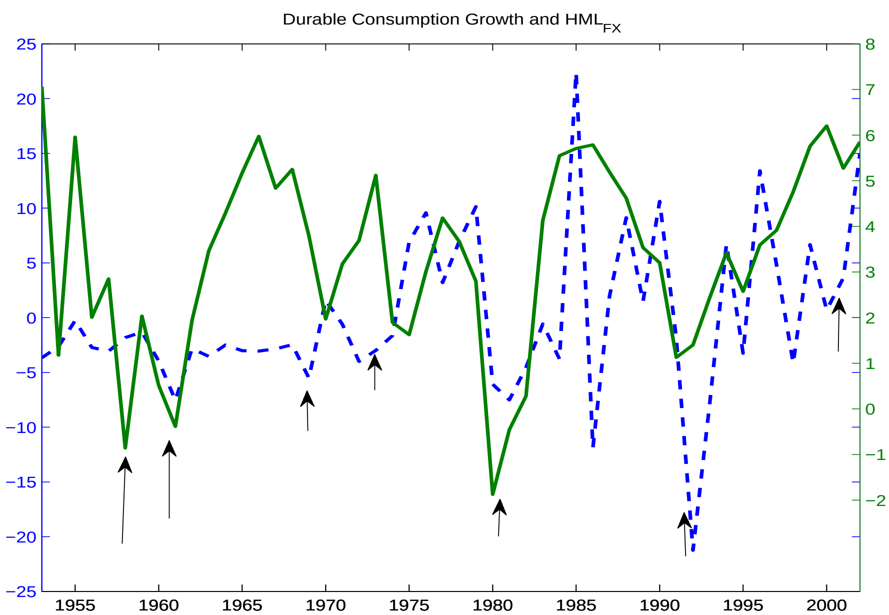

Figure 2: Durable Consumption Growth and HML_{FX}

The dotted line is HML_{FX}. The arrows indicate NBER recessions. The data are annual and the sample is 1953-2002.

A Preference Parameters

Finally, Burnside also claims that the preference parameters implied by our risk price estimates are nonsensical. In Table 13 in the appendix we report the preference parameter estimates corresponding to Table 5 in the paper, after for correcting for the typo in the published version of Yogo (2006)'s appendix. In the *EZ-DCAPM*, the risk aversion is high, around 110. The point estimate for the EIS is -.03, not significantly different from $1/\gamma$ which is the case of time-separable utility, and α is estimated to be larger than one, but the confidence interval includes values much smaller than one

We find very similar preference parameter estimates on the long-short test assets, reported in Table 6 and Table 8. The GMM point estimates for α are .6 in the *DCAPM* and .7 in the *EZ-DCAPM*.

B Additional Tables

[Table 12 about here.]

[Table 13 about here.]

[Table 14 about here.]

Table 12: Estimation of Factor Betas for 25 Fama-French Portfolios sorted on Size and Book-to-Market

Portfolios	1	2	3	4	5	6	7	8	9	10	11	12	132	14	15	16	17	18	19	20	21	22	23	24	25
													1953-2002												
Nondurables	−1.50	0.04	0.99	1.59	1.91	−1.30	−0.35	1.57	1.49	1.55	−0.94	1.13	1.92	2.18	2.21	−0.47	0.97	1.16	2.74	1.56	0.86	0.35	0.95	1.92	2.28
	[3.35]	[2.53]	[2.15]	[1.95]	[1.94]	[2.95]	[2.15]	[2.03]	[1.92]	[1.77]	[2.86]	[2.16]	[1.74]	[1.87]	[1.62]	[2.83]	[2.18]	[1.83]	[1.52]	[1.80]	[2.28]	[1.94]	[2.03]	[1.47]	[1.97]
Durables	−3.59	−3.46	−2.09	−2.04	−2.54	−3.86	−3.34	−1.98	−2.88	−3.57	−3.24	−2.71	−2.30	−2.50	−2.33	−2.74	−2.43	−2.90	−1.97	−2.92	−1.77	−2.29	−2.19	−1.75	−2.90
	[2.34]	[1.94]	[1.49]	[1.42]	[1.64]	[1.95]	[1.38]	[1.28]	[1.37]	[1.69]	[1.48]	[1.17]	[1.22]	[1.48]	[1.88]	[1.44]	[1.12]	[1.27]	[1.56]	[1.88]	[1.27]	[1.17]	[1.25]	[1.17]	[1.67]
Market	1.45	1.37	1.10	1.05	1.13	1.30	1.09	1.08	1.04	1.07	1.22	1.04	0.96	1.02	1.01	1.11	0.93	0.96	0.97	1.13	1.03	0.91	0.85	0.91	1.03
	[0.20]	[0.18]	[0.17]	[0.19]	[0.18]	[0.13]	[0.12]	[0.14]	[0.14]	[0.16]	[0.08]	[0.10]	[0.12]	[0.14]	[0.16]	[0.07]	[0.10]	[0.12]	[0.13]	[0.16]	[0.07]	[0.07]	[0.09]	[0.13]	[0.15]

Notes: Each entry reports OLS estimates of β^j in the following time-series regression of exces returns on the 25 FF equity portfolios on the factor for each portfolio j: $R^{j,e}_{t+1} = \beta^j_0 + \beta^j_1 f^i t + 1 + \epsilon^j_{t+1}$. The estimates are based on annual data. Results for 1953-2002. We use 25 annually re-balanced equity portfolios sorted on size and book-to-market. We use Newey-West heteroskedasticity-consistent standard errors with an optimal number of lags to estimate the spectral density matrix following Andrews (1991).

Table 13: Estimation of Linear Factor Models with 8 Currency Portfolios sorted on Interest Rates

	CCAPM	DCAPM	EZ-CCAPM	EZ-DCAPM
Factor Prices				
	-0.693	-3.057	-0.525	-2.943
	[0.954]	[0.839]	[1.046]	[0.855]
	(1.582)	(2.049)	(1.809)	(2.209)
	{1.538}	{1.659}	{1.743}	{1.751}
Nondurables	1.938	1.973	2.021	2.194
	[0.917]	[0.915]	[0.845]	[0.830]
	(1.534)	(2.245)	(1.476)	(2.154)
	{1.369}	{1.343}	{1.460}	{1.360}
Durables		4.598		4.696
		[0.987]		[0.968]
		(2.430)		(2.518)
		{1.653}		{1.695}
Market			8.838	3.331
			[7.916]	[7.586]
			(13.917)	(19.754)
			{12.336}	{11.216}
Parameters				
γ	90.191	102.778	92.757	111.107
	[42.676]	[54.374]	[41.869]	[38.910]
σ			-0.008	-0.032
			[0.460]	[0.037]
α		1.104		1.147
		[0.530]		[0.555]
Stats				
MAE	2.041	0.650	1.989	0.325
R^2	0.178	0.738	0.199	0.869
$p-value$	0.025	0.735	0.024	0.628

Notes: This table reports the Fama-MacBeth estimates of the risk prices (in percentage points) using 8 annually re-balanced currency portfolios as test assets. The sample is 1953-2002 (annual data). The factors are demeaned. The OLS standard errors are reported between brackets. The Shanken-corrected standard errors are reported in (). The last three rows report the mean absolute pricing error (in percentage points), the R^2 and the p-value for a χ^2 test.

Table 14: Estimation of Linear Factor Models with 8 Currency Portfolios sorted on Interest Rates -No Constant

	CCAPM	DCAPM	EZ-CCAPM	EZ-DCAPM
Factor Prices				
Nondurables	1.083	1.166	1.283	1.543
	[0.889]	[0.890]	[0.782]	[0.775]
Durables		4.856		5.267
		[1.221]		[1.144]
Market			11.379	0.057
			[8.143]	[8.071]
RX_{FX}	0.362	0.201	0.359	0.168
	[0.830]	[0.829]	[0.830]	[0.828]
Stats				
MAE	1.287	0.846	1.358	0.560
R^2	0.125	0.600	0.189	0.799
$p-value$	0.000	0.143	0.000	0.087

Notes: This table reports the Fama-MacBeth estimates of the risk prices (in percentage points) using 8 annually re-balanced currency portfolios as test assets. The sample is 1953-2002 (annual data). The factors are demeaned. We did not include a constant in the regression of average returns on betas. RX_{FX} -the additional factor- is the average excess return on all eight portfolios. The OLS standard errors are reported between brackets. The last three rows report the mean absolute pricing error (in percentage points), the R^2 and the p-value for a χ^2 test.

Table 15: High Minus Low - No Constant

	1953-2002		**1971-2002**	
Factor Prices				
HML_{FX}	2.891	**4.099**	3.924	**6.200**
	[1.810]	[1.157]	[2.793]	[1.636]
	(2.202)	(1.684)	(3.426)	(2.541)
RX_{FX}		0.253		0.310
		[0.831]		[1.279]
		(1.273)		(2.052)
Stats				
MAE	1.145	0.891	1.131	0.705
R^2	0.506	0.513	0.680	0.765
$p-value$	0.084	0.094	0.608	0.666
Factor Mean				
HML_{FX}	**5.323**		**6.924**	
RX_{FX}		0.128		0.255

Notes: This table reports the Fama-MacBeth estimates of the risk prices (in percentage points) using 8 annually re-balanced currency portfolios as test assets. The factors are demeaned. We did not include a constant in the regression of average returns on betas. The OLS standard errors are reported in brackets []. The Shanken-corrected standard errors are reported in (). The last three rows report the mean absolute pricing error (in percentage points), the R^2 and the p-value for a χ^2 test.