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

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Abstract

We find that the US consumption growth beta of an investment strategy that goes long in high interest rate currencies and short in low interest rate currencies is larger than one. These consumption beta estimates are statistically significant, contrary to what is claimed in Burnside (2007). With these consumption betas, the Consumption-CAPM can account for the average return on this investment strategy of 5.3 percent per annum with a market price of consumption growth risk that is about 5 percent per annum, lower than the price of consumption risk implied by the US equity premium over the same sample. When we formally estimate the model on currency portfolios in a two-step procedure, our estimate of the price of consumption risk is significantly different from zero, even after accounting for the sampling uncertainty introduced by the estimation of the consumption betas, while the constant in the regression of average returns on consumption betas is not significant.

JEL codes: F31,G12. **Keywords**: Exchange Rates, Asset Pricing.

1 Introduction

Our paper In our paper on "The Cross-Section of Currency Risk Premia and Consumption Growth" (cf Lustig and Verdelhan (2007)), we show that US consumption growth risk

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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 this comes at the cost of bearing more US aggregate risk. To analyze currency returns, we sort currencies into eight portfolios based on their interest rate, because this procedure 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. 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 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 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 portfolio. 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). Third, 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) 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. In this note, we address these two claims.

1. Burnside claims there is no statistical evidence that aggregate consumption growth risk is priced in currency markets and that currency excess returns do not co-vary with US consumption growth. This is his most important claim, and it is wrong.

    (a) Let us define HML_{FX} as the difference in returns between the high interest rate portfolio and the low interest rate portfolio. We focus on the seventh portfolio minus the first portfolio because this produces the largest spread (5.3 percent per annum). By construction, the consumption β of HML_{FX} is the difference between the consumption beta of the seventh and the first portfolio ($\beta^{HML} = \beta^7 - \beta^1$). So, we can simply test Burnside's claim by regressing HML_{FX} on consumption growth.

    The consumption growth beta of the return on the high minus the return on the low interest rate currency portfolio (HML_{FX}) is 1 for non-durable and durable consumption growth in a long sample starting in 1953. As a result, the Consumption-CAPM can account for the average return on this investment strategy of 5.3 percent per annum with a market price of consumption risk between around 5 percent per annum. This spread in betas is significant, and this market price of risk is not excessive. As a comparison, the consumption beta of the return on the US stock market (the return on the value-weighted CRSP index) is .97 over the same sample. To explain the average annual stock market excess return of almost 7 percent in the standard consumption-CAPM, the price of consumption risk has to be 7.1 percent per annum. This implies a substantial spread of $7.1 = 1.0 \times 7.1$

percent on the HML strategy, compared to 5.3 percent in the data.[1] As a result, if we simply use risk prices from stock markets, then the model predicts a slightly larger cross-section of currency returns than what we document in the data.[2]

Table 1: Estimation of Consumption Betas for HML_{FX}

Panel I: Simple Regression						
	β_c^{HML}	$p(\%)$	R^2	β_d^{HML}	$p(\%)$	R^2
	Panel A: Nondurables			*Panel B: Durables*		
1953 − 2002	1.00		4.04	1.06		9.07
	[0.44]	2.23		[0.40]	0.89	
1971 − 2002	1.54		8.72	1.65		14.02
	[0.52]	0.28		[0.60]	0.63	

Panel II: Multivariate Regression				
	β_c^{HML}	β_d^{HML}	χ^2	R^2
1953 − 2002	0.07	1.03		9.07
	[0.68]	[0.62]	9.40	
1971 − 2002	0.28	1.48		14.90
	[1.20]	[1.24]	14.15	

Notes: In Panel I, each entry of this table reports OLS estimates of β_1 in the following time-series regression of the spread on the factor: $HML_{FX,t+1} = \beta_0 + \beta_1^{HML} f_t + \epsilon_{t+1}$. $HML_{FX,t+1}$ is the return on the seventh minus the return on the first portfolio. 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). The p-values (reported in %) are for a t-test on the slope coefficient. The factor f_t is non-durable consumption growth (Δc) in the left panel and durable consumption growth (Δd) in the right panel. In Panel II, we report the multivariate regressions $HML_{FX,t+1} = \beta_0 + \beta_1^{HML}\mathbf{f}_t + \epsilon_{t+1}$. with $\mathbf{f}_t = [\Delta c_t, \Delta d_t]$. The χ^2 are for a Wald-test that the slope coefficients are zero.

In addition, the spread in consumption betas is statistically significant. In the simple regression case, the p-values (reported in percent) for a t-test are smaller than 2.5 percent in all of the four cases that we consider: non-durables in the 1953-2002 sample and the 1971-2002 sample, durables in the 1953-2002 sample and the 1971-2002 sample. Panel II in Table 1 reports the multivariate regression results of HML_{FX} on non-durables and durables. We report the χ^2-values for a Wald test that both of the consumption β's are zero. The Wald test statistic's p-values are both below 1 percent.[3] Why does Burnside reach a different conclusion? In

[1] The spread in consumption betas on currencies is about 1.5 in the post-Bretton Woods sample. The consumption beta of the return on the US stock market is 1.2 over the 1971-2002 sample. To explain the average annual stock market excess return of 5.75 percent over the same sample in the standard consumption-CAPM, the price of consumption risk has to be 4.9 percent. This implies a substantial spread of $7.4 = 1.5 \times 4.9$ percent on the HML strategy, compared to 6.9 percent in the data over the 1971-2002 sample.

[2] 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. Burnside does not discuss this evidence.

[3] Note that, even in the multivariate case, the beta on HML_{FX} is simply the difference in betas between

the multivariate case, the only case he considers, Burnside mistakenly focuses on the t-stats of the individual β's; the strong correlation of the consumption factors renders the individual coefficient estimates imprecise.[4] Obviously, two low t-stats on the consumption growth betas in the multiple regression do <u>not</u> imply that consumption growth does not co-vary with currency returns.

Finally, it is not the case that *all* of the consumption betas should be statistically different from zero. The interesting economic question is whether betas are different from each other, not different from zero. Since, for example, the average excess returns on the fifth and sixth currency portfolios are very close to zero, we should expect to see β's close to zero for these portfolios, and this is why we focus on the "corner portfolios".

(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.[5] Burnside does not discuss the standard errors 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 percent level. In 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.

Moreover, in section II.D of our paper, we use the average interest rate gap with the US for each portfolio as conditioning variables to estimate conditional consumption betas, because this delivers more precise estimates if 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 consump-

the high and low interest rate portfolios.

[4]This inference problem is commonly referred to as multi-collinearity in textbooks.

[5]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.

tion betas, because the exchange rates of low interest rate currencies depreciate in US recessions, and they depreciate by more as foreign interest rates decrease. Burnside does not discuss these results.

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. The constant is large (about 300 basis points), but is not precisely estimated and it is not significantly different from zero. Since the rest of Burnside's comment is exclusively about estimation uncertainty, we are puzzled by the emphasis on the point estimate for the constant without even mentioning the standard error.

   This constant is difficult to estimate precisely because these currency excess 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 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 percent of the variation in currency excess 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.

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

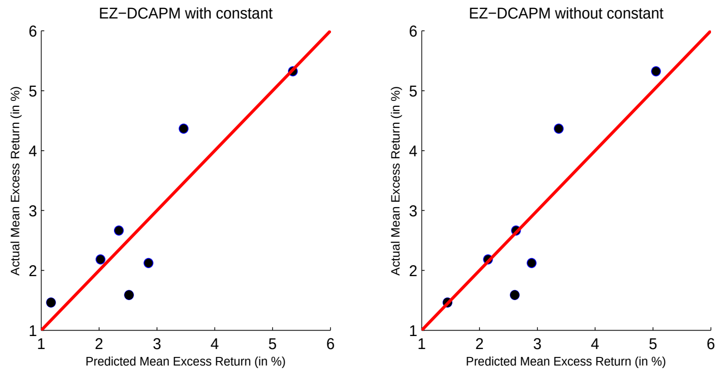

Figure 1: Short in Low and Long in High Interest Rate Currencies

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.

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.[6] As a result, all tables in the paper and in this note can be easily replicated.

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

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

portfolio j:

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

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 β^j:

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

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 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. In addition, he argues that the consumption β's are all indistinguishable from zero. This is wrong. We start with the consumption β estimates.

2.1 Consumption Betas

Currency Carry Trades As we have already shown in Table 1, the consumption growth betas on a simple currency carry trade strategy (borrowing in low interest rate currencies and lending in high interest rate currencies) vary between 1 over the entire sample and 1.5 in the post-Bretton woods sample. All these betas are statistically significant at the 5 % confidence level and economically meaningful. This simple fact contradicts ruins Burnside's argument.

All Currency Returns We report the univariate consumption betas and standard errors for all the currency portfolios in Table 6 of the published paper, reproduced here in Table 2. 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 with Burnside's comment 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 11 in the Appendix of this note). Most of the consumption betas are not significantly different from zero. However, that does not mean that 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.

Table 2: 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).

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 Lustig and Verdelhan (2007). Burnside does not discuss this evidence. We reproduce it in Table 3 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.[7] 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.

2.2 Prices of risk

We start by comparing the evidence in Lustig and Verdelhan (2007) on risk price estimates against Burnside's claim; in our 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 4 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

[7] 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.

Table 3: 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.

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.

Shanken-correction Table 4 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 4 reports the 2-stage linear GMM estimates obtained on the same test assets. These standard errors also reflect the estimation uncertainty for these

Table 4: 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. We consider currency portfolios (column 1), equity portfolios (column 2), equity and currency portfolios (column 3), equity and bond portfolios (column 4), and finally, equity, bond and currency portfolios (column 5). 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 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 the same test assets and sample. 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.

betas. Again, the price of non-durable consumption risk is significant (3.2 with a standard error of .9); likewise, the price of durable consumption risk is positive and significant (3.4 with a standard error 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 that 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 In Table 5 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.[8] 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 standard error of 1.8 and bootstrapped errors around 1.2. The price of durable consumption risk is around 5.15, with a Shanken-corrected standard error 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 6 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.

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.[9]

Finally, Burnside also claims that the preference parameters implied by our estimates are nonsensical. In Table 10 in the appendix we report the preference parameter estimates corresponding to Table 5 in the paper, after correcting for the typo in the published version of Yogo (2006)'s appendix. In the *EZ-DCAPM*, the risk aversion γ is high, around 110. The point estimate for the elasticity of intertemporal substitution is -.03, not significantly different from $1/\gamma$ which is the case of time-separable utility, and the utility weight on durable consumption α is estimated to be larger than one, but the confidence interval includes values

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

[9] 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.

Table 5: 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.

much smaller than one.[10] We find very similar preference parameter estimates on the long-short test assets, reported in Table 7 and Table 8. In the latter, the GMM point estimates for α are .6 in the *DCAPM* and .7 in the *EZ-DCAPM*.

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

[10] When the depreciation rate is 1 and the elasticity of substitution between the durable and non-durable goods is 1, then α is the share of durables in total consumption, but not in any other case because D is the stock rather than the expenditures of the durable good, cf footnote 2, page 6 in Yogo (2006).

Table 6: 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.

the evidence and then 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^2s, 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 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?[11]

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 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 $\Delta e_{t+1} = (1/I) \sum_{i=1}^{I} \Delta e^i_{t+1}$ 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.[12]

In theory, one could estimate the intercept accurately 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 and most durable consumption series, for example, are not available. This also implies that

[11] 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. Even for the three-factor Fama-French model, Shanken and Zhou (2007) find that the constant is positive and significant (see Table 12, page 73).

[12] 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 to changes in the intercept when estimating a model on a cross-section of stock portfolios.

forcing the intercept to be zero in the estimation only makes sense if one wants to test whether the model can explain the average foreign currency return for every foreign investors. That is not what our paper is about.

3.2 Long in High and Short in Low Interest Rate Currencies

Using the data that we have posted on-line, we can simply test the model's performance on currency carry trade strategies that go long in the high interest rate currency portfolios and short in the first low interest rate portfolio. 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 our interpretation of the constant is correct, we should observe a smaller intercept λ_0.

FMB Table 7 reports the results for the Fama-Macbeth estimation of the linear factor models on these test assets. In the benchmark *EZ-DCAPM* (column 5), the constant λ_0 drops from 290 basis points to -60 basis points, and it is not significantly different from zero. The R^2 is 81 %.[13] 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.

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 cross-section.

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

[13] 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 %.

this additional factor, but including a constant. These results are reported in Table 12 in the Appendix of this note.

As a result, 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.

Table 7: Long in High and Short in Low Interest Rate Currency Portfolios: FMB

	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.

Measures of fit 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.[14] 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 9 reports the results. The price of non-durable and durable consumption risk are 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 errors on the first and seventh portfolios are close to zero in both cases.

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.

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

Table 9: 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.

4 Conclusion

Our paper on "The Cross-Section of Currency Risk Premia and Consumption Growth" demonstrates 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.

Burnside et al. (2006) and Burnside, Eichenbaum and Rebelo (2007) argue that predictable excess returns in currency markets are orthogonal to risk factors, but instead can be attributed to market frictions (e.g bid-ask spreads and price pressure in Burnside et al. (2006), or time-varying adverse selection in Burnside et al. (2007)). 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.

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Table 10: 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 11: 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. The sample is 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 12: 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.