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

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ABSTRACT

The U.S. consumption growth beta of an investment strategy that goes long in high interest rate currencies and short in low interest rate currencies is large and significant. 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. The constant in the regression of average returns on consumption betas is not significant. In addition, the consumption and market betas of this investment strategy increase during recessions and times of crisis, when risk prices are high, implying that the unconditional betas understate its riskiness. We use the recent crisis as an example.

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An online appendix is available at:
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I Introduction

Generations of economists have been trained to think that investors ex ante should expect to earn a zero return on investments in high interest rate currencies that are funded in low interest rate currencies. In this view, large carry trade returns are an anomaly. Our work shows that this is a reasonable view of the world only if you are willing to believe that investors in currency markets take on aggregate risk without being compensated for it.[1]

Investments are risky if they offer low returns in bad times, when the typical investor experiences higher marginal utility growth than average. As an example, let us examine what transpired in currency markets during the current crisis.

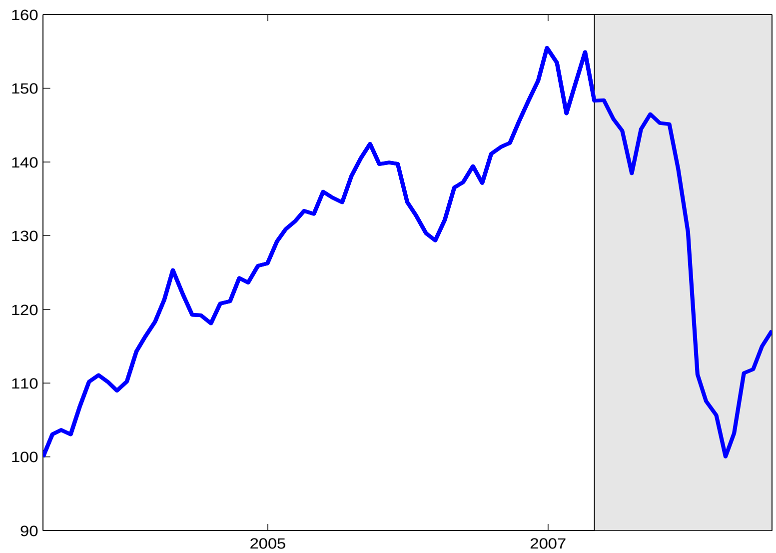

Figure 1: Currency Carry Trade Excess Returns

This figure presents an index of monthly excess returns on currency carry trades. Carry excess returns correspond to the Deutsche Bank Carry Harvest Index, which is available online at *http://www.dbfunds.db.com/Dbv/index.aspx* and starts in 1993. In Lustig and Verdelhan (2007), the sample ends in 2002. We focus here on the recent sample from December 2002 to June 2009. The index is equal to 100 at the end of 2002. The gray area, which starts in December 2007, corresponds to the latest recession.

The current recession, which started in December 2007, provides an interesting out-of-sample test

[1] Other authors have pursued a risk-based explanation of the forward premium puzzle. Our work is closest to Burton Hollifield and Amir Yaron (2001) and Campbell Harvey, Bruno Solnik and Guofu Zhou (2002). Hollifield and Yaron (2001) find some evidence that real factors, not nominal ones, drive most of the predictable variation in currency risk premia. Using a latent factor technique on a sample of international bonds, Harvey, Solnik and Zhou (2002) find empirical evidence of a factor premium that is related to foreign exchange risk. More recently, Michael J. Brennan and Yihong Xia (2006) show that their estimates of currency risk premia derived in affine term structure models satisfy the Eugene F. Fama (1984) necessary conditions for explaining the forward premium puzzle. Their paper builds on David Backus, Silverio Foresi and Chris Telmer (2001) who delineate the class of affine models that satisfy the Fama (1984) necessary conditions.

case for our claims. We start with the real side of the economy. In the fourth quarter of 2008, the United States recorded a 4.9% (annualized) drop in real personal consumption expenditures on nondurable goods, following a 5.6% (annualized) decrease in the third quarter. These growth rates are 3 standard deviations below the mean U.S. consumption growth rate in postwar data. Table I summarizes the evidence for 2008: -0.8% in nondurable consumption growth, -4.5% in durable expenditures growth, and -38.4% in the U.S. stock market return. The drop in durable consumption expenditures translates into a weak increase in the stock of durable goods (our measure of durable consumption growth) of 2.7%, much lower than the post–WWII average of 3.6%. It seems safe to say that the average investor in the United States experienced a higher than usual growth rate in marginal utility, regardless of which model is employed.

What happened in currency markets during the same period? High interest rate currencies depreciated and low interest rate currencies appreciated. As a result, returns on currency carry trades were low exactly in bad times. No computation needed here; market data are readily available. For example, the Deutsche Bank G10 Carry Harvest Index consists of long futures contracts on the three G10 currencies associated with the highest interest rates and of short futures contracts on the three G10 Currencies associated with the lowest interest rates. We use this index as one measure of carry trade returns because the corresponding exchange traded fund is easily available to any investor. The evolution of this index is clear. The current crisis has erased almost all of the carry trade gains made since the end of 2002. Figure 1 plots the evolution of this carry index (we normalized the index to 100 at the end of 2002). Carry traders first enjoyed a long period of steadily high returns. The index peaked at 155 in June 2007, but by the end of 2008 it was back down to 105. Thus, a 55% cumulative gain was followed by a like decrease. During the last two quarters of 2008, we witnessed a decrease of more than 31% of the Deutsche Bank carry trade index, a negative return equivalent to three standard deviations. The currency portfolios constructed by Lustig, Roussanov and Verdelhan (2008) (denoted LRV) cover more contracts than the Deutsche Bank Index. Even so, applying the long-short strategy still yielded a decrease of more than 10% if one invested in both developed and emerging countries.

In a reversal, the Deutsche Bank carry trade index recovered 16.8% during the first two quarters of 2009. Nondurable expenditures increased by 4.11%, and expenditures on durables increased by 7.4%. The U.S. stock market recovered 4.9 %. Again, if Burnside were right then the opposite pattern would have been just as likely: investors would have fled from the dollar and yen directly to the Australian dollar and the Icelandic Krone during the fall of 2008, ignoring the higher-yield currencies when the crisis abated.

One might argue this is not surprising; maybe all risky zero-cost investment strategies had similar returns during this episode. Far from it. Panel C shows the returns for small-minus-big (long in small, short in large stocks), high-minus-low (long in high book-to-market stocks, short in low book-to-market stocks) and momentum (long in winners over the past 12 months, short in losers over the past 12 months) equity investment strategies. All of these are zero-cost strategies that historically have produced large and positive average excess returns. All three did well during the last two quarters of 2008, while HML and momentum actually did very poorly during the first two quarters of 2009, when U.S. consumption rebounded. The exact opposite from what we see in currency markets. Clearly, U.S. investors with positions in the carry trade have incurred larger portfolio losses and consumption drops than those experienced by the average U.S. investor. Remarkably, this was not the case for other popular zero-cost investment strategies that have proven profitable: size, value and momentum stock investment strategies all posted positive returns during 2008.

Table I: Currency Excess Returns and Risk Factors — Subprime Crisis

	2007–2008	*2008:III–2008:IV*	*2009:I–2009:II*
		Panel A: Carry Trade Returns	
Deutsche Bank Returns	−28.8%	−31.4%	16.8%
LRV Returns — All Countries	−10.2%	−6.1%	3.6%
LRV Returns — Developed Countries	−31.1%	−20.1%	8.6%
		Panel B: Risk Factors	
Expenditure — Nondurables	−0.8%	−1.6%	4.1%
Expenditure — Durables	−4.5%	−10.2%	7.4%
Consumption — Durables	2.7%	0.4%	.9%
U.S. Stock Market	−38.4%	−29.9%	6.06%
		Panel C: Other Zero-Cost Investment Strategies	
SMB	4.2%	5.6%	10.3%
HML	1.0%	10.3%	−10.7%
Momentum	13.4%	.4%	−45.4%

Notes: The annual sample starts on 31 December 2007 and ends on 31 December 2008 (first column). The quarterly data correspond to the sum of the 2008 third- and fourth-quarter (second column) growth rates and to the sum of the first two quarters of 2009 (third column). The Deutsche Bank Carry Harvest Index is available online at *http://www.dbfunds.db.com/Dbv/index.aspx*. The *LRV* returns were computed by Lustig, Roussanov, and Verdelhan (2008), updated through June 2009. Section A in the separate appendix contains a detailed description of our series. SMB, HML and the momentum factor were taken from Kenneth French's web site.

In a previous paper entitled "The Cross-Section of Currency Risk Premia and Consumption Growth" (Hanno Lustig and Adrien Verdelhan 2007) we showed that, on average, high interest rate currencies are more exposed to aggregate consumption growth risk than low interest rate currencies in a sample with 81 currencies spanning 50 years of data. High interest rate currencies do not depreciate as much as the interest rate difference, and as a result, U.S. investors can generally earn positive excess returns by investing in these currencies. However, these high interest rate currencies tend to depreciate relative to low interest rate currencies in bad times for U.S. investors.

Furthermore, the average risk factor loadings of currency portfolios that we reported in Lustig and Verdelhan (2007) tend to understate the true risk inherent in the carry trade because the exposure of carry trade returns to aggregate risk factors increases dramatically during crisis episodes and recessions, exactly when the price of risk should increase in currency markets. To anyone who kept track of recent developments in currency markets, this may now seem obvious, but some economists still insist that "there is no relation between risk factors and currency returns" (see Craig Burnside, Martin Eichenbaum, Isaac Kleshchelski and Sergio Rebelo 2008).

Our Previous Paper In his comment on our previous paper, Burnside does not question our findings, yet he starts a debate about their statistical significance. In that paper, we departed from the literature by examining the cross-sectional relation between average returns on foreign currency investments and interest rates rather than examining the time-series relation.[2] To do so, we sorted currencies into eight portfolios based on their current interest rate. This approach (developed in Lustig and Verdelhan 2005) is helpful because it averages out changes in exchange rates that are purely idiosyncratic and hence are not priced in currency markets. We found that investors on average earn large excess returns simply by taking long positions in baskets of currencies with currently high interest rates and taking short positions in baskets of currencies with currently low interest rates, regardless of the history of interest rate differences for individual currency pairs. The average excess returns increase from the first portfolio, with currently low interest rate currencies, to the last portfolio, with currently high interest rate currencies. Moreover, we established that currencies sorted by interest rates share a lot of common variation, a necessary condition for a risk-based explanation. Finally, we tied this common variation to aggregate risk exposure — more specifically, to durable consumption growth, nondurable consumption growth, and the market return. Our paper was the first to make this point. The literature that precedes it focusses almost exclusively on currency-specific variation in bilateral exchange rates.

[2] See Lars P. Hansen and Robert J. Hodrick (1980) and Fama (1984) for earlier examples of time-series tests. Geert Bekaert and Hodrick (1993) investigate biases as an explanation of the forward premium puzzle. Hodrick (1987) and Karen K. Lewis (1995) provide extensive surveys and updated regression results.

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 U.S. aggregate consumption growth risk explains a large share of the variation in average returns for 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 U.S. 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 nonseparable utility from nondurable and durable consumption and also for nonseparable utility over time. In Lustig and Verdelhan (2007), our analysis proceeds in two steps. First, as is standard in modern macroeconomics, we calibrate the actual model. We adopted the structural parameters from Motohiro Yogo (2006), who estimates these parameters based on stock returns and macroeconomic 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 (see section I.E). When confronted with the postwar sample of foreign currency returns and U.S. 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, for example, Kenneth Froot and Richard Thaler 1990).

Second, as is standard in empirical finance, we linearize the model (in section II of the paper). We then estimate the factor betas for this linearized model by regressing the currency portfolio returns on the three factors (nondurable and durable consumption growth and the market return). Finally, we regress average returns on these betas in order to estimate the risk prices. This exercise confirms our earlier results. The risk prices of nondurable and durable consumption are large, and they are in line with what we and others have found using different test assets (like stocks and bonds). Our paper concludes by explaining why low interest rate currencies tend to appreciate when U.S. 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 currency portfolios as test assets. He agrees that the consumption betas are aligned with the returns on these currency portfolios. In other words, there is no question that consumption risk is priced if you accept the consumption betas in our sample. However, 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 —after all, 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.

Our Reply In this paper, we first address these two claims. We show that they have no merit with respect to our initial sample that ends in 2002. Furthermore, extending the sample to 2009 actually reinforces our points.

First claim: Burnside claims there is no statistical evidence that aggregate consumption growth risk is priced in currency markets or that currency excess returns do not covary with U.S. consumption growth. This is his most important claim, and it is false.

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% per annum). By construction, the consumption β of HML_{FX} is the difference between the consumption β of the seventh and the first portfolio ($\beta^{HML} = \beta^7 - \beta^1$). Hence, we can simply test Burnside's claim by regressing HML_{FX} on consumption growth.

The consumption growth beta of HML_{FX} is 1 for nondurable and durable consumption growth in a long sample starting in 1953. As a result, the consumption Capital Asset Pricing Model (CAPM) can account for the average return on this investment strategy of 5.3% per annum given a market price of consumption risk of around 5% per annum. This spread in betas is economically significant. As a benchmark, the consumption beta of the return on the U.S. stock market (the return on the value-weighted CRSP index) is 0.97 for the same sample. In order to explain the average annual stock market excess return of almost 7% in the standard consumption CAPM, the price of consumption risk has to be 7.1% per annum. This implies a substantial carry trade premium of $7.1 = 1.0 \times 7.1$ % on the HML strategy, compared to 5.3% in the data. We obtain similar results in the post–Bretton Woods sample. As a result, if we simply use risk prices from stocks then the model already predicts a sizable carry trade risk premium. Moreover, in section IV.C of our 2007 paper, we show that the risk prices we found for currency excess returns are similar to those obtained when estimating the same model on other test assets such as equity and bonds, even though these currency returns are not spanned by the usual factors of value and size. Burnside does not discuss this evidence.

In addition, the spread in consumption betas is statistically significant. In the simple univariate regression case, the p-values for a t-test are smaller than 2.5 % in all of the four cases that we consider: nondurables in the 1953-2002 and the 1971-2002 samples, durables in the 1953-2002 and the 1971-2002 samples. The multivariate regressions lead to the same conclusion. The Wald test statistic's p-values are both below 1%.[3]

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 betas close to zero for these portfolios. This is why we focus on the "corner portfolios".

Burnside then argues that the price of consumption risk estimated for currency portfolios is not significantly different from zero once you correct for the fact that the betas are estimated. In our previous paper, we report the standard errors obtained by bootstrapping samples from the observed consumption and return data (see section IV.C). 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. The separate appendix presents additional evidence from generalized least squares (GLS) and generalized method of moments (GMM) estimates that were omitted from the published version. All the evidence indicates that the price of consumption risk is statistically significant.

Second claim: 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 overpredicts the returns on the eight currency portfolios. The constant is large (about 300 basis points), but it is not precisely estimated and is not significantly different from zero. Since the rest of Burnside's comment is exclusively about estimation uncertainty, we are puzzled by his emphasis on the point estimate for the constant without emphasizing the large standard error.

[3] Why does Burnside reach a different conclusion? In the multivariate case, the only case he considers, Burnside mistakenly focuses on the t-statistics of the individual betas; the strong correlation of the consumption factors renders the individual coefficient estimates imprecise. This inference problem is commonly referred to as multicollinearity in textbooks. Obviously, two low t-stats on the consumption growth betas in the multivariate regression do not imply that consumption growth does not covary with currency returns.

This constant measures the price of that part of dollar risk that is not explained by our risk factors. All of our currency portfolios share the same loading on dollar fluctuations, so the cross-section of currency returns cannot be informative about the price of dollar risk. In other terms, the first principal component of currency returns is a dollar factor; all currency portfolios have essentially the same loadings on this factor. We never claim that consumption growth risk explains the returns on investing in a basket of all foreign currencies (i.e., the dollar risk premium), and no such claim is supported by the data.

If instead we use test assets that go long in high interest rate portfolios and short in low interest rate portfolios, we eliminate the dollar risk factor. These test assets are now dollar-neutral. In this case the estimated constant is much smaller and insignificant, as expected, and the model does even better on these test assets. Figure 2 plots the benchmark model's predicted excess returns against the realized excess returns for these seven test assets. The model's predicted excess returns are a linear combination of the factor betas. The left panel reflects the inclusion of a constant and the right panel does not, but there is hardly any difference in the fit. The consumption CAPM model explains 80% of the variation in currency excess returns regardless of whether we include a constant. Even though we agree that it overpredicts 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 is the essence of the forward premium puzzle. 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 of the conclusion.

Our paper is not about dollar risk. We agree with Burnside that consumption risk does not explain the average returns earned by U.S. investors on a basket of all foreign currencies, and we have never claimed that it did. Our focus is on the returns obtained by going long in high interest rate currencies and short in low interest rate currencies for this is how the carry trade is defined.

Additional Evidence To address Burnside's concerns about estimation uncertainty, we also bring new evidence to bear on the relation between aggregate risk and currency returns.

First, the statistical link between asset returns and macroeconomic factors is always weaker than the link between asset returns and return-based factors. That is why John H. Cochrane (2001) warns against pointless horse races between models with macroeconomic factors and those with return-based factors. Statistical uncertainty is not a license to ignore the link with the macroeconomy. However, an alternative is to use return-based risk factors. This reduces statistical uncertainty at the cost of economic content.

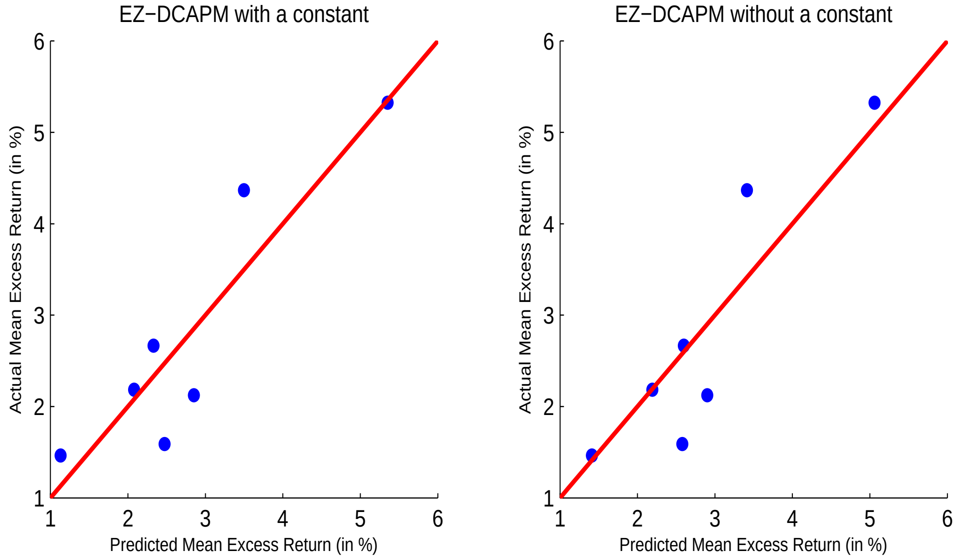

Figure 2: Short in Low and Long in High Interest Rate Currencies

This figure plots actual versus predicted excess returns for seven test assets. Currencies are sorted into eight portfolios according to their interest rates. The seven test assets are obtained by subtracting the returns on the first portfolio from the returns on the other portfolios. These test assets correspond to an investment strategy of going long in the high interest rate currency portfolios and short in the low interest rate currency portfolio. The risk factors are nondurable consumption growth, durable consumption growth and stock market returns. Risk prices are estimated following Fama and James D. MacBeth (1973). The data are annual from Lustig and Verdelhan (2007), and the sample is 1953–2002.

Return-based risk factors are more precisely measured, and their loadings and prices are precisely estimated. Lustig, Nick Roussanov, and Verdelhan (2008) use the first two principal components of the currency excess returns as risk factors to explain the cross-sectional variation in monthly currency returns. The second principal component captures the common time-series variation in exchange rates of currencies sorted by their interest rates. As a risk factor, it can account for 65% of the cross-sectional variation in average excess returns across our eight portfolios. This factor's market price of risk is precisely estimated and has t-stats well in excess of the 1% significance level. Clearly, average currency excess returns can be attributed to covariances between returns and common risk factors. Going back to the sample of countries we used in Lustig and Verdelhan (2007), we obtain similar results on long series of quarterly returns.

Second, by looking at monthly and daily return data, we show that the risk factor loadings of carry returns vary over time. We emphasize that these loadings tend to increase during recessions and other crisis episodes, when the price of risk tends to be higher. As a result, the average factor loadings reported in our previous paper tend to understate, not overstate, the true riskiness of these currency portfolios. For example, consider what happened during the recent crisis. At daily frequency, the correlation between one-month returns on the U.S. stock market and carry trade returns increased to 0.70 during the subprime crisis. We report similar findings for the LTCM crisis, the Mexican "Tequila" crisis, and the Brazilian/Argentine crisis.

Taking Stock Burnside replicates our point estimates of both the quantity and price of consumption growth risk and he concludes that "it is impossible to reject [our] model using formal statistical tests". He ignores the current crisis, which provides a striking counterexample to his own claims. In the face of all this evidence, what is the point of a lengthy digression on standard errors? In a series of papers, Burnside has consistently argued against risk-based explanations of the carry trade, and he has explored such other avenues as adverse selection and price pressure in attempting to explain the forward premium puzzle.[4] Most recently, however, Burnside et al. (2008) find some support for the disaster risk model of Robert Barro (2006). They conclude that "the same value of the stochastic discount factor that rationalizes the average payoffs to the carry trade also rationalizes the equity premium." Thus it seems like they finally ended up where we started off. We agree that neither price pressure nor adverse selection is a plausible explanation of the forward premium puzzle. Remarkably, Burnside et al. (2008) still insists that carry excess returns are "uncorrelated to traditional risk factors".

The evidence presented in our original paper, and in this reply, presents a serious challenge to the view that risk is not priced in currency markets. All the data used in Lustig and Verdelhan (2007) and in this reply are available on-line.[5] As a result, all tables in the 2007 paper and in this reply can be easily replicated.

Outline The rest of this paper is structured as follows. Section II addresses Burnside's first claim in detail. We have already addressed his second claim. To save space, we present additional results and robustness checks in the separate appendix. Section III studies the low frequency changes in carry trade risk premia and risk exposure. Section IV documents the high frequency time variation in risk prices

[4] Burnside, Eichenbaum, Kleshchelski, and Rebelo (2006) argue that the average payoffs to currency-speculation strategies are low and not risky, and that spot and forward prices move against currency traders in response to 'price pressure', thus explaining the forward premium puzzle. However, in a reversal, Burnside, Eichenbaum and Rebelo (2007) consider emerging market currencies and do obtain large excess returns, but they still argue against a risk-based explanation. The same claim appears in Burnside, Eichenbaum and Rebelo (2008) and Burnside, Eichenbaum, Kleshchelski, and Rebelo (2008). Burnside, Eichenbaum, and Rebelo (2009) argue that the forward premium puzzle is explained by adverse selection, but not by risk. In the first version of their paper, Burnside, Eichenbaum, Kleshchelski, and Rebelo (2008) conclude that excess returns in currency markets are not risk compensation arguing that "the peso problem cannot be a major determinant of the payoff to the carry trade." The second version reaches a different conclusion. Finally, Burnside and Bing Han and David Hirshleifer and Tracy Yue Wang (2010) argue that investors' overconfidence explains the forward premium puzzle.

[5] Data sets are available at `http://hlustig2001.squarespace.com/` and at `http://web.mit.edu/adrienv/www/`.

and quantities on currency markets, and section V concludes. The separate appendix, available on-line, details our sources and the construction of each variable, and provides additional evidence.

II The Exposure to Consumption Risk Across Currencies

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 betas are all indistinguishable from zero. This is incorrect. We start with the consumption beta estimates.

We first refute Burnside's claim with respect to our previous sample, which ends in 2002. We then check the robustness of our results by extending our annual sample through 2008.

A Consumption Growth Betas in Previous Sample

Let us first recall our previous results. Table 6 of our previous paper reports the univariate consumption betas and standard errors. The (nondurable and durable) consumption betas for the seventh currency portfolios are significantly different from zero, but this does not hold for the other portfolios. We obviously agree with Burnside's comment that consumption betas are not estimated as precisely as return-based betas.

Are the consumption betas all indistinguishable from zero and not different across portfolios, as Burnside claims? We show in Table II that 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. As already reported, this spread in betas over the entire sample easily accounts for the carry trade excess return. The same is true on the post–Bretton Woods sample. The consumption beta of the return on the U.S. stock market is 1.2 over the 1971–2002 sample. To explain the average annual stock market excess return of 5.75% over the same post–Bretton Woods sample in the standard consumption CAPM, the price of consumption risk has to be 4.9%. This implies a substantial spread of $7.4 = 1.5 \times 4.9\%$ on the HML strategy, compared with 6.9% in the data for the 1971–2002 sample.

All these betas are statistically significant at the 5% confidence level and are economically meaningful as well.

Table II: 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 reports OLS estimates of β_1 in the following time-series regression of the spread on the factor: $HML_{FX,t+1} = \beta_0 + \beta_1^{HML} f_t + \epsilon_{t+1}$, where $HML_{FX,t+1}$ is the return on the seventh portfolio minus the return on the first portfolio. The estimates are based on annual data; standard errors are reported in brackets. Following Donald W. K. Andrews (1991), we use Newey–West heteroscedasticity-consistent standard errors with an optimal number of lags to estimate the spectral density matrix. The p-values (reported in %) are for a t-test on the slope coefficient. The factor f_t is nondurable consumption growth (Δc) in the left panel and durable consumption growth (Δd) in the right panel. In Panel II, we report the multivariate regressions $HML_{FX,t+1} = \beta_0 + \beta_1^{HML} \mathbf{f}_t + \epsilon_{t+1}$, where $\mathbf{f}_t = [\Delta c_t, \Delta d_t]$. The χ^2 are for a Wald test that the slope coefficients are zero. The data are annual, and the samples cover 1953–2002 and 1971–2002.

B Durable Consumption Growth and Currency Returns in Longer Sample

We now extend our sample through 2008. In order to conserve space, we consider only one risk factor: durable consumption growth. This risk factor matters for asset prices if preferences are nonseparable. In addition, durable consumption growth has intuitive appeal as an asset pricing factor because it is highly cyclical. Our baseline measure of durable consumption growth, denoted Δd_1, corresponds to the change in the stock of consumer durables, as in Yogo (2006). As a robustness check we also use a second measure, Δd_2, which is the log change in the quantity index for consumer durable goods from the U.S Bureau of Economic Analysis (BEA) fixed asset tables. Durable consumption growth is strongly pro-cyclical, more so than nondurable consumption growth. In fact, Joao Gomes, Leonid Kogan and Yogo (2009) find that an investment strategy that is long on the durable-good producers portfolio and short on the service industry portfolio earns a risk premium exceeding 4% annually.

In Figure 3, we plot the average currency excess returns against the durable consumption growth betas, for the first measure Δd_1, estimated on the 1953-2008 sample; we also add standard-error bands. The univariate durable consumption growth betas are reported in Table VII in Section B of the separate appendix, where Newey-West standard errors are reported in brackets. The durable consumption beta of low interest rate currencies (portfolio 1) differs from the consumption beta of high interest rate currencies (portfolio 7). Clearly, portfolios with higher interest rates tend to have higher consumption betas. What does this mean? It means that investing in low interest rate currencies does not carry the same risk as investing in high interest rate currencies. Low interest rate currencies tend to appreciate in bad times while high interest rate currencies tend to depreciate. The betas on the last portfolios (7 and 8) are statistically different from zero. The betas of the intermediate portfolios are lower, but so are the excess returns on these portfolios. Recall that we are only using one risk factor here; hence Burnside's discussion about the rank of the beta matrix is meaningless.

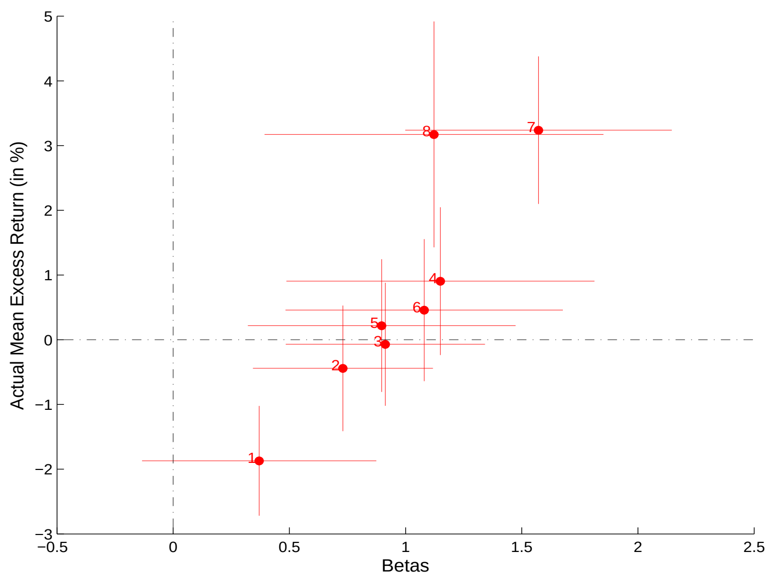

Figure 3: Durable Consumption Growth Betas and Average Currency Excess Returns

The dots represent point estimates; the lines represent one standard deviation above and below the point estimates. The sample is 1953–2008 and data are annual. Durable consumption growth (Δd_1) is defined in Section A in the separate appendix.

Durable consumption growth betas are higher for higher interest rate currencies. For the most comprehensive sample of currencies, which includes developing countries, we find that the durable consumption growth betas (i.e., the factor loadings on Δd_1) increase from 0.70 for the first portfolio to 1.37 for the last portfolio. We find similar results using the second measure; the loadings on Δd_2 increase from 0.51 to 1.44. The variation in betas increases in the post–Bretton Woods sample: the loadings on Δd_1 increase from 1.19 on the first portfolio to 2.19 on the sixth portfolio, declining to 1.64 on the last portfolio; the loadings on Δd_2 range from 0.83 to 1.60. For the sample of developed currencies, the spreads in loadings are even larger. The loading increases from 0.44 on the first portfolio to 1.46 on the last portfolio for our baseline measure, Δd_1, and from 0.46 to 1.11 for Δd_2. In the post–Bretton Woods sample, the loadings increase from 0.68 to 2.83 on the sixth portfolio, and then decreases to 2.27 on the last portfolio; from 0.67 to 1.83 to the fifth portfolio.

We now examine Burnside's (2007) claim that there is no statistically significant relation between durable consumption growth and the excess returns on the currency portfolios. For Δd_1, this null hypothesis is rejected at the 1% significance level for portfolio 7, at the 5% level for portfolio 3 and at the 10% level for portfolios 2 and 6. For Δd_2, this null is rejected at the 5% significance level for five portfolios out of 8 and at the 10% level for an additional portfolio. In the post–Bretton Woods sample, this claim is rejected at the 1% significance level for three portfolios, at the 5% level for two more, and at 10% for a last one. In sum, six of the eight portfolios have significant consumption betas, and the results are similar for the second measure of durable consumption growth (Δd_2). The results are even stronger if we consider only developed currencies. When using the first measure the null is rejected at the 5% level four times and once at the 10% level. In the post-Bretton-Woods sample, the null is rejected at the 1% level for three portfolios and at the 5% level for two more. Durable consumption growth betas are statistically different from zero.

Are the loadings on the other currency portfolios statistically different from those on the first portfolio? To answer this question, we look again at loadings on returns on zero-cost portfolios that short the first basket of currencies and go long in the other currencies. These results are reported in the bottom panel of Table VII in the separate appendix (Section B). Without exception, the loadings are positive. For the sample of developed currencies, the loadings are significantly different from zero at the 10% level in half the cases, when we use Δd_1, the first measure of durable consumption growth. In the post–Bretton Woods sample, all of the differences (excepting the fourth) are statistically different from zero. Clearly, the null that the betas on the first portfolio of funding currencies are identical to those of the other currency portfolios is overwhelmingly rejected at conventional levels of significance. These loadings for the $8-1$ portfolio vary between 0.68 (resp. 0.45) in the entire (post–Bretton Woods) sample for all currencies, and 1.02 (resp. 1.59) in the sample of developed currencies.

We have established that the average high interest rate currency *depreciates* relative to the average low interest rate currency in case of negative durable consumption growth innovations and *appreciates* in the case of positive durable consumption growth innovations. This effect is statistically significant, and also economically meaningful; for the sample of developed currencies, the beta of the zero-cost portfolio is around 2, which indicates a 2% depreciation for every 1% drop in durable consumption growth below its mean.

We have shown in this section that average excess returns correspond to covariances between returns and risk factors. The quantities and prices of risk are precisely estimated when one uses return-based risk factors. Consequently, it is not possible to ignore a risk-based approach to exchange rates merely on the basis of the large standard errors obtained with consumption-based risk factors.

We now rapidly turn to Burnside's second claim. Burnside stresses that the constant in the second stage of our regression is large and negative. We have already largely addressed this claim in the introduction of this paper. The constant is not significant. It simply represents the dollar factor. For the sake of completeness, we re-estimate our model on a set of 7 assets, built from the original 8 portfolios by going short in first portfolio (low interest currencies) and long in the other ones. We use Fama and MacBeth (1973) and GMM procedures. To save space, we report these additional results in Section C of the separate appendix. Figure 2 in this paper already shows the main point: estimating the model on dollar-neutral portfolios gives similar results with or without a constant.

In the asset pricing literature, standard errors on the estimates of macroeconomic factor loadings are typically large. We report in our previous paper and in this reply estimates that are nonetheless significant. To reduce the estimation uncertainty, an alternative strategy is to construct return-based factors. These are much more precisely measured and thus deliver better estimates of the loadings and the risk prices. In Lustig, Roussanov, and Verdelhan (2008), we pursue this approach using monthly currency return data constructed from one-month forward contracts (not T-bills) over the period from 1983.1 to 2008.12. Our results extends to a longer sample of quarterly returns starting either in 1953:I or in 1971:I. The second principal component is a slope factor that explains a large share of the cross-sectional variation in average returns on the currency portfolios. There is much less estimation uncertainty when we use return-based factors. The market price of risk has a t-statistic of 3.4. Again to save space, we report these additional results in Section D in the separate appendix.

III Low Frequency Variation in the Carry Trade Risk Premium and Risk Exposure

We have refuted Burnside's two claims regarding our previous sample. We turn now to the additional evidence made available by extending the sample in time and frequency. The sample used in Lustig and Verdelhan (2007) ends in 2002. Since we published our first paper, new data have, if anything, strengthened our case. The correlation between currency returns and risk factors has increased. For example, the correlation between nondurable consumption growth and returns on the Deutsche Bank carry trade index is equal to 0.4 in the 1993:II–2009:II sample, and it increases to 0.6 in the 2003:I to 2009:II subsample. Thanks to the long sample of high frequency currency returns of Lustig, Roussanov, and Verdelhan (2008), we can compare the increased correlation of returns and risk factors to the average excess returns on currency markets.

Is this increased correlation between consumption growth and carry returns evident in currency risk premia? It turns out that during the last ten years the carry trade risk premium has doubled. Table III provides an overview of the returns on the monthly currency portfolios of Lustig, Roussanov, and Verdelhan (2008), where the monthly returns are annualized. The table reports the moments for the entire sample in the first panel. The second panel shows the sample moments for the first half of the sample (1983–1995), the third panel for the second half of the sample (1996–2009). Finally, the fourth panel shows the results for the sample starting in 2000. The carry risk premium, the average return on a long position in the last portfolio and a short position in the first portfolio, has increased from 2.9% (2.16% for developed currencies) in the first half of the sample to 6.84% (3.65%) in the second half of the sample. The premium essentially doubles in the second half of the sample.

Table IV gives the estimated factor loadings for the different subsamples of monthly data. In the panel on the left, we report the contemporaneous betas for durable expenditure growth, nondurable expenditure growth, and the market return. In the panel on the right we report forward-looking betas, which are obtained by regressing returns at $t+1$ on consumption growth and returns between t and $t+3$.[6]

In the 1983.11–1995.12 sample, which is characterized by a small carry risk premium, a long position in the highest interest rate currencies and a short position in the first position did not expose investors to durable expenditures risk. The estimated loadings are close to zero and sometimes negative. Similarly, the market betas are small or negative. However, in the second half of the sample (1996.1–2009.6), the

[6] Jonathan Parker and Christian Julliard (2005) show that future consumption growth improves the explanatory power of the standard consumption CAPM, presumably because of lags in household consumption adjustment, which are especially relevant at monthly frequencies. These are the time-series regressions we run:

$$R^{j,e}_{t+1} = \theta^{j,k}_0 + \theta^{j,k}_1 f^k_{t\to t+3} + \eta^{j,k}_{t+1},$$

where $R^{j,e}_{t+1}$ denotes the excess return on portfolio j and $f^k_{t\to t+3}$ denotes the risk factor.

Table III: Monthly Currency Portfolio Returns

	All Currencies	Developed Currencies
Portfolios	6 minus 1	5 minus 1
	Panel A: 1983.11–2009.6	
Mean	4.97%	2.94%
Standard Deviation	9.03%	9.74%
Sharpe Ratio	0.55	0.30
	Panel B: 1983.11–1995.12	
Mean	2.90%	2.16%
Standard Deviation	9.14%	8.72%
Sharpe Ratio	0.42	0.25
	Panel C: 1996.1–2009.6	
Mean	6.84%	3.65%
Standard Deviation	8.92%	10.61%
Sharpe Ratio	0.77	0.34
	Panel D: 2000.1–2009.6	
Mean	6.52%	4.42%
Standard Deviation	7.84%	11.26%
Sharpe Ratio	0.83	0.39

Notes: The mean and standard deviation are annualized. Results reported for monthly currency portfolio returns from Lustig, Roussanov, and Verdelhan (2008), updated through June 2009.

durable expenditure beta of the long short position is 0.16, increasing to 0.19 in the subsample that starts in 2000. The market beta is 0.20. The results for developed currencies are similar but not significant. In the forward-looking consumption betas, the differences are even more striking. The durable expenditure beta increases from -0.46 in the first half of the sample to 0.45 in the second half, and the market beta increases from -0.01 to 0.25. These differences are statistically significant. In short, the large increase in the carry trade risk premium over the past 14 years has been matched by a similar increase in the consumption and market betas of carry trade returns.

Finally, the last panel reports betas using industry returns for non-durables and durables from Kenneth French's web site; he creates 10 industry portfolios using all stocks traded on NYSE-AMEX-NASDAQ. These betas confirm our earlier results. The durable beta increases from 0.02 to 12, the non-durable beta from 0.07 to 0.16. In the sample of developed currencies, the increases in consumption betas measured using industry returns are even larger. In the second half of the sample (1996.1–2009.6), where carry trade excess returns are the largest, industry betas are all significant.

IV High Frequency Variation in the Carry Trade Risk Premium and Risk Exposure

In the first sections of this paper, we have only used unconditional betas, obtained by regressing returns on the factors over the entire sample, to measure the quantity of risk inherent in each portfolio of currencies. However, as we have just seen, the factor betas of currencies vary over time. As demonstrated by Jagannathan and Wang (1996), this time variation matters for asset pricing if the betas co-vary with risk prices. We show that this does occur in currency markets. The factor betas of these currencies increase whenever the market price of risk increases. Hence, the average factor loadings reported by Lustig and Verdelhan (2007) tend to understate the riskiness of these currencies.

Consider two assets with zero average market betas: one asset has a beta of 1.5 during recessions and crises whereas the other asset has a beta of -1.5 during those times. Considering only average betas, one might conclude that investors should be indifferent between these two assets, but this conjecture is false unless the price of risk is constant (see Jagannathan and Wang 1996 for a detailed analysis). Yet it seems reasonable to assume that the price of risk *increases* during episodes like the financial crisis.[7]

In Lustig and Verdelhan (2007) we start from Epstein-Zin preferences and use the implied three components of the stochastic discount factor as risk factors – namely, consumption growth in nondurables and durables and stock market returns. This last factor proxies for the return on wealth. In unconditional asset pricing tests, this return factor does not play much of a role in explaining the cross-section of currency excess returns. Using the annual returns of Lustig and Verdelhan (2007), we show in Section E of the separate appendix that unconditional equity market betas are too low, and thus lead to implausibly high market prices of risk. However, the stock market risk in carry trades increases during crisis episodes. To make this point we turn again to higher-frequency data, the daily and monthly returns of Lustig, Roussanov, and Verdelhan (2008). We start with the current recession and then show that similar results obtain for previous crises.[8]

The recent subprime mortgage crisis offers a good example of the changing nature of the connection between currency and equity markets. Figure 4 plots the monthly returns on a carry trade at daily

[7] Time variation in the quantity and price of risk is indeed well established on equity and bond markets. It underlies the leading dynamic asset pricing models: the habit preferences of John Campbell and John H. Cochrane (1999), the long-run risk model of Ravi Bansal and Yaron (2004), and the time-varying disaster risk of Xavier Gabaix (2009). We show that a similar result holds for currency markets: the factor loadings of carry trades tend to increase dramatically during these episodes. This implies that the average betas tend to understate the true riskiness of these investments.

[8] We could make a similar point about the subprime crisis and the CAPM using the Deutsche Bank currency index as in Section **??**. However, because this index does not begin until 1993, we would not be able to present results for previous crises.

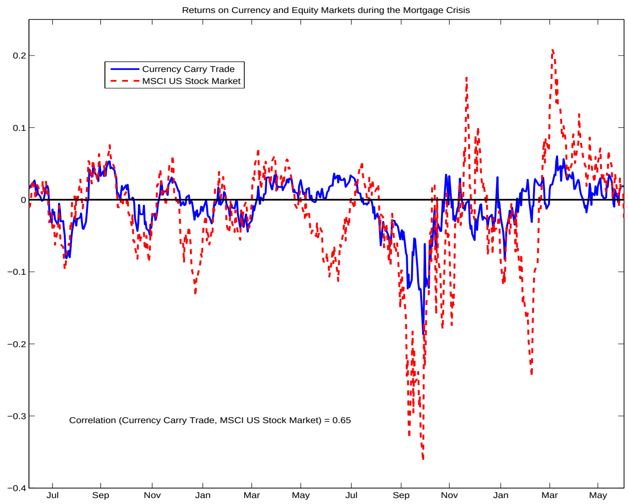

Figure 4: Currency Carry Trade and U.S. Stock Market Returns during the Mortgage Crisis (July 2007 through June 2009).

This figure plots one-month carry trade returns at daily frequency against one-month returns on the U.S. MSCI stock market index at daily frequency. Currency carry trade returns come from Lustig, Roussanov, and Verdelhan (2008). They correspond to the returns on their last portfolio (i.e., high interest rate currencies) minus the return on their first portfolio (i.e., low interest rate currencies). The sample period is 2 July 2007 through 30 June 2009.

frequencies against the U.S. stock market return. Clearly, a U.S. investor who was long in these high interest rate currencies and short in low interest rate currencies was heavily exposed to U.S. aggregate stock market risk during the subprime mortgage crisis, and therefore should have been compensated by a risk premium ex ante.

This increase in correlations is not specific to the recent mortgage crisis. We compute the correlation between one-month currency returns and the return on the value-weighted U.S. stock market return using 12-month rolling windows on daily data over the entire 1983-2009 sample. Figure 5 plots the difference between the correlation of the 6th and the 1st portfolio with the U.S. stock market excess return. We also plot the stock market beta of HML_{FX}, defined again as the difference in returns between high and low interest rate portfolios. The shaded areas indicate NBER recessions and financial crises. These market correlations exhibit enormous variation. In times of crisis and during U.S. recessions, the difference in market correlation between high and low currencies increases significantly. During the Mexican, Asian, Russian, and Argentinean crises, the correlation difference jumps up by 50–90 basis points.

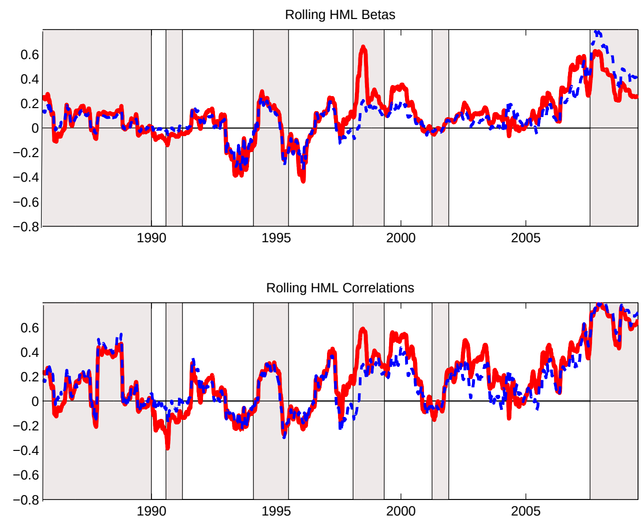

Figure 5: HML Market Correlations and Betas

This figure first presents $Corr_\tau[R_t^m, HML_{FX,t}]$, where $Corr_\tau$ is the sample correlation over the previous 12 months (253 days; i.e, over the sample $[\tau - 253, \tau]$), R^m is the return on the U.S. stock market, and HML_{FX} denotes currency carry trade returns. We use monthly returns at daily frequency. The figure also presents β_{HML}, the stock market beta of HML_{FX}. The stock market return is the return on U.S. MSCI index. The solid red line uses carry returns on a large sample of developed and emerging countries; the dotted blue line corresponds to a sample of developed countries. The sample period is 31 October 1983 through 30 June 2009.

Market Betas We explore time variation in the stock market betas over the 1983–2008 sample in more detail for three crisis episodes. To estimate the market β^m, we use daily observations on monthly currency and stock market returns. During these three crisis episodes, stock market betas of carry trades increase dramatically.[9] Table V reports the market betas of currency portfolios for 6-month windows before the end of May 1998, August 2007 and June 2009.

We observe that β^m_{HML} increases to 1.08 in the run-up to the Russian default in 1998, implying that high interest rate currencies depreciate on average by 1.08% relative to low interest rate currencies when the stock market goes down by 1%. In times of crisis, low interest rate currencies provide a hedge against market risk while high interest rate currencies expose U.S. investors to more market risk. And as expected the estimated market betas increase monotonically as we move from low to high interest rate currency portfolios. Similarly, β^m_{HML} increases to 0.64 at the start of the mortgage crisis in August of 2007 and drops to 0.13 toward the end of the crisis in June 2009.

[9] Charlotte Christiansen, Angelo Ranaldo, and Paul Soderlind (2009) find similar results on a shorter sample (1995–2008) using a logistic smooth transition regression model.

Consumption Betas Also, there is substantial high frequency variation in the consumption betas as well. These consumption betas tend to increase during recessions and financial crises. In Table VI, we revisit the same three episodes and report the consumption exposure in these long–short currency strategies. We use 18-month windows to estimate the factor betas. The nondurable betas of the $6-1$ (HML_{FX}) zero-cost portfolio are 0.99 in May 1998, 1.53 in August 2007, and 1.73 in June 2009 (as estimated using the most comprehensive sample of currencies). For the sample of developed currencies, the respective consumption betas are 1.98, 3.76 and 2.07. The panel on the right reports the corresponding industry return betas. These produce the same pattern.

V 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 by interest rates. These portfolios average out the idiosyncratic risk in exchange rate changes, and this produces a sharper picture of the risk-return trade-off in currency markets. In our sample, low interest rate currency portfolios have low consumption growth betas and high interest rate currency portfolios have high consumption growth betas. This implies that the forward premium puzzle has a risk-based explanation. Burnside argues that the data are not informative about the relation between consumption growth and foreign currency returns. We disagree and 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 precisely. Our portfolios do not allow us to identify the price of dollar risk which is Burnside's second point. We agree, but this is not what either our paper or the forward premium puzzle is about. Our paper is concerned with 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.

What Have We Learned? Our approach of building portfolios of currencies has helped to establish that low and high interest rate currencies have undeniably different risk characteristics. The current crisis provides a painful lesson to anyone who doubts this. Had researchers started by looking at these currency portfolios 25 years ago, the forward premium would probably not have been assigned the "puzzle" label. In fact, these researchers would have been puzzled to find that uncovered interest rate parity actually holds.

Our approach to studying currencies has been adopted by several authors recently (including Burnside). It enabled, for example, Roberto A. DeSantis and Fabio Fornari (2008), Jakub W. Jurek (2008), Lukas Menkho, Lucio Sarno, Maik Schmeling and Andreas Schrimpf (2009), Farhi, Samuel Fraiberger, Gabaix, Romain Ranciere and Verdelhan (2009), Christiansen, Ranaldo and Soderlind (2009), Andrew Ang and Joseph S. Chen (2010) to make further progress on the road to a better understanding of exchange rates.

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Table IV: Estimation of Factor Loadings on Monthly Data

	Betas		Forward-Looking Betas		Industry Return Betas	
	All Currencies	Developed Currencies	All Currencies	Developed Currencies	All Currencies	Developed Currencies
Portfolios	6 minus 1	5 minus 1	6 minus 1	5 minus 1	6 minus 1	5 minus 1
			Panel A: 1983.11–2009.6			
Durables	0.04	0.04	-0.07	-0.04	0.08***	0.10***
	[0.05]	[0.06]	[0.23]	[0.20]	[0.02]	[0.03]
Nondurables	0.25	0.41	1.03*	1.12	0.11**	0.14**
	[0.26]	[0.32]	[0.63]	[0.88]	[0.04]	[0.05]
Market	0.15***	0.17***	0.14**	0.19	0.15***	0.17***
	[0.03]	[0.05]	[0.07]	[0.10]	[0.03]	[0.05]
			Panel B: 1983.11–1995.12			
Durables	-0.02	0.01	-0.46	-0.28	0.02	0.03
	[0.05]	[0.05]	[0.28]	[0.19]	[0.04]	[0.04]
Nondurables	0.21	0.15	0.89	0.93	0.07	0.09
	[0.37]	[0.33]	[0.85]	[0.65]	[0.04]	[0.05]
Market	0.07	0.10	-0.01	0.05	0.07	0.10*
	[0.05]	[0.05]	[0.08]	[0.08]	[0.05]	[0.05]
			Panel C: 1996.1–2009.6			
Durables	0.16*	0.09	0.45**	0.28	0.12***	0.14**
	[0.09]	[0.13]	[0.22]	[0.32]	[0.03]	[0.05]
Nondurables	0.28	0.63	1.11	1.25	0.16**	0.22**
	[0.36]	[0.52]	[0.94]	[1.45]	[0.06]	[0.09]
Market	0.20***	0.22***	0.25***	0.28**	0.20***	0.22***
	[0.04]	[0.07]	[0.08]	[0.12]	[0.04]	[0.07]
			Panel D: 2000.1–2009.6			
Durables	0.19**	0.17	0.44*	0.54	0.11***	0.16***
	[0.09]	[0.13]	[0.25]	[0.36]	[0.03]	[0.05]
Nondurables	0.35	0.72	1.33	2.32	0.08	0.32**
	[0.40]	[0.61]	[1.05]	[1.59]	[0.09]	[0.12]
Market	0.17***	0.29***	0.25***	0.44***	0.17***	0.29***
	[0.05]	[0.08]	[0.09]	[0.13]	[0.05]	[0.08]

Notes: The left panel reports contemporaneous betas; the middle panel reports forward-looking betas; the right panel reports non-durable and durable goods industry betas (10 industry portfolios downloadable from French's web site). Results reported for monthly currency portfolio returns from Lustig, Roussanov, and Verdelhan (2008), updated through June 2009. Standard errors are reported in brackets. Following Andrews (1991), we use Newey–West heteroscedasticity-consistent standard errors with an optimal number of lags to estimate the spectral density matrix. We use one asterisk to denote significance at the 10% level, two for 5%, and three for 1%. Section A in the separate appendix contains a detailed description of the monthly consumption data.

Table V: Estimation of Market Loadings on Daily Data — Three Case Studies

Portfolio	α_m^i	β_m^i	R^2	α_m^i	β_m^i	R^2	α_m^i	β_m^i	R^2
Sample	*Panel A: 26 May 1998*			*Panel B: 31 Aug 2007*			*Panel C: 30 Jun 2009*		
1	-0.83	-0.06	0.63	0.18	-0.14^{***}	13.52	0.37	0.31^{***}	60.69
	[0.73]	[0.15]		[0.37]	[0.05]		[0.45]	[0.03]	
2	-0.52	-0.04	0.39	0.15	0.22^{*}	28.11	0.47	0.22^{***}	54.97
	[1.03]	[0.17]		[0.36]	[0.06]		[0.33]	[0.03]	
3	-1.29^{***}	0.19^{*}	11.35	0.72^{***}	0.18^{***}	29.72	0.14	0.33^{***}	49.16
	[0.42]	[0.10]		[0.26]	[0.05]		[0.78]	[0.06]	
4	-1.67^{***}	0.33^{**}	7.13	0.29	0.21^{***}	39.76	0.62	0.29^{***}	52.48
	[0.59]	[0.14]		[0.26]	[0.04]		[0.54]	[0.05]	
5	-3.13^{**}	0.58^{**}	15.44	0.46^{**}	0.15^{***}	43.24	0.87	0.43^{***}	59.35
	[1.50]	[0.27]		[0.15]	[0.04]		[0.76]	[0.07]	
6	-3.68^{***}	1.03^{***}	40.47	0.60	0.51^{***}	54.47	0.74	0.45^{***}	71.19
	[1.08]	[0.23]		[0.43]	[0.11]		[0.77]	[0.09]	
6 minus 1	-2.85^{***}	1.08^{***}	53.36	0.42	0.64^{***}	67.39	0.37	0.13^{**}	20.04
	[0.63]	[0.16]		[0.37]	[0.08]		[0.67]	[0.08]	

Notes: This table reports estimates of the CAPM betas during crises. The sample period is 129 days (6 months) before and including the mentioned date in each panel. The table reports the intercept α_m^i, slope coefficient β_m^i, and R^2 in a regression of each portfolio i's currency excess returns on a constant and the U.S. stock market return. The intercept α_m^i and the R^2 are reported in percentage points. The Newey–West standard error correction is computed with 20 lags. We use one asterisk to denote significance at the 10% level, two for 5%, and three for 1%. We use the daily currency portfolios from Lustig, Roussanov, and Verdelhan (2008), updated through June 2009 and the MSCI return on the U.S. stock market.

Table VI: Estimation of Factor Loadings on Monthly Data — Three Case Studies

	Betas		Industry Return Betas	
	6 minus 1	5 minus 1	6 minus 1	5 minus 1
	All Currencies	Developed Currencies	All Currencies	Developed Currencies
	Panel A: May 1998			
Durables	0.38	0.21	0.15	-0.10
	[0.51]	[0.27]	[0.24]	[0.11]
Nondurables	0.99	1.98**	0.43**	0.03
	[2.14]	[0.95]	[0.17]	[0.11]
Market	0.34*	-0.07	0.34*	-0.07
	[0.19]	[0.13]	[0.19]	[0.13]
	Panel B: August 2007			
Durables	-0.01	-1.15	0.29*	0.30*
	[0.49]	[0.77]	[0.18]	[0.22]
Nondurables	1.53*	3.76***	0.55**	0.47*
	[0.95]	[1.42]	[0.14]	[0.26]
Market	0.46***	0.27	0.46***	0.27*
	[0.16]	[0.18]	[0.16]	[0.18]
	Panel C: June 2009			
Durables	0.91*	1.79**	0.11**	0.22***
	[0.56]	[0.81]	[0.04]	[0.08]
Nondurables	1.73**	2.07**	0.29***	0.65***
	[0.68]	[1.05]	[0.12]	[0.19]
Market	0.26***	0.54***	0.26***	0.54***
	[0.10]	[0.12]	[0.10]	[0.12]

Notes: Table entries are regressions on the 18 months preceding the event, including the month itself. Following Andrews (1991), we use Newey–West heteroscedasticity-consistent standard errors with an optimal number of lags to estimate the spectral density matrix. We use one asterisk to denote significance at the 10% level, two for 5%, and three for 1%. The left panel reports actual betas for consumption and the market returns. The right panel reports non-durable and durable goods industry betas (10 industry portfolios downloadable from French's web site).

The Cross-Section of Foreign Currency Risk Premia and Consumption Growth Risk: A Reply

- Supplementary Online Appendix -

NOT FOR PUBLICATION

This separate appendix complements the paper "The Cross-Section of Foreign Currency Risk Premia and Consumption Growth Risk: A Reply". We report additional results on:

- Durable consumption growth betas of returns and exchange rates: see Table VII and VIII
- Estimation on dollar-neutral portfolios: see Tables IX and X (with and without a constant), and GMM estimates in Table XI
- GLS estimation of linear factor models: see Table XII and XIII
- FMB estimation of linear factor models and preference parameters: see Table XIV
- Estimation of linear factor models without a constant but with an additional factor: see Table XV
- Principal components of currency portfolio returns: see Table XVI
- Estimation of the CAPM and conditional CAPM: see Table XVII

This separate appendix first details our data in Section A. We focus on the quantity of risk and report in Section B additional regression results on durable consumption growth betas. This additional evidence addresses Burnside's first claim. We then turn to the market prices of risk in Section C. We address Burnside's second claim and presents robustness checks on the model's estimation. We show how to extract risk factors from exchange rates at quarterly frequency in Section D. Finally, we report evidence on time-varying equity prices of risk in Section E.

A Appendix: Data

Deutsche Bank Carry Trade Return Index The Deutsche Bank G10 Currency Harvest Index consists of long futures contracts on the three G10 currencies associated with the highest interest rates and short futures contracts on the three G10 Currencies associated with the lowest interest rates. This index re-evaluates interest rates quarterly and, based on the evaluation, reweights the futures contracts it holds. Immediately after each reweighting, the index will reflect an investment on a 2 to 1 leveraged basis in the three long futures contracts and in the three short futures contracts. The index is available online at *http://www.dbfunds.db.com/Dbv/index.aspx* and starts in 1993. The PowerShares DB G10 Currency Harvest Fund (symbol: DBV), which replicates the Deutsche Bank index, has been listed on the NYSE since 18 September 2006. As a result, the recent returns and losses that we report were accessible to many investors.

LRV Carry Trade Returns Lustig, Roussanov, and Verdelhan (2008) use monthly data on one-month forward and spot exchange rates to construct currency portfolio returns. We have updated these data through June 2009. Lustig, Roussanov, and Verdelhan (2008) construct portfolios of currencies sorted by their forward discounts. We construct six portfolios on the entire set of currencies, five for the developed currencies. The data are available online in the Excel worksheet `overview-data-monthly`.

Monthly Consumption Data The U.S. National Income and Product Accounts (NIPA) has recently begun constructing monthly consumption series. We use the growth rate in real total consumption expenditures. Because we do not have monthly data on the stock of durable consumption goods, we cannot construct durable consumption growth. However, we check our results on the growth rate of durable consumption expenditures. The monthly consumption data is from Table 2.8.1. (Percent Change from Preceding Period in Real Personal Consumption Expenditures by Major Type of Product, Monthly). Line 1 is Personal Consumption Expenditures; we call this total consumption growth. Line 2 is Durable goods; we call this durable expenditures. Line 3 is nondurables; we call this nondurable consumption. We do not divide these series by the number of households. The monthly stock return is the real returns on the CRSP-VW index.

Crises Dates for the Tequila crisis and the Long Term Capital Management crisis were taken from Kho, Lee, and Stulz (2000). Dates for the less-developed-country crisis are from the FDIC web site, available on-line.

Annual Consumption Data The annual personal consumption expenditures in dollars are obtained from Table 2.3.5. (Personal Consumption Expenditures by Major Type of Product), and the price series are obtained from Table 2.3.4. (Price Indexes for Personal Consumption Expenditures by Major Type of Product). The nondurable consumption series is constructed as the sum of nondurable goods (line 6) deflated by its price index (line 6 in Table 2.3.4), services (line 13) deflated by its price (line 13), housing services (line 14) deflated by its price (line 14), and clothes and shoes (line 8) deflated by its price (line 8).

What enters the average investor's utility function is the service flow provided by the stock of durables, not the expenditures on durable consumption goods. The stock is our measure of the service flow provided by the durables. Following Yogo (2006), we measure durable consumption growth as the change in the stock of consumer durables D. Instead of constructing our own measure (using the perpetual inventory method) to extend the sample, we simply used the fixed asset tables, which are available only at annual frequency. For the 1953–2002 sample, our measure is very close to Yogo's. We use two different measures of durable consumption growth.

Our baseline measure of durable consumption growth, denoted Δd_1, is the log change in the deflated current cost stock of consumer durables (line 13 from Fixed Assets Table 1.1: Current-Cost Net Stock of Fixed Assets and Consumer Durable Goods) divided by the number of U.S. households. We also use a second measure, Δd_2; it is the log change in the quantity index for consumer durable goods (line 13 from Fixed Assets Table 1.2: Chain-Type Quantity Indexes for Net Stock of Fixed Assets and Consumer Durable Goods). We divide these series by the number of households.

Return Factors SMB and HML are obtained from Kenneth French's web site. These are listed as Fama-French factors. SMB goes long in a basket of small stocks and short in a basket of long stocks. The momentum factor is also downloadable from Kenneth-French's web site; it goes long in a portfolio of winners and short in a portfolio of losers.

B Appendix: Durable Consumption Growth Betas of Returns and Exchange Rates

In this section, we focus on the quantity of risk. Table VII reports durable consumption growth betas of returns on currency portfolios. Durable consumption growth betas are higher for higher interest rate currencies. For the most comprehensive sample of currencies, which includes developing countries, we find that the durable consumption growth betas (i.e., the factor loadings on Δd_1) increase from 0.70 for the first portfolio to 1.37 for the last portfolio. We find similar results using the second measure; the loadings on Δd_2 increase from 0.51 to 1.44. The variation in betas increases in the post–Bretton Woods sample: the loadings on Δd_1 increase from 1.19 on the first portfolio to 2.19 on the sixth portfolio, declining to 1.64 on the last portfolio; the loadings on Δd_2 range from 0.83 to 1.60. For the sample of developed currencies, the spreads in loadings are even larger. The loading increases from 0.44 on the first portfolio to 1.46 on the last portfolio for our baseline measure, Δd_1, and from 0.46 to 1.11 for Δd_2. In the post–Bretton Woods sample, the loadings increase from 0.68 to 2.83 on the sixth portfolio, and then decreases to 2.27 on the last portfolio; from 0.67 to 1.83 to the fifth portfolio.

What is the source of this covariance between currency returns and consumption growth: interest rates or exchange rates? We can disentangle the contribution of changes in exchange rates and interest rates to the factor loadings of returns that we have reported. To do so, we run a simple time-series regression of the annual exchange rate changes in portfolio j (averaged over all the currencies in this portfolio) on durable consumption growth. The results of this time-series regression are reported in Table VIII. For developed currencies, all of the consumption exposure is a consequence of the exchange rate exposure; for less developed currencies, some of the exposure is due to interest rates. High interest rate currencies depreciate when U.S. durable consumption growth is low and appreciate when consumption growth is high.

Table VIII reports durable consumption growth betas of exchange rates. We start with developed currencies. The top panel reports the results for long positions in each portfolio. The bottom panel reports the results for short positions in the first portfolio and long positions in the other portfolios. For Δd^1, the loadings of exchange rate changes increase from .52 on the first portfolio to 2.76 on the sixth portfolio; the loadings on the last portfolio is 2.35. These loadings are statistically different from zero at the 5 % level, except for the loadings on portfolio 1 and portfolio 4. For Δd^2, the loadings of exchange rate changes increase from .64 on the first portfolio to 1.91 on the fifth portfolio; the loadings on the last portfolio is 1.75. These loadings are statistically significantly different from zero, except for the ones on portfolios 1 and 4. We also tested whether these loadings differ significantly across portfolios by looking at the loadings of zero-cost portfolios of currencies that go long in the high interest rate portfolio and short in the low interest rate portfolio. For the first factor Δd^1, the loadings are significantly different from zero at the 10 % level for all portfolios except for portfolios 4 and 7. The loadings are larger in the post-Bretton-Woods sample, as one would expect. When we include the less developed currencies, the loadings on the higher interest rate portfolios are less precisely estimated. Moreover, they sometimes are smaller for the higher interest rate portfolios. For these currencies, the interest rates themselves contribute significantly to the loadings of returns on durable consumption growth (i.e., higher interest rate differences with the US when US durable consumption growth is high).

C Appendix: Prices of Currency Risk

In this section, we focus on the prices of aggregate risk. We first focus on Burnside's second claim and then report several robustness checks on the model's estimation that were left out from our main paper.

.1 What Is the Price of Dollar Risk?

We start with Burnside's second claim. Burnside stresses that the constant in the second stage of our regression is large and negative. He then argues (a) that a risk-based explanation can be discounted because our model overpredicts the returns on all eight currency portfolios and (b) that our R^2 overstates the fit of the model because it includes this constant.

The constant measures the risk price of variations in the dollar (relative to all other currencies) that cannot be attributed to consumption growth. In Section D below, we show that the first principal component of currency returns is a dollar factor. All currency portfolios have essentially the same loadings on this factor.[10] As a result, our cross-section of currency portfolios is not informative about the price of dollar risk. In order to estimate the price of dollar risk, currencies need to be sorted by their exposure to the dollar risk factor, not by interest rates.

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

To confirm that the constant actually measures the price of dollar risk, we 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. 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

[10] It is important to note that this is typically *not* true for stock portfolios and the loadings on the market return.

Euler equations should also be satisfied for these zero-cost strategies, but the returns are not affected by the fluctuations in the dollar. If our interpretation of the constant is correct, we should observe a smaller intercept λ_0.

GMM Estimation In Table XI we report the GMM estimates obtained on these seven test assets. The factors are de-meaned. The consumption risk prices are 2.8 and 4.8, respectively, and these values 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 then 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. The resulting risk price estimates are much like the ones we obtained on the same test assets without this additional factor but including a constant. These results are reported in Table XV.

As a result, the *EZ-DCAPM* model overpredicts 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 is what the forward premium puzzle and our previous paper are about.

FMB Estimation As a robustness check, Table IX reports the results for the Fama–MacBeth (1973) estimation of the linear factor models using these test assets. In the benchmark *EZ-DCAPM* (column 5), the constant λ_0 drops 350 basis points (from 290 to -60 basis points), and is not significantly different from zero. The R^2 is 81%. This measure is based on the regression *with* a constant, yet *without* a constant is (the case considered in the next paragraph) the R^2 "drops" to 79 %. 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, and we find similar results for the second subsample. 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.

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 the literature. So let us turn again to those 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 use Burnside's preferred measure of fit. Table X reports the results. The price of nondurable and durable consumption risk are significantly different from zero, and the model accounts for 79% of the variation in these returns. Figure 2 compares the models estimated with and without the constant plotting 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.

Our paper is not about dollar risk. We agree with Burnside that consumption risk does not explain the average returns earned by U.S. investors on a basket of all foreign currencies, and we have never claimed that it did. Our focus is on the returns obtained by going long in high interest rate currencies and short in low interest rate currencies for this is how the carry trade is defined.

.2 Robustness checks

Having checked the quantities of risk and the role of the constant, we turn now to the prices of carry risk. We now compare the evidence in Lustig and Verdelhan (2007) on risk price estimates against Burnside's claim. In our previous paper we report bootstrapped standard errors, Shanken-corrected standard errors, and GMM standard errors. In this appendix we add GLS standard errors.

Notation For the reader's convenience, we briefly review the notation and methodology of our previous paper. We keep the same notation in this reply. Starting from the Euler equation, one can derive a linear factor model whose factors are nondurable U.S. consumption growth Δc_t, durable U.S. consumption growth Δd_t, and the log of the U.S. market return r_t^m. The U.S. 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 nests as special cases the *CCAPM* with Δc_t as the only factor, the *DCAPM* with factors Δc_t and Δd_t, the *EZ-CCAPM*, with factors Δc_t and r_t^m, and the *CAPM*. 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 multiplied by the amount of risk 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 classic estimation proceeds in two stages. In the first stage, we estimate the betas (β^j) by running a time-series regression of returns on the factors. In the second stage, we estimate the market prices of risk for all the factors (λ) by running a cross-sectional regression of average returns on the betas.

Bootstrap In panel B of Table 14, we report the standard errors in braces $\{\cdot\}$ obtained by bootstrapping the whole estimation. These standard errors take into account the uncertainty in the first stage of the estimation as well as the small sample size. They were generated by running the estimation procedure on 10,000 samples constructed by drawing with replacement both from the observed returns and factors 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 for 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, then the bootstrapping exercise would have revealed this. Instead, it confirms that our results are significant.

Shanken correction Table 14 also reports the Shanken-corrected standard errors in parenthesis $(\cdot)$. The Shanken correction (see Jay Shanken 1992), which is only valid asymptotically, produces substantially larger standard errors than the ones we generated by bootstrapping. Ravi Jagannathan and Jiang Wang (1998) show that the uncorrected Fama and MacBeth (1973) 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 our previous paper that conditional heteroskedasticity is the key to understanding these currency betas.

Generalized Method of Moments In addition, panel A of Table 14 reports the two-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 nondurable consumption risk is significant (3.2 with a standard error of 0.9); likewise, the price of durable consumption risk is positive and significant (3.4 with a standard error of 1.2). Burnside also discards the GMM evidence because he insists on estimating the mean of the factors; adding three 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 an appealing outcome. Yogo (2006) encounters a similar problem and adjusts the weighting matrix to deal with it, as he explains in an appendix (p. 575). Because of these issues, our approach of *not* estimating the mean of the factors is actually more standard.[11]

Generalized Least Squares In Table XII of this paper, we report the GLS estimates that were left out of the previously published version. We remark that GLS estimators are more efficient than OLS estimators because they put more weight on the more informative moment conditions.[12] Clearly, for the *D-CAPM* and the *EZ-DCAPM*, the market price of durable consumption risk is significant at the 5% level, even when we use the asymptotic Shanken correction upon which Burnside insists. The price of nondurable consumption risk is about 3.2, with a Shanken-corrected standard error of 1.8 and bootstrapped errors of about 1.2. The price of durable consumption risk is approximately 5.15, with a Shanken-corrected standard error of about 2.3 and bootstrapped errors of about 1.7. The measures of fit are lower because GLS does not simply minimize the squared pricing errors; it minimizes the weighted sum. Table XIII reports similar results for the post–Bretton Woods subsample. Burnside's claim that the risk prices are not statistically different from zero does not hold for that sample either.

Table XIII reports results from a the GLS estimation on the post-Bretton Woods sample. Table XIV reports the FMB estimation of linear factor models and preference parameters. Table XV presents the estimation of linear factor models without a constant but with an additional, which is the average return on currency markets.

[11] For example, Martin Lettau and Sydney Ludvigson (2001) report results from a GMM estimation of their linear factor model, yet do not to estimate the mean of the factors.

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

D Appendix: Extracting Risk Factors from Currency Returns

The current recession provides additional support for a risk-based explanation of exchange rates. The number of recessions, however, remains limited. Not surprisingly, the standard errors on the estimates of macroeconomic factor loadings are typically large. This is a well-known fact in the asset pricing literature. To reduce the estimation uncertainty, an alternative strategy is to construct return-based factors. These are much more precisely measured and thus deliver better estimates of the loadings and the risk prices. The Arbitrage Pricing Theory developed by Stephen A. Ross (1976) provides a theoretical underpinning for this methodology.

In Lustig, Roussanov, and Verdelhan (2008), we pursue this approach using monthly currency return data constructed from one-month forward contracts (not T-bills) over the period from 1983.1 to 2008.12. We find that these currency returns exhibit a clear factor structure. The first two principal components of the returns account for most of the time-series variation in the returns on the currency portfolios. All excess returns load approximately equally on the first principal component (PC_1), which is close to the mean of all currency returns. We call this the *dollar* factor. However, loadings on the second principal component (PC_2) increase monotonically with the interest rates. This is why we refer to it as the *slope* factor. Exchange rates of various currencies covary in the right way when we sort these currencies by their interest rates: high interest rate currencies move together, and so do low interest rate currencies. This is a necessary condition for a risk-based explanation.

We show in this paper that our results extends to a longer sample of quarterly returns on the portfolios of foreign T-bills used in Lustig and Verdelhan (2007). The annual series are not as informative about the covariance matrix and the factor structure of exchange rates. Starting either in 1953:I or in 1971:I, a clear factor structure emerges. The second principal component is a slope factor that explains a large share of the cross-sectional variation in average returns on the currency portfolios.[13] There is much less estimation uncertainty when we use return-based factors. The market price of risk has a t-statistic of 3.4. Table XVI reports the estimates we obtained on the currency portfolios. The left-hand side of the table corresponds to our long sample (1953:I–2008:IV); the right-hand side corresponds to the post–Bretton Woods sample (1971:I–2008:IV). In both cases, the risk factors are the first two principal

[13]The first principal component loads on all portfolios in the same way. On the 1953:IV–2009:II sample, its loadings are [0.12, 0.12, 0.13, 0.12, 0.12, 0.11, 0.11, 0.11]. The second principal component loads very differently on low and high interest rate portfolios. On the same sample, its loadings are [−0.65, −0.45, −0.10, 0.13, 0.25, 0.22, 0.44, 0.20]. We obtain similar results on the post–Bretton Woods sample.

components of the currency excess returns.

Cross-Sectional Regressions The top panel of Table XVI reports estimates of the market prices of risk λ, the adjusted R^2, the square root of mean-squared errors $RMSE$, and the p-values of χ^2 tests (in percentage points).

The dollar risk factor, PC_1, has an estimated risk price of 148 basis points. Its value is not precisely estimated, which is not surprising because all the portfolios have similar betas with respect to this dollar factor. As a result, this factor explains none of the cross-sectional variation in portfolio returns. Although the dollar factor does not explain any of the cross-sectional variation in expected returns, it is important for the level of average returns; it mimics a constant in the cross-sectional regression because all of the portfolios have the same exposure to this factor.

This is the same issue mentioned in Section C. The intercept measures the price of that component of dollar risk not explained by our risk factors (consumption growth and the market return). Why do we call it "dollar risk"? The first principal component of all currency excess returns has the same loadings on all currency portfolios. Hence, in the cross-sectional regression of average returns on factor loadings, the dollar risk loadings behave as a constant. The risk price of this factor is hard to estimate precisely, exactly because it affects all of the currencies in the same way. The cross-section of currency returns is obviously not informative here. Suppose we eliminated the dollar risk factor from the regressions. Then all of the alphas in panel II of Table XVI would increase by 148 basis points, but we would still be explaining 65% of the cross-sectional variation.

The slope factor, PC_2, has a risk price of 293 basis points per annum. This means that an asset with a beta of 1 earns a risk premium of 2.93% per annum.[14] The FMB standard error is 83 basis points. The risk price is more than 3 standard errors removed from zero and thus is highly statistically significant. The lambdas indicate whether risk is priced, and the second principal component clearly captures aggregate risk on currency markets. Overall, the pricing errors are small. The $RMSE$ is around 63 basis points, and the adjusted R^2 is 63%. The null that the pricing errors are zero cannot be rejected, regardless of the estimation procedure. All these results hold also in a smaller sample starting in 1971, as shown on the right-hand side of Table XVI.

[14]Because the factors are returns, the absence of arbitrage implies that the risk prices of these factors should equal their average excess returns. This condition stems from the fact that the Euler equation applies to the risk factor itself, which clearly has a regression coefficient β of 1 on itself. In our estimation, this no-arbitrage condition is automatically satisfied because the two risk factors are orthogonal; in this case, the risk prices exactly equal the factor means.

Time-Series Regressions The bottom panel of Table XVI reports the intercepts and the slope coefficients obtained by running time-series regressions of each portfolio's currency excess returns on a constant and risk factors. The returns and alphas are in percentage points per annum. The first column reports alpha estimates, which are not statistically significant at the 5% significance level. The null that the alphas are jointly zero cannot be rejected at either the 5 or either the 10% level.

The third column of the same panel reports the estimated betas for the second principal component. These betas increase monotonically from -0.65 for the first portfolio to 0.44 for the seventh currency portfolio, decreasing slightly to 0.20 for the last portfolio. These betas are estimated very precisely. The first two portfolios have betas that are negative and significantly different from 0; the last five have betas that are positive and significantly different from zero. The second column shows that betas for the dollar factor are essentially all equal to 1. Obviously, this dollar factor does not explain any of the variation in average excess returns across portfolios, but it helps to explain the average level of excess returns. These results are comparable to the ones obtained using a shorter sample (reported on the right-hand side of the table).

E Appendix: Time Variation in Equity Risk Prices

On a long sample at annual frequency, we report in Lustig and Verdelhan (2007) that the CAPM accounts for less than 5% of the cross-sectional variation in returns. Using monthly returns and a shorter sample leads to more favorable results. But the CAPM is still not a good description of the currency returns. The top panel of Table XVII reports CAPM results with 6 portfolios. The US stock market excess return can account for a large share of the cross-sectional variation in returns. However, the estimated price of US market risk is close to 29 percent, while the actual annualized excess return on the market is only 5.5 percent over this sample. The risk price is more than 5 times too large. The CAPM β's vary from -.03 for the first portfolio to .12 for the last one. Low interest rate currencies provide a hedge, while high interest rate currencies expose US investors to more stock market risk. These β's increase almost monotonically from low to high interest rates, but they are too small to explain these excess returns. Therefore, the cross-sectional regression of currency returns on market β's implies market price of risk that are far too high.

The bottom panel of Table XVII reports results obtained on 12 test assets (the original 6 currency portfolios and the same ones multiplied by the lagged VIX index). Risk factors are the Fama-French

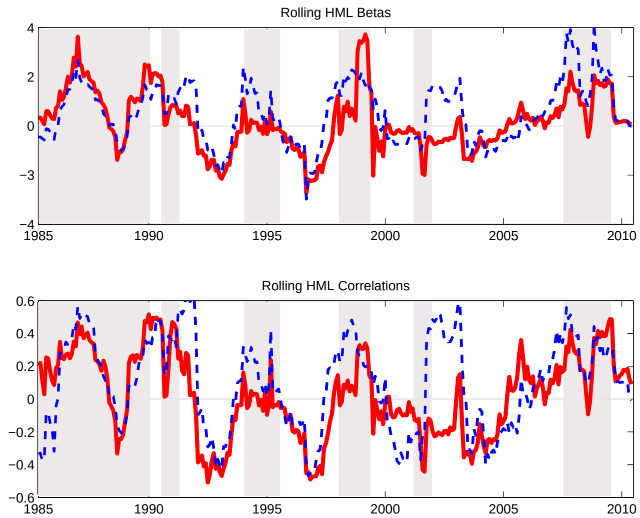

Figure 6: HML Consumption Correlations and Betas

We plot the 18-month nondurable consumption growth betas of HML (upper panel) and 18-month correlations of nondurable consumption growth and HML (lower panel). The red solid line is for all currencies; the blue dotted line is for developed currencies. Monthly currency portfolios from Lustig, Roussanov, and Verdelhan (2008) are updated through May 2010. The shaded areas are NBER recessions and the LDC crisis, the Tequila crisis, and the LTCM crisis.

value-weighted stock market excess return R^M and $R^M z$, which is R^M multiplied by the lagged value of the VIX index (scaled by its standard deviation). We find that the market price of risk increases significantly in bad times (when the stock market volatility index VIX is high). However, taking into account such time-variation is not enough to justify the CAPM: market prices of risk are still too high.

Table VII: Durable Consumption Growth Betas of Returns

	All Currencies								Developed Currencies							
Portfolios	1	2	3	4	5	6	7	8	1	2	3	4	5	6	7	8
								Panel A: 1953–2008								
Durables (1)	0.37	0.73*	0.91**	1.15*	0.90	1.08*	1.57***	1.12	0.23	1.90**	1.56*	0.62	1.04	1.69**	1.90**	1.85**
	[0.50]	[0.39]	[0.43]	[0.66]	[0.58]	[0.60]	[0.57]	[0.73]	[0.80]	[0.86]	[0.93]	[0.59]	[0.85]	[0.77]	[0.78]	[0.81]
Durables (2)	0.33	0.68**	0.69**	0.90*	0.67	0.91**	1.10**	1.24**	0.31	1.23*	1.02	0.68	0.99	1.11*	1.59**	1.41**
	[0.40]	[0.33]	[0.37]	[0.51]	[0.47]	[0.49]	[0.47]	[0.58]	[0.59]	[0.69]	[0.76]	[0.54]	[0.65]	[0.59]	[0.68]	[0.62]
								Long–Short in 1								
Durables (1)		0.36	0.54	0.78	0.53*	0.71*	1.20***	0.75		1.68***	1.33*	0.39	0.82	1.47*	1.67**	1.62***
		[0.37]	[0.36]	[0.48]	[0.27]	[0.42]	[0.41]	[0.67]		[0.65]	[0.81]	[0.58]	[0.64]	[0.82]	[0.73]	[0.55]
Durables (2)		0.35	0.36	0.56	0.33	0.58	0.77**	0.90*		0.92	0.71	0.37	0.68	0.80	1.28**	1.10**
		[0.29]	[0.31]	[0.35]	[0.27]	[0.36]	[0.37]	[0.52]		[0.56]	[0.65]	[0.48]	[0.51]	[0.66]	[0.61]	[0.48]
								Panel B: 1971–2008								
Durables (1)	0.67	1.11*	1.56***	2.21***	1.66**	1.82**	2.38***	1.22	0.37	2.81**	2.50**	0.99	1.72	2.77***	2.98***	2.85***
	[0.71]	[0.57]	[0.48]	[0.73]	[0.78]	[0.83]	[0.80]	[1.08]	[1.20]	[1.12]	[1.23]	[0.85]	[1.18]	[0.98]	[1.01]	[1.08]
Durables (2)	0.57	0.96**	1.10**	1.52**	1.13*	1.45**	1.53**	1.29*	0.47	1.68*	1.51	0.94	1.52*	1.73**	2.33***	2.06***
	[0.52]	[0.45]	[0.48]	[0.60]	[0.60]	[0.59]	[0.64]	[0.77]	[0.82]	[0.90]	[1.00]	[0.73]	[0.83]	[0.71]	[0.80]	[0.76]
								Long–Short in 1								
Durables (1)		0.44	0.89*	1.54***	0.99**	1.15**	1.71***	0.55		2.44***	2.13**	0.62	1.35	2.40***	2.61***	2.49***
		[0.50]	[0.50]	[0.56]	[0.41]	[0.58]	[0.61]	[0.98]		[0.78]	[1.03]	[0.87]	[0.84]	[0.97]	[0.82]	[0.63]
Durables (2)		0.39	0.53	0.95**	0.56	0.89**	0.96*	0.73		1.21	1.04	0.47	1.06	1.26	1.86***	1.60**
		[0.38]	[0.41]	[0.43]	[0.39]	[0.45]	[0.53]	[0.65]		[0.78]	[0.91]	[0.68]	[0.67]	[0.84]	[0.71]	[0.63]

Notes: Results obtained on annual currency portfolio returns. The returns are from Lustig and Verdelhan (2007), updated through 2008; standard errors are reported in brackets. Following Andrews (1991), we use Newey–West heteroscedasticity-consistent standard errors with an optimal number of lags to estimate the spectral density matrix. We use one asterisk to denote significance at the 10% level, two for 5%, and three for 1%. The first measure, denoted *Durables (1)*, is the log change in the deflated current cost stock of consumer durables; the second measure, *Durables (2)*, is the log change in the quantity index for consumer durable goods. The Data Appendix contains a detailed description of our variables.

Table VIII: Estimation of Durable Consumption Growth Betas of Exchange Rates

	All Currencies								Developed Currencies							
Portfolios	1	2	3	4	5	6	7	8	1	2	3	4	5	6	7	8
Panel A: 1953-2008																
Long																
Durables (1)	-0.03	0.43	0.73^{**}	1.00^{*}	0.36	0.64	0.77	-0.60	0.03	1.74^{**}	1.32^{*}	0.42	0.79	1.57^{**}	1.74^{**}	1.62^{**}
	[0.38]	[0.39]	[0.35]	[0.52]	[0.47]	[0.49]	[0.60]	[0.90]	[0.77]	[0.73]	[0.78]	[0.49]	[0.71]	[0.66]	[0.70]	[0.63]
Durables (2)	0.09	0.50	0.56^{**}	0.71^{*}	0.22	0.58	0.39	-0.51	0.21	1.13^{**}	0.89	0.55	0.84	1.05^{**}	1.52^{***}	1.26^{**}
	[0.33]	[0.35]	[0.28]	[0.43]	[0.37]	[0.43]	[0.53]	[0.84]	[0.58]	[0.57]	[0.62]	[0.45]	[0.55]	[0.50]	[0.58]	[0.51]
Long-Short in 1																
Durables (1)		0.46	0.76^{**}	1.02^{**}	0.39	0.67	0.80	-0.57		1.72^{**}	1.30	0.39	0.76	1.54^{*}	1.71^{**}	1.59^{**}
		[0.38]	[0.33]	[0.47]	[0.31]	[0.45]	[0.61]	[0.89]		[0.67]	[0.84]	[0.61]	[0.70]	[0.90]	[0.86]	[0.68]
Durables (2)		0.41	0.47	0.62^{*}	0.14	0.49	0.30	-0.60		0.92	0.68	0.34	0.63	0.84	1.30^{*}	1.05^{*}
		[0.30]	[0.30]	[0.37]	[0.28]	[0.39]	[0.55]	[0.77]		[0.61]	[0.67]	[0.50]	[0.54]	[0.72]	[0.72]	[0.58]
Panel B: 1971-2008																
Long																
Durables (1)	0.07	0.64	1.22^{**}	1.76^{***}	0.85	1.23^{*}	1.34	-0.05	0.06	2.50^{***}	2.08^{**}	0.65	1.31	2.57^{***}	2.71^{***}	2.55^{***}
	[0.58]	[0.59]	[0.50]	[0.68]	[0.70]	[0.69]	[0.82]	[0.92]	[1.15]	[0.97]	[1.01]	[0.85]	[0.99]	[0.83]	[0.91]	[0.90]
Durables (2)	0.22	0.67	0.85^{**}	1.19^{**}	0.51	1.06^{**}	0.68	0.05	0.32	1.48^{**}	1.27	0.72	1.30^{**}	1.62^{***}	2.20^{***}	1.90^{***}
	[0.46]	[0.50]	[0.36]	[0.53]	[0.51]	[0.51]	[0.70]	[0.93]	[0.80]	[0.73]	[0.81]	[0.67]	[0.70]	[0.60]	[0.71]	[0.62]
Long-Short in 1																
Durables (1)		0.57	1.14^{**}	1.69^{***}	0.78^{*}	1.16^{*}	1.26	-0.12		2.44^{***}	2.02^{*}	0.59	1.25	2.50^{**}	2.65^{***}	2.49^{***}
		[0.53]	[0.48]	[0.62]	[0.46]	[0.60]	[0.87]	[0.96]		[0.84]	[1.09]	[0.91]	[0.89]	[1.04]	[0.99]	[0.75]
Durables (2)		0.45	0.63	0.97^{**}	0.29	0.84^{*}	0.46	-0.18		1.16	0.95	0.40	0.98	1.30	1.88^{**}	1.58^{**}
		[0.38]	[0.43]	[0.49]	[0.41]	[0.51]	[0.77]	[0.89]		[0.83]	[0.95]	[0.72]	[0.70]	[0.92]	[0.84]	[0.71]

Results obtained on annual currency portfolio returns. The returns are from Lustig and Verdelhan (2007), updated to 2008. The standard errors are reported in brackets. We use Newey-West heteroscedasticity-consistent standard errors with an optimal number of lags to estimate the spectral density matrix following Andrews (1991). We use one asterisk to denote significance at the 10 % level, two for 5%, three for 1%. The first measure, denoted Δd^1, is the log change in the deflated current cost stock of consumer durables. The second measure, Δd^2, is the log change in the quantity index for Consumer durable goods. The Data Appendix contains a detailed description.

Table IX: 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.

Table X: FMB Estimation: Dollar-Neutral Currency Portfolios, 1953-2002, 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 FMB estimates of the risk prices (in percentage points) using 7 annually re-balanced currency portfolios as test assets. No constant included. 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 XI: GMM Estimation: Dollar-Neutral Currency Portfolios, 1953–2002

	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 two-stage GMM estimates of the factor prices (in percentage points) using seven annually rebalanced currency portfolios as test assets. These test assets go long in the nth currency portfolio and short in the first portfolio. The sample is 1953–2002. The data are annual, from Lustig and Verdelhan (2007). 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). Standard errors are reported in brackets. The factors are de-meaned. The pricing errors correspond to the first-stage estimates. The last two rows report the mean absolute pricing error (in percentage points) and the p-value for a χ^2 test.

Table XII: GLS Estimation of Linear Factor Models, 1953–2002

	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}
Nondurables	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 generalized least squares (GLS) estimates of the risk prices (in percentage points) using eight annually rebalanced currency portfolios as test assets. The sample is 1953–2002. The data are annual, from Lustig and Verdelhan (2007). The factors are demeaned. The OLS standard errors are reported in brackets; the Shanken-corrected standard errors in parentheses, and the boostrapped errors are in braces. The last three rows report the mean absolute pricing error (in percentage points), the $\bar{R}^2$, and the p-value for a χ^2 test.

Table XIII: GLS Estimation of Linear Factor Models: 1971-2002

	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 XIV: Estimation of Linear Factor Models and Preference Parameters: 1953-2002

	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 XV: Estimation of Linear Factor Models: No Constant but Additional Factor

	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 β's. 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 XVI: Asset Pricing — Principal Components

	1953:IV–2009:II					1971:IV–2009:II				
	Panel I: Risk Prices									
	PC_1	PC_2	R^2	$RMSE$	χ^2	PC_1	PC_2	R^2	$RMSE$	χ^2
FMB	1.48	2.93	66.54	0.63		1.80	2.82	47.09	0.87	
	[0.98]	[0.86]			9.08	[1.44]	[1.26]			13.58
	(0.98)	(0.86)			11.32	(1.44)	(1.26)			15.54
Mean	1.48	2.93				1.80	2.82			
	Panel II: Factor Betas									
Portfolio	*Intercept*	PC_1	PC_2	R^2		*Intercept*	PC_1	PC_2	R^2	
1	−0.52	1.04	−0.65	87.92		−0.79	1.05	−0.65	88.40	
	[0.47]	[0.04]	[0.05]			[0.68]	[0.04]	[0.05]		
2	0.80	1.06	−0.45	85.69		0.94	1.06	−0.45	86.01	
	[0.55]	[0.05]	[0.06]			[0.78]	[0.05]	[0.06]		
3	−0.24	1.20	−0.10	86.73		0.04	1.20	−0.08	87.43	
	[0.49]	[0.05]	[0.04]			[0.69]	[0.05]	[0.05]		
4	−0.26	1.07	0.13	80.84		−0.24	1.08	0.14	82.18	
	[0.57]	[0.05]	[0.04]			[0.78]	[0.05]	[0.04]		
5	−1.03	1.07	0.25	81.40		−1.67	1.07	0.27	83.08	
	[0.63]	[0.05]	[0.07]			[0.83]	[0.05]	[0.07]		
6	0.45	1.03	0.22	70.74		0.51	1.01	0.14	75.08	
	[0.71]	[0.06]	[0.07]			[0.98]	[0.06]	[0.07]		
7	−0.02	1.03	0.44	84.72		0.15	1.03	0.46	85.97	
	[0.49]	[0.04]	[0.05]			[0.68]	[0.04]	[0.04]		
8	0.94	0.98	0.20	83.93		1.17	0.98	0.20	84.75	
	[0.42]	[0.04]	[0.04]			[0.60]	[0.04]	[0.04]		
All	7.37	49.74%				6.65	57.47%			

Notes: Panel I reports results from GMM and Fama–MacBeth asset pricing procedures. Market prices of risk λ, the adjusted R^2, the square root of mean-squared errors $RMSE$, and the p-values of χ^2 tests on pricing errors are reported in percentage points. Excess returns used as test assets and risk factors take into account bid–ask spreads. All excess returns are multiplied by 4 (annualized). Shanken-corrected standard errors are reported in parentheses. We do not include a constant in the second step of the FMB procedure. Panel II reports OLS estimates of the factor betas. The R^2s and p-values are reported in percentage points. The standard errors in brackets are Newey and West (1987) standard errors computed with the optimal number of lags according to Andrews (1991). The χ^2 test statistic $\alpha' V_\alpha^{-1} \alpha$ tests the null that all intercepts are jointly zero. This statistic is constructed from the Newey–West variance-covariance matrix (one lag) for the system of equations (see Cochrane 2001, p. 234). Data are quarterly, from Global Financial Data. The sample includes only developed countries. Portfolios are rebalanced every quarter. The sample period is 1953:IV–2009:II for the left panel and 1971:IV-2009:II for the right panel. The alphas are annualized and reported in percentage points.

Table XVII: Estimation of Linear Factor Model Risk Prices: CAPM

	$Market$	$Market - VIX$	R^2	$RMSE$	χ^2
	Panel A: Unconditional CAPM				
FMB	27.00		70.50	1.08	
	[14.60]				9.80
	(16.21)				18.77
Mean	**5.67**				
	Panel B: CAPM with VIX as conditioning variable				
FMB	29.91	75.46	69.64	2.07	
	[11.75]	[29.51]			25.57
	(13.80)	(34.22)			55.13
Mean	**5.04**	**10.53**			

Notes: This table reports results from Fama-McBeth asset pricing test. Market prices of risk λ, the adjusted R^2, the square-root of mean-squared errors $RMSE$ and the p-values of χ^2 tests are reported in percentage points. In the top panel, the risk factor is the Fama-French value-weighted stock market excess return R^m. In the bottom panel, the risk factors are the value-weighted stock market excess return R^m and $R^m VIX$, which is R^m multiplied by the lagged value of the VIX index (scaled by its standard deviation). The portfolios are constructed by sorting currencies into six groups at time t based on the interest rate differential at the end of period $t-1$. Portfolio 1 contains currencies with the lowest interest rates. Portfolio 6 contains currencies with the highest interest rates. In the bottom panels, we use 12 test assets: the original 6 portfolios and 6 additional portfolios obtained by multiplying the original set by the conditioning variable (VIX). Data are monthly, from Barclays and Reuters (Datastream). The sample is 11/1983–6/2009 for the top panel and 02/1990–6/2009 for the bottom panel. Standard errors are reported in brackets. Shanken-corrected standard errors are reported in parentheses. We do not include a constant in the second step of the FMB procedure.