Global Macro Risks in Currency Excess Returns

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Abstract

We study the cross-sectional variation of carry-trade-generated currency excess returns in terms of their exposure to global macroeconomic fundamental risk. The risk factor is the cross-country high-minus-low (HML) conditional skewness of the unemployment gap. It is robustly priced in currency excess returns and provides a measure of global macroeconomic uncertainty. A widening of the HML gap signifies increasing divergence, disparity, and inequality of economic performance across countries.

Keywords: Currency excess returns, beta-risk, carry trade, global macro risk, uncertainty
JEL: F3,F4, G1

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Introduction

In this paper, we study the cross-sectional variation of carry-trade-generated portfolios of currency excess returns as a function of their exposure to systematic risk. The proposed risk factors are high-minus-low (HML) differences between the top and bottom quartiles of conditional moments of country-level macroeconomic performance indicators. Movements in these easily computable risk factors reflect variations in global economic uncertainty. The HML concept is heavily used in finance. By sorting into quartiles, our HML variable is similar to the interquartile range, which is a robust measure of distributional dispersion. This measure captures an important aspect of global uncertainty. Our emphasis on conditional second and third moments, draws attention to a second dimension of economic uncertainty.

Our main result is that the HML skewness of the unemployment gap is a macroeconomic fundamental risk factor that is robustly priced into the carry-trade-generated currency excess returns. In a globally integrated financial market, it makes sense that investors pay attention to the state of the global economy. Our results are not dependent on emerging-market economies, nor are they driven by the global financial crisis.

The factor is constructed by computing the conditional skewness of each country's unemployment gap and subtracting the average of the bottom quartile of countries from the average of the top quartile. Countries in the top quartile have a high probability of above-normal unemployment, and a high chance of entering the bad state. Countries in the bottom quartile have a high probability of below-normal unemployment, and a high chance of entering the good state. Movements in the factor reflect variations in divergence, disparity, and inequality of fortunes across national economies. The factor is robust to alternative conditional moments (mean and volatility) and alternative macro fundamentals (changes in the unemployment rate, output gap, output growth, real exchange rate gap, real exchange rate depreciation, consumption growth rate, and inflation rate). The significance of the conditional skewness measure underscores the importance of asymmetries in the state of nature which are obscured by volatility measures of uncertainty.

A legacy literature sought to understand currency excess returns by trying to resolve the forward premium anomaly–recognized as an empirical regularity since Hansen and Hodrick (1980), Bilson (1981), and Fama (1984).[1] Although the forward premium anomaly implies non-zero currency excess returns, they are two different and distinct phenomena (Hassan and Mano (2014)). Recent research in international finance has de-emphasized the forward premium anomaly, focused directly on currency excess returns, and has produced new insights. A methodological innovation introduced by Lustig and Verdelhan (2007), was to change the observational unit from individual returns to portfolios of returns.

[1] Regressions of the future currency depreciation on the interest differential typically give a negative slope coefficient in violation of the zero-profit uncovered interest rate parity (UIP) condition. Hodrick (1987), Engel (1996), and Lewis (1995) survey earlier work on the topic, which viewed excess returns as risk premia and emphasized the time-series properties of individual currency excess returns. Whether through estimation or quantitative evaluation of asset pricing models, explanatory power was low and this body of work was unable to produce or identify mechanisms for risk premia that were sufficiently large or acceptably correlated with the excess returns. This is not to say interest in the topic has waned. See, for example, Alvarez et al. (2009), Bansal and Shaliastovich (2012), Chinn and Zhang (2015), Engel (2016), and Verdelhan (2010).

Identification of systematic risk in currency excess returns has long posed a challenge to this research, and the use of portfolios aids in this identification by averaging out idiosyncratic return fluctuations. Since the returns are available to global investors, and portfolio formation allows diversification of country-specific risk, presumably only global risk factors remain to drive portfolio returns.

Following the recent literature, our test assets are interest-rate ranked portfolios of currency excess returns. While the HML unemployment gap skewness factor looks like a risk factor to the portfolio returns, the mechanism differs across portfolios. The betas for low interest (and hence low currency excess return) portfolios are negative. This is due primarily to the exchange rate component. Currencies in these portfolios lose substantial value when there is an increase in global uncertainty, as measured by the factor. In contrast, beta on the high interest portfolio is positive, primarily on account of the interest rate component of returns. When the factor spikes up, the yields in this portfolio increase as global investors flee the debt of these countries. The currencies of the high interest portfolio countries also fall in the bad state but not enough to offset the increase in the interest differential.

To provide context and interpretation for the empirical results, we draw on an affine yield model (adapted from Lustig et al. (2011) and Backus et al. (2001)) of the term structure of interest rates, applied to pricing currency excess returns. In the model, countries' log stochastic discount factors (SDFs) exhibit heterogeneity in the way they load on a country-specific factor and a common global risk factor (the HML skewness of the unemployment gap). We estimate the model parameters by simulated method of moments and show that the model can qualitatively replicate key features of the data.

Our paper is part of a literature that studies portfolios of currency excess returns in the context of asset pricing models and is closest to the absolute asset pricing strand of the literature, which examines currency returns in terms of their exposure to macroeconomic fundamental risk (Lustig and Verdelhan (2007), Burnside, Eichenbaum, and Kleshchelski (2011), Jorda and Taylor (2012), Hassan (2013), Ready, Roussanov, and Ward (2015), Menkhoff, Lukas, Sarno, Schmeling, and Shrimpf (2013), and Della Corte, Riddiough, and Sarno (2013)). The relative asset pricing strand (Lustig, Roussanov, and Verdelhan (2011), Daniel, Hodrick, and Lu (2014), and Ang and Chen (2010)) studies risk factors built from other asset returns. Clarida, Davis, and Pederson (2009) and Christiansen, Ranaldo, and Söderlind (2011) focus on regime switches. Our paper is also aligned with a strand of the literature that connects notions of uncertainty to currency excess returns. Menkhoff, Sarno, Schmeling, and Shrimpf (2012) price returns to global foreign exchange volatility, Della Corte, Sarno, Schmeling, and Wagner (2015) price currency returns to sovereign risk, Brunnermeier, Nagel, and Pederson (2008), Jurek (2014), and Lettau, Maggiori, and Weber (2014) study the relation of returns to crash risk.

Although our paper is mainly empirical, from a macroeconomic modeling perspective, an improved understanding of currency excess returns can help inform future developments in modeling uncovered interest rate parity shocks. Frequently, macro models impose exogenous dynamics into deviations from uncovered interest rate parity (UIP) for the models to generate realistic exchange rate dynamics (Kollmann (2002), Devereux and Engel (2003), Engel (2015), and Itskhoki and Mukhin (2016)). Empirical analyses, such as ours, may aid in developing general equilibrium models with endogenous deviations from UIP.

The remainder of the paper is organized as follows. The next section discusses the construction of portfolios of currency excess returns. Section 2 describes the data. Section 3 implements the main empirical work. Section 4 presents the affine asset pricing model, and Section 5 concludes.

1 Portfolios of Currency Excess Returns

Identification of systematic risk in currency returns has long posed a challenge in international finance. In early research on single-factor models (e.g., Frankel and Engel (1984), Cumby (1988), and Mark (1988)), the observational unit was the excess U.S. dollar return against a single currency. Lustig and Verdelhan (2007) innovated on the methodology by working with portfolios of currency excess returns instead of returns for individual currencies. This is a useful way to organize the data because it averages out noisy idiosyncratic and non-systematic variation and improves the ability to uncover systematic risk. Since global investors have access to these returns, they can form such portfolios and diversify away country-specific risk. In a world of integrated financial markets, only undiversifiable global risk factors should be priced.

Before forming portfolios, we start with the bilateral carry trade. Let there be $n_t + 1$ currencies available at time t. Let the nominal interest rate of country i be $r_{i,t}$ for $i = 1, ..., n_t$, and the U.S. nominal interest rate be $r_{0,t}$. The United States will always be country '0.' In the carry, we short the U.S. dollar (USD) and go long in currency i if $r_{i,t} > r_{0,t}$. The expected bilateral excess return is

$$E_t \left((1 + r_{i,t}) \frac{S_{i,t+1}}{S_{i,t}} - (1 + r_{0,t}) \right) \simeq E_t \left(\Delta \ln \left(S_{i,t+1} \right) \right) + r_{i,t} - r_{0,t}, \tag{1}$$

where $S_{i,t}$ is the USD price of currency i (an increase in $S_{i,t}$ means the USD depreciates relative to currency i). If $r_{0,t} > r_{i,t}$, we short currency i and go long in the USD.[2]

Next, we extend the carry trade to a multilateral setting. We rank countries by interest rates from low to high in each time period and use this ranking to form portfolios of currency excess returns. As in Lustig et al. (2011), we form six such portfolios, called $P_1, \ldots, P_6$. The portfolios are rebalanced every period. Portfolios are arranged from low (P_1) to high (P_6), where P_6 is the equally weighted average return from those countries in the highest quantile of interest rates and P_1 is the equally weighted average return from the lowest quantile of interest rates. Excess portfolio returns are stated relative to the U.S.,

$$\frac{1}{n_{j,t}} \sum_{i \in P_j} (1 + r_{i,t}) \frac{S_{i,t+1}}{S_{i,t}} - (1 + r_{0,t}), \tag{2}$$

for $j = 1, \ldots, 6$. In this approach, the exchange rate components of the excess returns are relative to the USD. The USD is the funding currency if the average of P_j interest rates are higher than the U.S. rate and vice-versa. An alternative, but equivalent approach would be to short any of the $n_t + 1$ currencies and to go long in the remaining n_t currencies. Excess returns would be constructed by 'differencing' the portfolio return, as in Lustig et al. (2011) and Menkhoff et al. (2013), by subtracting the P_1 return

[2]The right hand side of (1) is the log-approximated excess return.

from P_2 through P_6.[3] It does not matter, however, whether excess returns are formed by the 'difference' method or by subtracting the U.S. interest rate. As Burnside (2011a) points out, portfolios formed by one method are linear combinations of portfolios formed by the other. The next section describes the data we use to construct the portfolios of currency excess returns as well as some properties of the excess return data.

2 The Data

The raw data are quarterly and have a maximal span from 1973Q1 to 2014Q2. When available, observations are end-of-quarter and point-sampled. Cross-country data availability varies by quarter. At the beginning of the sample, observations are available for 10 countries. The sample expands to include additional countries as their data become available, and contracts when data vanish (as when countries join the euro). Our encompassing sample is for 41 countries plus the euro area. The countries are Australia, Austria, Belgium, Brazil, Canada, Chile, Colombia, the Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Iceland, India, Indonesia, Ireland, Israel, Italy, Japan, Malaysia, Mexico, the Netherlands, New Zealand, Norway, the Philippines, Poland, Portugal, Romania, Singapore, South Africa, South Korea, Spain, Sweden, Switzerland, Taiwan, Thailand, Turkey, the United Kingdom, and the United States. The data set consists of exchange rates, interest rates, consumption, gross domestic product (GDP), unemployment rates, and the consumer price index (CPI). The macro data are *not* seasonally adjusted. Census seasonal adjustment procedures impound future information into today's seasonally adjusted observations, which is generally unwelcome. We remove the seasonality ourselves with a moving average of the current and three previous quarters of the variable in question.

Currency returns are formed using interbank interest rates and spot exchange rates. The exchange rate, $S_{j,t}$, is expressed as USD per foreign currency units so that a higher exchange rate represents an *appreciation* of the *foreign* currency relative to the USD. The data source from 1996Q1 to 2014Q2 is *Datastream* for three-month yields and *Bloomberg* for exchange rates. Before 1996, coverage from both sources was very thin. To extend the sample back to 1973Q1, exchange rates and interest rates for Australia, Belgium, Canada, France, Germany, Italy, Japan, the Netherlands, Switzerland, the United Kingdom, and the United States are from the Harris Bank *Weekly Review.* These are quotations from the last Friday of the quarter from 1973Q1 to 1995Q4.

One consideration in selecting countries in our sample was based on the availability of rates on interbank or Eurocurrency loans, which are assets for which traders can take short positions. Because the rates for alternative currencies are often quoted by the same bank, Eurocurrency/interbank rates net out cross-country differences in default risk. Imputing interest rates from the foreign exchange forward premium is not a good idea since covered interest parity has been reported to fail since the onset of

[3] If there are $n_{j,t}$ currencies (excluding the reference currency) in portfolio P_j, the USD ex post $P_6 - P_1$ excess return is

$$\frac{1}{n_{6,t}} \sum_{i \in P_6} (1 + r_{i,t}) \frac{S_{i,t+1}}{S_{i,t}} - \frac{1}{n_{1,t}} \sum_{k \in P_1} (1 + r_{k,t}) \frac{S_{k,t+1}}{S_{k,t}}. \tag{3}$$

the global financial crisis (Pinnington and Shamloo (2016) and Du, Tepper and Verdelhan (2016)).[4] Additional details on the interest rate data are provided in Appendix A.

Real consumption and GDP are from Haver Analytics. The unemployment rate and the consumer price index ($P_{j,t}$) are from the FRED database at the Federal Reserve Bank of St. Louis. The log real exchange rate between the United States (country '0') and country j is $q_{j,t} \equiv \ln\left(\left(S_{j,t}P_{j,t}\right)/P_{0,t}\right)$.

In many cases, because of the relatively short time span of the data, the real exchange rate and unemployment rate appear to be non-stationary. To induce stationarity in these variables, we work with their 'gap' versions. The gap variables are cyclical components from a recursively applied Hodrick-Prescott (1997) (HP) filter. The HP filter is applied recursively so as not to introduce future information into current observations. The GDP gap is constructed similarly.

In the next subsection, we construct portfolios of currency excess returns using the raw data described above and outline some key properties of this data.

2.1 Some properties of the data

Following Lustig et al. (2011), we sort countries by the interest rate in each time period into six equally weighted portfolios. The U.S. interest rate is subtracted from each portfolio return to form excess returns that are stated in percent per annum.

Table 1, Panel A, shows the log-approximate portfolio mean returns, mean excess returns, and their Sharpe ratios. As we describe below, construction of the factors requires 20 start-up observations, so the sample ranges from 1978Q1–2014Q2. Both the mean excess returns and the mean returns increase monotonically across the portfolios. There is not much variation in average excess returns and average returns between P_4 and P_5. There is a sizable jump in the average return and excess return from P_5 to P_6. These six portfolios form the cross-section of returns that we analyze below.

Figure 1 plots the cumulated portfolio excess returns from shorting the dollar and going long in the foreign currency portfolios. The carry trade performs poorly before the mid 1980s, but its profitability takes off around 1985. The observations available in the 1970s are mostly for European countries, who held a loose peg against the deutschemark, initially through the 'Snake in the Tunnel,' and then in 1979 through the European Monetary System. During this period, there is not much cross-sectional variation across countries, especially in their exchange rate movements against the USD. The U.S. nominal interest rate was also relatively high during this time period.

For additional context, Figure 2 plots the cumulated P_6 excess return together with the cumulated excess return on the Standard and Poor's 500 index over the same time span. The P_6 excess return is seen to be first-order large and important.

Table 1, Panel B, decomposes log-approximate portfolio excess returns into contributions from the interest rate differential and the exchange rate components. Interestingly, on average, there is no forward premium anomaly in the portfolio excess returns. To read the table, the average U.S. interest rate is 2.9% higher than the average P_1 interest rate. Uncovered interest rate parity (UIP) predicts an average

[4] We also found imputed interest rates to be excessively volatile and were often negative (in periods before central banks began paying negative interest).

dollar depreciation of 2.9%. The actual average dollar depreciation of 1.8% for P_1 currencies goes in the direction of UIP. Similarly, the average P_6 interest rate is 16.4% higher than the average U.S. interest rate. UIP predicts an average dollar appreciation of 16.4% and the dollar actually appreciates 9.5% against P_6 currencies, on average. If the forward premium anomaly were present, the dollar would have depreciated. Figure 3 plots the relationship between the portfolio interest rate differential and the dollar depreciation. The relationships between the average interest rate differential and the average depreciation is reminiscent of Chinn and Merideth's (2004) findings of long-horizon UIP.

What about the short-run relationship between interest rates and exchange rate returns? Table 2 reports estimates of the Fama (1984) regression for the six portfolios, which is the regression of the one-period-ahead dollar depreciation of the P_j portfolio ($j = 1, ..., 6$) on the U.S. − P_j interest differential. Let $\Delta s_{t+1}^{P_j} \equiv \frac{1}{n_{j,t}} \sum_{i \in P_j} \ln\left(\frac{S_{j,t+1}}{S_{j,t}}\right)$ be the dollar depreciation against portfolio j and $r_t^{P_j} \equiv \frac{1}{n_{j,t}} \sum_{i \in P_j} r_{j,t}$ be the yield on portfolio j's. The regression is

$$\Delta s_{t+1}^{P_j} = \alpha_j + \beta_{F,j}\left(r_{0,t} - r_t^{P_j}\right) + \epsilon_{j,t+1}. \tag{4}$$

According to the point estimates, there is a forward premium anomaly only for P_1, P_2, and P_3. These are portfolios whose interest rates are relatively close to the U.S. interest rate. There is no forward premium anomaly for portfolios P_4, P_5, and P_6, whose interest rates are high relative to the United States. In particular, the slope for P_5 exceeds 1. Currencies of countries whose interest rates are systematically high relative to the United States tend to depreciate in accordance with UIP.

Tables 1 and 2 are indicative of how, in our data set, currency excess returns and the forward premium anomaly are different and distinct phenomena. We find no forward premium anomaly in the portfolios that earn the largest excess returns. We do find a forward premium anomaly associated with the portfolios that earn the smallest excess returns.

Hassan and Mano (2014) showed econometrically, how the forward premium anomaly and currency excess returns are distinct phenomenon. The distinction can also be seen as follows. Let $M_{j,t}$ be the *nominal* stochastic discount factor (SDF) for country j. The investors' Euler equations for pricing nominal bonds give $r_{0,t} - r_{j,t} = \ln(E_t M_{j,t+1}) - \ln(E_t M_{0,t+1})$. In a complete markets environment (or an incomplete markets setting with no arbitrage), the stochastic discount factor approach to the exchange rate (Lustig and Verdelhan (2012)) gives $\Delta \ln(S_{j,t+1}) = \ln(M_{j,t+1}) - \ln(M_{0,t+1})$. The forward premium anomaly is a story about the covariance between relative log SDFs and relative log conditional expectations of SDFs

$$\mathrm{Cov}_t\left(\Delta \ln(S_{j,t+1}), r_{0,t} - r_{j,t}\right) = \mathrm{Cov}_t\left(\ln\left(\frac{M_{j,t+1}}{M_{0,t+1}}\right), \ln\left(\frac{E_t M_{j,t+1}}{E_t M_{0,t+1}}\right)\right),$$

being negative.

The expected currency excess return, on the other hand, is a story about relative conditional variances of the log SDFs.[5] Following from the investors' Euler equations, $E_t\left(\Delta \ln(S_{j,t+1}) + r_{j,t} - r_{0,t}\right) = \ln\left(\frac{E_t M_{0,t+1}}{E_t M_{j,t+1}}\right) - \left[E_t\left(\ln(M_{0,t+1})\right) - E_t\left(\ln(M_{j,t+1})\right)\right]$. If the stochastic discount factors are log-normally

[5]If the log SDF is not normally distributed, Backus et al. (2001) show that the expected currency excess return depends on a series of higher-ordered cumulants of the log SDFs.

distributed, the expected currency excess return simplifies to the difference in the conditional variance of the log SDFs,

$$E_t\left(\Delta \ln\left(S_{j,t+1}\right) + r_{j,t} - r_{0,t}\right) = \frac{1}{2}\left(\mathrm{Var}_t\left(\ln\left(M_{0,t+1}\right)\right) - \mathrm{Var}_t\left(\ln\left(M_{j,t+1}\right)\right)\right). \quad (5)$$

Equation (5) says country j is 'risky' and pays a currency premium if its log SDF is less volatile than country '0' (the United States). When country j residents live in relative stability, the need for precautionary saving is low. Hence, bond prices in country j will be relatively low. The relatively high returns this implies contribute to a higher currency excess return.

3 Global Macro Fundamental Risk in Currency Excess Returns

This section addresses the central issue of the paper. Does the cross-section of carry-trade-generated currency excess returns vary in proportion to their exposure to risk factors based on macro-fundamentals? Burnside et al. (2011) found little evidence that any macro-variables were priced. Lustig and Verdelhan's (2007) analysis of U.S. consumption growth as a risk factor was challenged by Burnside (2011a). Menkhoff et al. (2012) price carry-trade portfolios augmented by portfolios formed by ranking variables used in the monetary approach to exchange rates.

The macroeconomic performance indicators we consider are,

1. Unemployment rate gap, UE^{gap}
2. Change in unemployment rate, ΔUE
3. GDP growth, Δy
4. GDP gap, y^{gap}
5. Real exchange rate gap, q^{gap}
6. Real exchange rate depreciation, Δq
7. Aggregate consumption growth, Δc
8. Inflation rate, π

The rationale for unemployment, consumption growth, and GDP measures is obvious. Inflation, especially at higher levels, is associated with the economic state by depressing economic activity. We try to obtain information on the international distribution of the log SDFs through consideration of the real exchange rate gap. In the SDF approach to exchange rates, the real depreciation is the foreign-U.S. difference in log real SDFs, $\Delta q_{i,t} = n_{i,t} - n_{0,t}$. Both the gap and rates of change are employed to induce stationarity in the real exchange rate, unemployment rate, and GDP observations.

For each country, we compute time-varying (conditional) skewness $sk_t(\bullet)$, volatilities $\sigma_t(\bullet)$, and means $\mu_t(\bullet)$ of the eight variables. Our primary interest is in the higher-ordered moments, but Menkhoff et. al (2013) found first-moments to be priced, so we include them for comparison. The conditional moments are estimated by sample moments computed from a backward-looking moving 20-quarter window.[6] We then form HML versions of these variables by subtracting the average value in the bottom quartile from the average in the top quartile.

An increase in HML conditional mean variables signifies greater inequality across countries in various measures of growth. We include volatility since it is a popular measure of uncertainty. The HML conditional skewness measure provides an alternative measure of macroeconomic uncertainty highlighting the role of distributional asymmetries. High (low) skewness means a high probability of a right (left) tail event. The HML construction is similar to the interquartile range, which captures the concept of global uncertainty.[7]

3.1 Estimation

We employ the two-pass regression method used in finance to estimate how the cross-section of carry-trade excess returns are priced by the HML macroeconomic risk factors described above. Inference is drawn using generalized method of moments (GMM) standard errors as described in Cochrane (2005).

Two-pass regressions. Let $\left\{r_{i,t}^e\right\}$, $i = 1, ...N, t = 1, ..., T$, denote our collection of $N = 6$ carry-trade excess returns. Let $\left\{f_{k,t}^{HML}\right\}$, $k = 1, .., K$, be a collection of potential HML macro risk factors. In the first pass, we run $N = 6$ individual time-series regressions of the excess returns on the factors to estimate the factor 'betas' (the slope coefficients on the risk factors),

$$r_{i,t}^e = a_i + \sum_{k=1}^{K} \beta_{i,k} f_{k,t}^{HML} + \epsilon_{i,t}. \tag{6}$$

Covariance is risk, and the betas measure the extent to which the excess return is exposed to, or covaries with, the $k-th$ risk factor (holding everything else constant). If this risk is systematic and undiversifiable, investors should be compensated for bearing it. The risk should explain why some excess returns are high while others are low. This implication is tested in the second pass, which is the single cross-sectional regression of the (time-series) mean excess returns on the estimated betas,

$$\bar{r}_i^e = \gamma + \sum_{k=1}^{K} \lambda_k \beta_{i,k} + \alpha_i, \tag{7}$$

where $\bar{r}_i^e = (1/T) \sum_{t=1}^{T} r_{it}^e$ and the slope coefficient λ_k is the risk premia associated with the $k-th$ risk factor.

In other contexts, the excess return is constructed relative to what the investor considers to be the risk-free interest rate. If the model is properly specified, the intercept γ should be zero. In the current

[6] We also considered using a 16-quarter and a 24-quarter window. The results are robust to these alternative window lengths. These results are reported in Appendix C.

[7] We point out that there is a literature that attempts to measure macroeconomic uncertainty. For example, Baker, Bloom, and Davis (2013) build their measure by counting the frequency with which newspaper articles mention words like 'policy uncertainty,' and Jurado, Ludvigson, and Ng (2013) which is based on the conditional volatility of forecast errors. In contrast, our measures are comparatively low tech and easily computable.

setting, the carry trades are available to global investors. When the trade matures, the payoff needs to be repatriated to the investor's home currency, which entails some foreign exchange risk. Hence, the excess returns we consider are not necessarily relative to 'the' risk-free rate, and there is no presumption that the intercept γ is zero.

To draw inference about the λs, we recognize that the betas in equation (7) are not data themselves, but are estimated from the data. To do this, we compute the GMM standard errors, described in Cochrane (2005) and Burnside (2011b), that account for the generated regressors problem and for potential serial correlation and heteroskedasticity in the errors. Cochrane (2005) sets up a GMM estimation problem using a constant as the instrument, which produces the identical point estimates for $\beta_{i,k}$ and λ_k as in the two-pass regression. The 'pricing errors,' α_i, should be zero if the model adequately describes the data. Also available is the GMM covariance matrix of the residuals α_i, which we use to test that they are jointly zero. We get our point estimates by doing the two-pass regressions with least squares and get the standard errors by 'plugging in' the point estimates into the GMM formulae. Additional details are given in Appendix B.

3.2 Empirical Results

We begin by estimating a single-factor model with the two-pass procedure, where the single factor is one of the HML global macro risk factors discussed above. Table 3 shows the the second stage estimation results for the single-factor model. In the first row, we see that the HML unemployment gap skewness factor is priced in the excess returns. The price of risk λ is positive, the t-ratio is significant, the R^2 is very high, and the constant γ is not significant.

Several other factor candidates also appear to be priced, such as two other HML conditional skewness measures ($sk_t\left(\Delta UE\right)$ and $sk_t\left(\Delta y\right)$) and HML conditional volatilities of y^{gap}, Δy, and Δc, and conditional means of UE^{gap} and y^{gap}. For these factor candidates, the t-ratios on λ estimates are significant, the estimated intercepts γ are insignificant, and many of the R^2 values are also quite high. However, it is not the case that generically formed HML specifications on conditional moments of macro fundamentals will automatically get priced. The HML conditional volatilities of unemployment rate changes and real exchange rate changes are not priced, and these specifications have R^2 values near zero.

The single-factor results give an informal impression that the HML $sk_t\left(UE^{gap}\right)$ factor dominates the alternative measures of the global risk factor. The price of risk has the highest t-ratio and the regression has the highest R^2. Figure 4 displays the scatter plot of the average portfolio currency excess returns against their HML unemployment gap skewness betas.

To assess more formally the impression that HML $sk_t(UE^{gap})$ dominates, we estimate a two-factor model with the HML $sk_t\left(UE^{gap}\right)$ as the maintained (first) factor and each of the alternative factor constructions as the second factor. Table 4 shows the two-factor estimation results.

Here, the HML unemployment gap skewness factor is significant at the 5% level in every case, while only HML $sk_t(\Delta UE)$ is significantly priced as a second factor at the 5% level. The two variables are alternative constructions to measure the same concept. We continue to find the constant and the Wald test on the pricing errors to be insignificant. These results suggest that the HML unemployment gap skewness factor is the global macro risk factor for carry trade excess returns.

To delve deeper into the risk-return relationship, Table 5 reports the decomposition of the betas of the log-approximate portfolio excess returns into contributions from the interest differential and the exchange rate return components.[8] Notice that the betas on exchange rate returns are uniformly negative. Currency values of all the portfolios decline relative to the USD in times when the factor is high. Similarly, the betas for the interest differentials for P_1 through P_4 are negative. However, the beta for the interest differentials for P_5 and P_6 are positive. For P_6, the magnitude is so large that it more than offsets the negative beta for the exchange rate return. Why do yields increase for these portfolios in times of high global uncertainty? Because global investors flee the debt of these countries in the bad state which drives bond prices down and yields up.

A visual of the factor is presented in Figure 5, which plots the high, low, and high-minus-low average values of skewness of the unemployment gap. Low skewness is typically negative. The figure also shows European and U.S. business cycle dating. The correspondence between the factor and U.S. and European business cycles is positive only about half of the time. Since the factor samples economies beyond the United States and Europe, the imperfect correspondence might be expected.

To see which countries are key in constructing the factor, Table 6 lists the top 10 countries that appear most frequently in construction of the HML unemployment gap skewness factor. They are roughly a mix of developed and emerging economies.

Pre-Crisis Sample. Since we are using quarterly observations due to the availability of the macro variables, we do not have a surplus of time-series observations. Nevertheless, we can do some limited subsample analyses. Here, we ask if our results are driven by the global financial crisis. Lustig and Verdelhan (2011) point specifically to the poor performance of the carry trade during the crisis as an example of the risk borne by international investors in the carry trade. To answer this question, we end the sample in 2008Q2. Table 7 shows the mean excess returns and Sharpe ratios for the interest rate sorted portfolios over this time span. Again, there is little difference between P_4 and P_5 average excess returns, but there is a large spread between returns on P_6 and P_1.

Table 8 shows the results from the single-factor estimation over the pre-crisis sample. The HML $sk_t\,(UE^{gap})$ factor again gives the highest R^2. Fewer of the alternative factor measures are significantly priced. This could be because they were more pronounced during the crisis or because we have a smaller sample, having lost 24 quarterly observations-a reduction of 16% of the time-series observations.

In Table 9, we evaluate robustness in the pre-crisis subsample, by maintaining HML $sk_t\,(UE^{gap})$ as the first factor and alternating the second factor. HML skewness in the unemployment gap remains significant at the 5% level in 18 specifications and at the 10% level in 3 specifications. The only alternative factor that is significantly priced at the 5% level is the HML conditional volatility of consumption growth.

[8]Statistical significance of the betas are not the key issue as the GMM standard error estimates on the λ estimates take into account that the betas are estimated.

Sorting excess currency returns by beta. In the foregoing analysis, we sorted countries into portfolios and found that their excess returns varied proportionately with their betas on the HML $sk_t\,(UE^{gap})$ factor. Additional evidence that this variable provides a risk-based explanation would be if the betas of individual excess returns vary and are increasing in those returns. To investigate along these lines, for each individual currency i, at time t, we create an excess return by going long (short) in that currency if its interest rate is higher (lower) than the U.S. interest rate. We then estimate beta for each currency individually and sort the excess returns into portfolios by their beta.

Table 10 shows the average excess returns from sorting into six beta-ranked portfolios. They are low for low-beta portfolios and high for high-beta portfolios. While they do not increase monotonically, average excess returns rise monotonically if we sort less finely into three quantiles instead of six.

Looking at individual currency excess returns reveals there are both positive beta and negative beta currencies. Table 11 shows the individual country betas and excess returns associated with the low and high tertile beta countries. The identification by individual country, while not exact, shows a clear tendency for excess returns to be correlated with betas. Figure 6 shows the scatter plot for all of the currency excess returns against their betas. Next, we eliminate those European countries that adopted the euro. Figure 7 shows, for this subsample of countries, the scatter plot of mean currency excess returns against their betas.

Our results share similarities with Lustig et al. (2011). In both papers, the global risk factor connects with the concept of global macroeconomic uncertainty. Their relative asset pricing work identifies the HML currency excess returns between P_6 and P_1 portfolios as the global risk factor, which they argue is associated with changes in global equity market volatility.

Developed Countries. Are our results driven entirely by emerging market economies? To address this question, we restrict the sample to developed economies.[9] As seen in Table 12, the factor is significantly priced into the 6 portfolio excess returns formed only by developed countries.

Relation to the U.S. Stock Market. In finance, the three Fama-French (1996) factors consisting of the market return, the HML book-to-market return, and the HML size ranked return, adequately price (virtually all) U.S. stock returns. Hence, knowing these three factors implies that one knows everything about the systematic behavior of any stock return. The only thing left to learn about is the behavior of these three factor portfolio returns. Due to the pervasive explanatory power of the Fama-French factors, we ask if they also price currency excess returns in our data set. If they do, then there is no puzzle associated with the carry trade. Table 14 shows results of this two-step estimation. These data were obtained from Kenneth French's web site.[10]

In the table, λ_{mkt} is the price of risk on the market excess return, λ_{SMB} is the price of risk on the small minus big firm excess return, and λ_{value} is the price of risk on the low to high book-to-market excess return.

[9] We omit Brazil, Chile, Colombia, Czech Republic, Hungary, Indonesia, India, Malaysia, Mexico, Philippines, Romania, South Africa, Thailand, and Turkey.

[10] http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html

In Column A, we include the three Fama-French factors. Only the SMB factor is significant. When we add the HML $sk_t\left(UE^{gap}\right)$ (Column B), none of the candidate factors are significantly priced. There is a degrees of freedom issue with six portfolios being priced by four factors, so in Columns C, D, and E, we maintain $sk_t\left(UE^{gap}\right)$ and include the Fama-French factors individually. In two cases, $sk_t\left(UE^{gap}\right)$ is significantly priced at the 5% level and in the third case, is nearly significant, while none of the Fama-French factors are significant.

While we find that HML $sk_t\left(UE^{gap}\right)$ prices currency excess returns, it does not price U.S. equities. In Table 14, we estimate the beta-risk model for the 25 Fama and French (1996) double-sorted test portfolios over the time-span of our sample.[11] Column A includes the three Fama-French factors and our HML $sk_t\left(UE^{gap}\right)$, which is not significantly priced into the U.S. equity portfolios during our sample. Column B shows estimates using HML $sk_t\left(UE^{gap}\right)$ as a single-factor are little changed.

4 Interpretation

The empirical work above does not say that countries with high (low) unemployment gap skewness have high (low) interest rates and pay out high (low) currency excess returns. It says investors pay attention to the HML $sk_t(UE^{gap})$ factor, as a global risk factor. We draw on a no-arbitrage model for interest rates and exchange rates as an interpretative framework for the empirical results. The model is closely related to Backus et al. (2001) and Lustig et al. (2011)'s affine-yield models of the term structure to pricing currency excess returns. To ease notation, we will call the global risk factor $z_{g,t}$ = HML $sk_t(UE^{gap})$. We model the way investors pay attention to this global risk factor by letting the global risk factor ($z_{g,t}$) and a country-specific risk factor ($z_{i,t}$) load on a country's log nominal SDF ($m_{i,t+1}$) according to

$$m_{i,t+1} = -\theta_i\left(z_{i,t} + z_{g,t}\right) - u_{i,t+1}\sqrt{\omega_i z_{i,t}} - u_{gt+1}\sqrt{\kappa_i z_{i,t} + \delta_i z_{g,t}}, \tag{8}$$

where

$$\begin{aligned} z_{g,t+1} &= \left(1-\phi_g\right)\chi_g + \phi_g z_{g,t} + u_{g,t+1}\sqrt{z_{g,t}}, && (9)\\ z_{i,t+1} &= \left(1-\phi_i\right)\chi_i + \phi_i z_{i,t} + u_{i,t+1}\sqrt{z_{i,t}}, && (10)\\ u_{g,t} &= \sigma_g v_{g,t}, && (11)\\ u_{i,t} &= \sigma_i\left(\rho_i v_{g,t} + v_{i,t}\sqrt{(1-\rho_i^2)}\right), && (12) \end{aligned}$$

and $v_{g,t}$ and $v_{i,t}$ are independent standard normal variates. Since the global factor must be built from an aggregation of country factors, we allow the country-specific innovation to be correlated with the global innovation $E\left(u_{i,t}u_{g,t}\right) = \rho_i$.

[11] From Kenneth French's website: the portfolios, constructed at the end of each June, are the intersections of 5 portfolios formed on size (market equity, ME) and 5 portfolios formed on the ratio of book equity to market equity (BE/ME). The size breakpoints for year t are the NYSE market equity quintiles at the end of June of year t. BE/ME for June of year t is the book equity for the last fiscal year end in $t-1$ divided by ME for December of $t-1$. The BE/ME breakpoints are NYSE quintiles.

The conditional mean ($\mu_{i,t}$) and conditional variance ($V_{i,t}$) of the log SDF are

$$
\begin{aligned}
\mu_{i,t} &= -\theta_i \left(z_{i,t} + z_{g,t}\right) \\
V_{i,t} &= \sigma_g^2 \delta_i z_{g,t} + \left(\sigma_g^2 \kappa_i + \sigma_i^2 \omega_i\right) z_{i,t} + 2\sigma_g \sigma_i \rho_i \sqrt{\omega_i z_{i,t}} \sqrt{\kappa_i z_{i,t} + \delta_i z_{g,t}}.
\end{aligned}
$$

From investor Euler equations, we obtain the pricing relationships

$$
\begin{aligned}
r_{i,t} &= \mu_{i,t} + 0.5 V_{i,t}, \\
\Delta s_{i,t} &= m_{i,t} - m_{0,t}, \\
R^e_{i,t+1} &= 0.5\left(V_{0,t} - V_{i,t}\right) + \epsilon_{i,t+1},
\end{aligned}
$$

where $R^e_{i,t+1} = \Delta s_{i,t+1} + r_{i,t} - r_{0,t}$ is the excess dollar return. The last equation comes from $E_t\left(R^e_{i,t+1}\right) = 0.5\left(V_{0,t} - V_{i,t}\right)$ and $\epsilon_{i,t+1}$ is the expectational error.

Countries with high $\mu_{i,t}$ and $V_{i,t}$ will have high interest rates. But for country i to also pay the carry-trade excess return, it must have low $V_{i,t}$ relative to $V_{0,t}$. This suggests a pattern of high $\mu_{i,t}$ and low $V_{i,t}$ to explain the data. The usual story is one of the precautionary saving motive. If $V_{i,t}$ is low relative to $V_{0,t}$, there is little need for precautionary saving. Bond prices in i will therefore be low and yields high. We note that heterogeneity in the risk-factor loadings on the log SDFs is not necessary to generate differences in conditional variances. Differences in the realizations of country-specific risk $z_{i,t}$ will do that. What is key, however, is that the log SDFs load on the global factor z_{gt}. If they do not, excess currency returns may be non-zero, but they will not be priced by the global risk factor.

We estimate the model by simulated method of moments.[12] We begin by estimating the process for the global risk factor (the HML skewness of the unemployment gap) $z_{g,t}$ separately. Parameters in equation (9) are estimated by simulated method of moments and are shown in Table 15.

Recall that we do not have a balanced panel. The time-span coverage varies by availability. Our data sample consists of 41 countries that can be bilaterally paired with the United States (country '0'). Of these 41 countries, 38 have sufficiently long time-series data that we use in the estimation. Estimation is done bilaterally. The 14 moments we use in the estimation are $E\left(h_{i,t}\right)$, where

$$
h'_{i,t} = \left(\Delta s_{i,t}, \Delta s^2_{i,t}, \Delta s_{i,t}\Delta s_{i,t-1}, \Delta s_{i,t}\Delta s_{i,t-4}, R^e_{i,t}, \left(R^e_{i,t}\right)^2, R^e_{i,t}R^e_{i,t-1}, R^e_{i,t}R^e_{i,t-4}, r_{i,t}, r^2_{i,t}, r_{i,t}r_{i,t-1}, r_{0,t}, r^2_{0,t}, r_{0,t}r_{0,t-1}\right).
$$

Table 16 shows the average of the parameter estimates. We also estimate two restricted versions of the model. In one version, the SDFs do not load on the global factor. In the other restricted model, the SDFs load only on the global factor but not on the country-specific factors. There is substantial heterogeneity across individual estimates. In the unconstrained model, the U.S. SDF loads more heavily on the global risk factor (δ) and on the country-specific component (ω) than the average on the other countries.

We simulate the three versions of the estimated model. In each of the 2,000 simulations, we generate 87 observations on exchange rate returns and interest rates across the 38 countries and the United States. In the data, we had, on average, 87 time-series observations. For each replication, we sort currencies

[12]See Lee and Ingram (1991). 100,000 is the length of the simulated time series.

into six interest rate ranked portfolios, compute their mean excess (over the U.S.) returns and Sharpe ratios, and estimate the single-factor beta-risk model. Table 17 reports the median values over the 2,000 simulations.

The simulated carry-trade-generated average excess returns are increasing as one moves from P_1 to P_6 (panel A). Volatility of simulated excess returns when SDFs do not load on country-specific factors and in the unconstrained model are too low, making the Sharpe ratios too high. When SDFs do not load on the global factor, there is no forward premium anomaly in the portfolios. Here, the Fama slope (in the regression of equation (4) on the simulations) is slightly positive (except P_6) but close to zero and does not vary across the portfolios. When the SDFs do not load on country-specific factors, a forward premium anomaly emerges but the slope does not vary across portfolios.

In panel B, the median estimates of the beta-risk model are shown. Here, it is verified that the global risk factor is unpriced if the SDFs do not load on that factor. The median t-ratio on λ is far from 2, even though the median R^2 value is quite high. The global risk factor is priced in the unconstrained model. The median R^2 is similar to that obtained from the data, while the estimated risk premium λ is overstated.

The point of the exercise in this section is not to replicate exactly the moments of the data but to illustrate the link between the global factor and carry-trade-generated excess returns. The unconstrained model captures three broad features of the data. Average excess returns are generally increasing in the carry-trade portfolios P_1 through P_6, the forward premium anomaly is more pronounced when portfolio interest rates are more similar to U.S. rates, and investor SDFs must load on the global factor. Probably, their SDFs load also on country-specific factors as well.

5 Conclusion

It has long been understood that systematic currency excess returns (deviations from uncovered interest parity) are available to investors. Less well understood is what risks are being compensated for by the excess returns. In a financially integrated world, excess returns should be driven by common factors. We find that a global risk factor, constructed as the high-minus-low conditional skewness of the unemployment gap, is priced into carry-trade-generated excess returns. Carry-trade-generated currency excess returns compensate for global macroeconomic risks.

There are three notable features of this risk factor. First, it is a macroeconomic fundamental. As Lustig and Verdelhan (2011) point out, since the statistical link between asset returns and macroeconomic factors is always weaker than the link between asset returns and return-based factors, the high explanatory power provided by this factor and its significance is noteworthy. Second, the factor is global in nature. It is constructed from averages of countries in the top and bottom quartiles of the unemployment gap skewness. Since the portfolios of carry-trade-generated excess returns are available to global investors, only global risk factors should be priced. Third, the factor measures something different from standard measures of global uncertainty. Unlike the standard measures of uncertainty, the HML global macro risk factor can capture asymmetries in the distribution of the global state that reflect the divergence, disparity, and inequality of fortunes across countries.

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Figure 1: Cumulated Excess Returns on Six Carry Portfolios

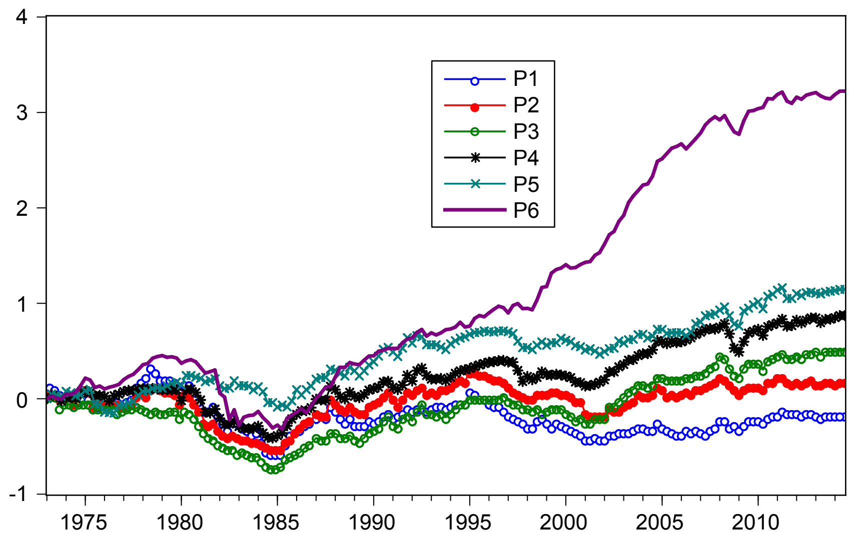


Figure 2: Cumulated Excess Returns on P_6 Carry Portfolio and the Standard and Poor's 500

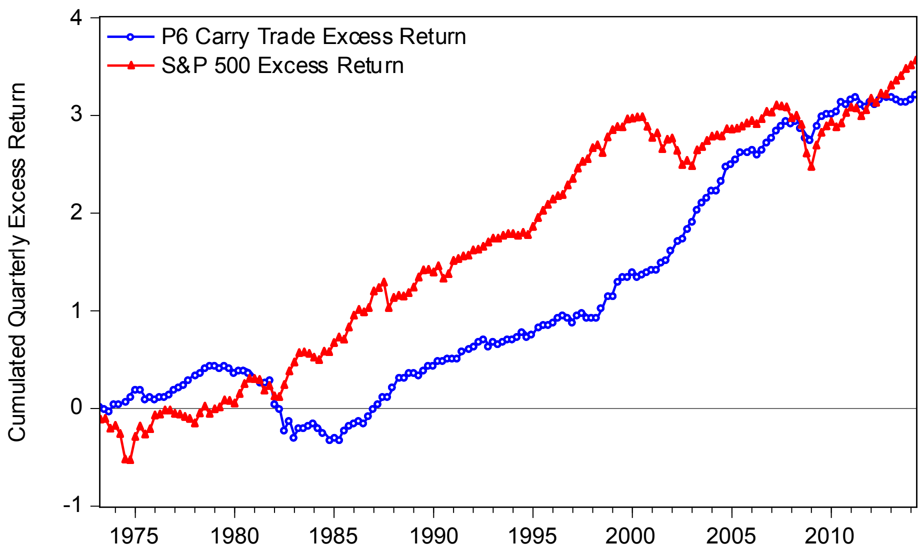


Figure 3: Decomposition of Carry Excess Returns (Log Approximation)

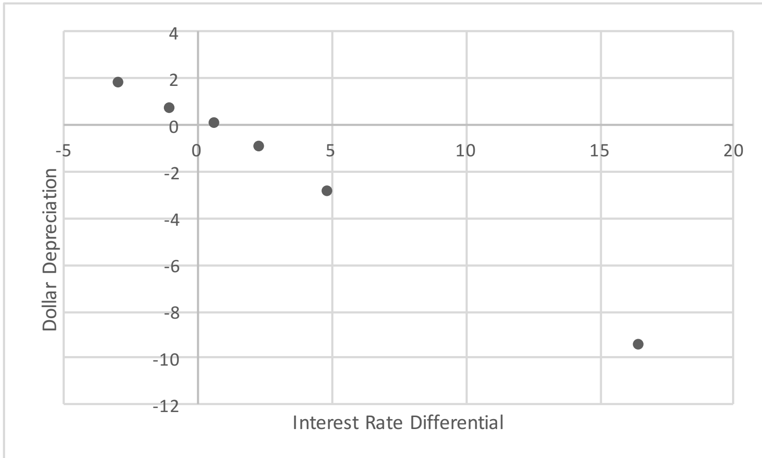
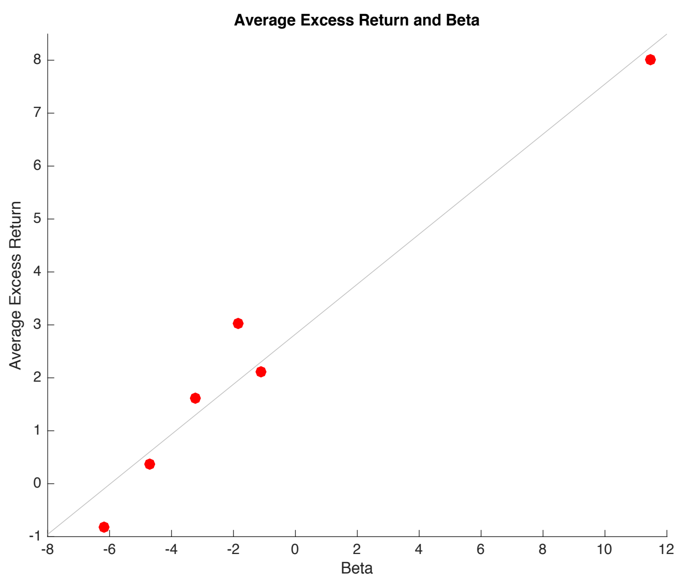


Figure 4: Average Excess Returns and Betas by HML Unemployment Gap Skewness Beta Model, 1978Q1–2014Q2



Notes: The raw data are quarterly (1973Q1 to 2014Q2) and, when available, are end-of-quarter and point-sampled. For each country (41 countries plus the euro area), we compute the 'conditional' unemployment gap skewness using a 20-quarter window. To form the portfolio returns, we sort by the nominal interest rate for each country from low to high. The rank ordering is divided into six categories, into which the currency returns are assigned. P_6 is the portfolio of returns associated with the highest interest rate quantile and P_1 is the portfolio of returns associated with the lowest interest rate quantile. The excess returns are the average of the USD returns in each category minus the U.S. nominal interest rate and are stated in percent per annum. The figure plots portfolio average excess returns against their betas.

Figure 5: High, Low, and High-Minus-Low (HML) Unemployment Gap Skewness, U.S. and European Recessions

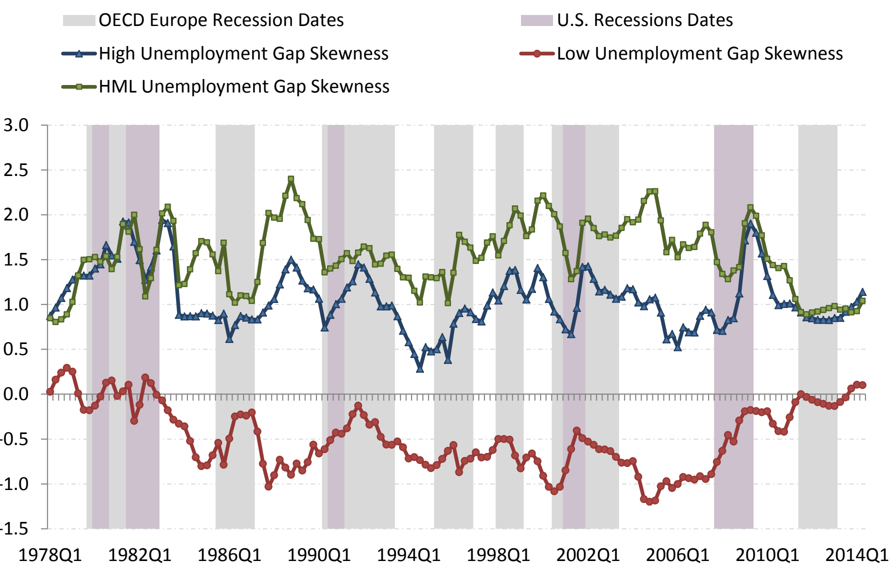


Figure 6: Country Mean Excess Returns and Betas

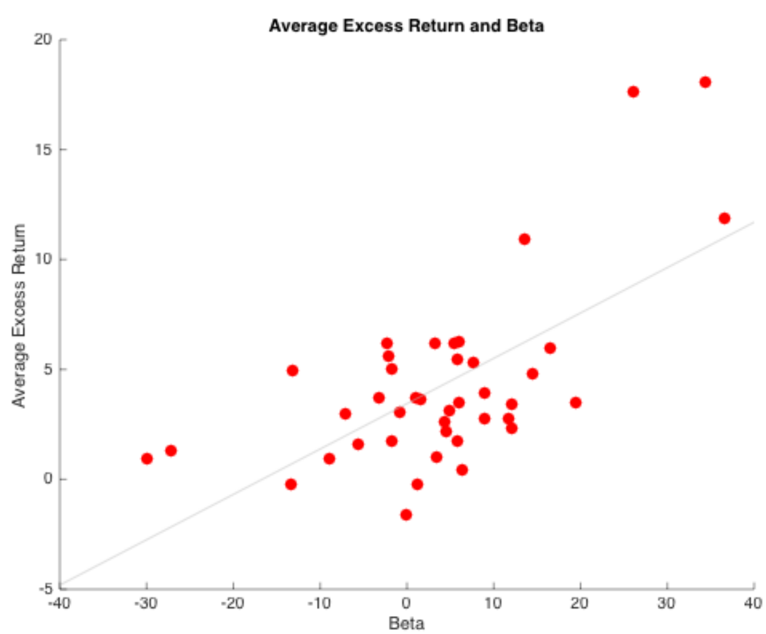


Figure 7: Non-Euro Mean Excess Returns and Betas

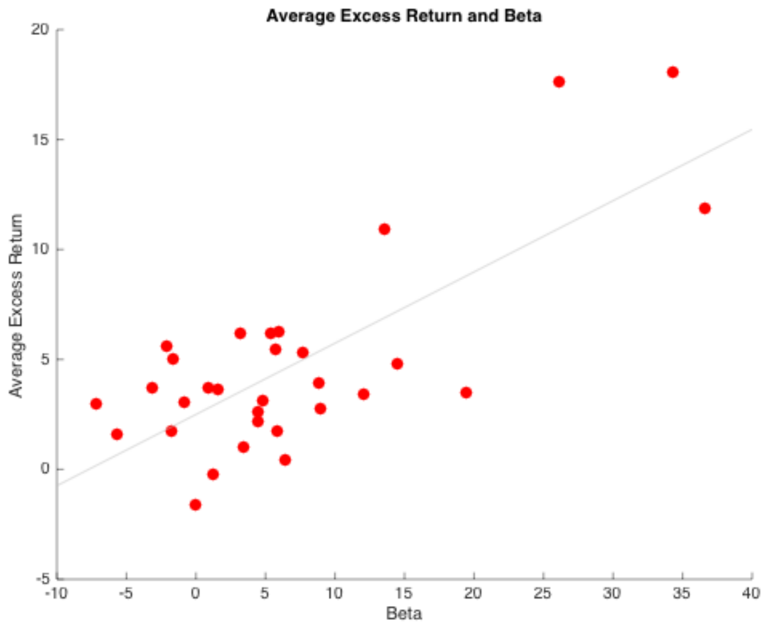


Table 1: Decomposition of Carry Excess Returns (Log Approximation), 1978Q1–2014Q2

	P_1	P_2	P_3	P_4	P_5	P_6
A. Portfolio Returns and Sharpe Ratios						
Mean return	4.929	5.723	6.735	7.408	8.027	13.031
Sharpe Ratio	0.277	0.320	0.427	0.409	0.427	0.558
Mean excess return	-1.161	-0.367	0.645	1.317	1.925	6.899
Sharpe Ratio	-0.064	-0.020	0.039	0.072	0.104	0.285
B. Components of Carry Excess Returns						
Interest rate differential	-2.916	-1.055	0.669	2.326	4.842	16.366
USD exchange rate depreciation	1.755	0.688	-0.024	-1.008	-2.917	-9.466

Notes: These are log-approximated returns and excess returns. USD exchange rate depreciation is Δs_{t+1} and is positive when the dollar falls in value. Interest rate differential is $r_t^{P_j} - r_{0,t}$, where $r_t^{P_j} \equiv \frac{1}{n_{j,t}} \sum_{i \in P_j} r_{j,t}$.

Table 2: Fama Regressions (Log Approximation), 1978Q1–2014Q2

Dependent Variable	Regressor	Slope	t-ratio ($\beta = 0$)	t-ratio ($\beta = 1$)
$\Delta s_{t+1}^{P_6}$	$r_{0,t} - r_t^{P_6}$	0.580	3.328	-2.407
$\Delta s_{t+1}^{P_5}$	$r_{0,t} - r_t^{P_5}$	1.302	2.174	0.504
$\Delta s_{t+1}^{P_4}$	$r_{0,t} - r_t^{P_4}$	0.251	0.353	-1.057
$\Delta s_{t+1}^{P_3}$	$r_{0,t} - r_t^{P_3}$	-0.942	-1.387	-2.859
$\Delta s_{t+1}^{P_2}$	$r_{0,t} - r_t^{P_2}$	-0.424	-0.675	-2.268
$\Delta s_{t+1}^{P_1}$	$r_{0,t} - r_t^{P_1}$	-0.561	-1.134	-3.153

Table 3: Two-Pass Estimation of the Single-Factor Beta-Risk Model on Carry Excess Returns, 1978Q1–2014Q2

	Single-Factor Model						
HML Factor	λ	t-ratio	γ	t-ratio	R^2	Test-stat	p-val.
$sk_t(UE^{gap})$	0.473	**3.090**	2.824	1.328	0.958	1.339	0.931
$sk_t(\Delta UE)$	0.577	**2.595**	2.351	0.875	0.893	2.849	0.723
$sk_t(y^{gap})$	0.614	1.320	6.134	1.095	0.559	4.042	0.543
$sk_t(\Delta y)$	0.536	**2.106**	4.796	1.807*	0.388	4.994	0.417
$sk_t(q^{gap})$	-0.664	-1.559	-0.528	-0.201	0.166	5.792	0.327
$sk_t(\Delta q)$	1.491	1.791*	7.620	**2.016**	0.446	3.185	0.671
$sk_t(\Delta c)$	0.274	1.669*	1.872	0.693	0.413	7.892	0.162
$sk_t(\pi)$	0.691	0.763	-9.086	-0.686	0.298	1.705	0.888
$\sigma_t(UE^{gap})$	2.684	1.419	2.529	1.014	0.122	7.894	0.162
$\sigma_t(\Delta UE)$	-0.028	-0.901	2.673	1.559	0.016	10.610	0.060
$\sigma_t(y^{gap})$	0.679	**2.224**	-0.008	-0.003	0.645	4.622	0.464
$\sigma_t(\Delta y)$	1.331	**2.031**	-0.838	-0.317	0.617	6.057	0.301
$\sigma_t(q^{gap})$	-5.491	-1.723*	2.246	0.786	0.333	4.434	0.489
$\sigma_t(\Delta q)$	4.938	1.470	0.225	0.114	0.124	9.082	0.106
$\sigma_t(\Delta c)$	1.857	**1.967**	-2.762	-0.755	0.706	4.177	0.524
$\sigma_t(\pi)$	1.501	1.450	1.091	0.547	0.215	9.826	0.080
$\mu_t(UE^{gap})$	3.100	**1.986**	0.926	0.292	0.829	5.285	0.382
$\mu_t(\Delta UE)$	0.072	1.891*	-0.064	-0.023	0.526	8.826	0.116
$\mu_t(y^{gap})$	0.477	**2.420**	0.863	0.324	0.807	5.405	0.368
$\mu_t(\Delta y)$	-0.174	-0.262	2.490	1.483	0.001	12.192	0.032
$\mu_t(q^{gap})$	1.140	1.353	1.731	0.956	0.036	9.163	0.103
$\mu_t(\Delta q)$	2.285	1.392	0.106	0.048	0.181	7.752	0.170
$\mu_t(\Delta c)$	-1.492	-1.936*	4.221	1.529	0.138	5.947	0.311
$\mu_t(\pi)$	-1.791	-0.972	2.615	1.275	0.021	10.208	0.070

Notes: The raw data are quarterly (1973Q1 to 2014Q2) and, when available, are end-of-quarter and point-sampled. 20 quarters start-up to compute initial HML factors. Model estimated on returns from 1978Q1 to 2014Q2. Δy, y^{gap}, Δc, ΔUE, UE^{gap}, π, q^{gap}, and Δq represent the GDP growth rate, GDP gap, consumption growth rate, change in the unemployment rate, unemployment gap, inflation rate, real exchange rate gap, and real exchange rate depreciation, respectively. For each country (41 countries plus the euro area) and each macroeconomic variable (x), we compute the 'conditional' mean ($\mu_t(x)$), volatility ($\sigma_t(x)$), and skewness ($sk_t(x)$) using a 20-quarter window. To form the portfolio returns, we sort by the nominal interest rate (carry) for each country from low to high. The rank ordering is divided into six portfolios, into which the currency returns are assigned. P_6 is the portfolio of returns associated with the highest nominal interest rate countries and P_1 is the portfolio of returns associated with the lowest nominal interest rate countries. This table reports the two-pass procedure estimation results from a one-factor model. In the first pass, we run $N = 6$ individual time-series regressions of the excess returns on the K factors to estimate the factor 'betas,' $r^e_{i,t} = a_i + \sum_{k=1}^{K} \beta_{i,k} f^{HML}_{k,t} + \epsilon_{i,t}$, where $r^e_{i,t}$ is the excess return, $\beta_{i,k}$ is the factor beta, and $f^{HML}_{k,t}$ is the high-minus-low (HML) macro risk factor. The factors considered include the HML values of the conditional mean, volatility, and skewness of Δy, y^{gap}, Δc, ΔUE, UE^{gap}, π, q^{gap}, and Δq. Each HML value is equal to the average in the highest quartile minus the average in the lowest quartile. In the second pass, we run a single cross-sectional regression of the (time-series) mean excess returns on the estimated betas, $\bar{r}^e_i = \gamma + \sum_{k=1}^{K} \lambda_k \beta_{i,k} + \alpha_i$, where $\bar{r}^e_i$ is the average excess return, γ is the intercept, λ_k is the risk premia, and α_i is the pricing error. The table reports the price of risk (λ) and its associated t-ratio (using GMM standard errors), the estimated intercept (γ) and its associate t-ratio, R^2 and the Wald test on the pricing errors (Test-stat) and its associated p-value (p-val.). Bold indicates significance at the 5% level. '*' indicates significance at the 10% level.

Table 4: Two-Pass Estimation of Two-Factor Beta-Risk Model on Carry Excess Returns, 1978Q1–2014Q2

		Two-Factor Model (First HML Factor is $sk_t(UE^{gap})$)							
		2^{nd} HML							
λ_1	t-ratio	Factor	λ_2	t-ratio	γ	t-ratio	R^2	Test-stat	p-val.
0.420	**2.603**	$sk_t(\Delta UE)$	0.393	**2.098**	2.510	1.130	0.943	1.530	0.910
0.474	**3.199**	$sk_t(y^{gap})$	0.189	0.839	2.506	0.946	0.960	1.012	0.962
0.474	**3.069**	$sk_t(\Delta y)$	0.099	0.369	2.653	1.002	0.959	0.919	0.969
0.425	**2.931**	$sk_t(q^{gap})$	-0.090	-0.277	2.500	0.927	0.950	1.410	0.923
0.472	**2.886**	$sk_t(\Delta q)$	0.005	0.008	2.846	0.865	0.958	1.196	0.945
0.555	**2.802**	$sk_t(\Delta c)$	-0.037	-0.238	3.127	1.401	0.987	0.929	0.968
0.570	**2.619**	$sk_t(\pi)$	-0.338	-0.810	8.367	1.128	0.993	0.132	1.000
0.356	**3.121**	$\sigma_t(UE^{gap})$	-0.494	-0.411	2.853	1.261	0.897	3.988	0.551
0.388	**3.152**	$\sigma_t(\Delta UE)$	-0.055	-1.396	2.794	1.189	0.865	4.043	0.543
0.307	**2.705**	$\sigma_t(y^{gap})$	0.151	0.932	2.150	0.916	0.928	2.720	0.743
0.413	**3.026**	$\sigma_t(\Delta y)$	-0.002	-0.004	2.506	1.054	0.844	4.366	0.498
0.458	**2.702**	$\sigma_t(q^{gap})$	-4.277	-1.658*	2.206	0.958	0.948	1.917	0.860
0.368	**3.006**	$\sigma_t(\Delta q)$	-1.212	-0.473	2.892	1.215	0.848	4.131	0.531
0.301	**2.458**	$\sigma_t(\Delta c)$	0.570	1.299	0.733	0.279	0.867	4.598	0.467
0.395	**3.082**	$\sigma_t(\pi)$	-0.477	-0.625	2.655	1.108	0.859	3.578	0.612
0.286	**2.431**	$\mu_t(UE^{gap})$	1.044	0.816	2.386	1.007	0.903	3.156	0.676
0.394	**3.012**	$\mu_t(\Delta UE)$	-0.009	-0.298	2.590	1.146	0.841	4.051	0.542
0.377	**2.325**	$\mu_t(y^{gap})$	0.116	0.501	2.870	1.224	0.902	3.058	0.691
0.522	**2.590**	$\mu_t(\Delta y)$	-1.744	-1.394	4.138	1.259	0.914	2.598	0.762
0.382	**3.209**	$\mu_t(q^{gap})$	0.668	0.805	2.308	1.051	0.906	3.944	0.558
0.376	**3.141**	$\mu_t(\Delta q)$	-0.559	-0.491	2.851	1.224	0.848	4.180	0.524
0.465	**3.057**	$\mu_t(\Delta c)$	-1.214	-1.574	3.454	1.321	0.888	3.141	0.678
0.335	**2.563**	$\mu_t(\pi)$	-1.751	-0.781	2.673	0.991	0.867	3.593	0.609

Notes: See notes to Table 3.

Table 5: Excess Return and Beta Decomposition, 1978Q1–2014Q2

	P_1	P_2	P_3	P_4	P_5	P_6
Total Excess Return	-1.161	-0.367	0.645	1.317	1.925	6.899
Beta	-6.367	-5.112	-3.698	-1.790	-2.246	11.437
t-ratio	-1.631	-1.325	-1.105	-0.485	-0.629	2.257
Interest Differential	-2.916	-1.055	0.669	2.326	4.842	16.366
Beta	-0.517	-0.700	-0.723	-0.168	1.229	13.819
t-ratio	-0.476	-0.927	-1.317	-0.258	1.463	3.651
Exchange Return	1.755	0.688	-0.024	-1.008	-2.917	-9.466
Beta	-5.850	-4.412	-2.975	-1.622	-3.475	-2.382
t-ratio	-1.475	-1.120	-0.936	-0.444	-0.922	-0.436

Notes: Log approximated currency excess returns. t-ratios are computed by Newey-West. Factor is HML unemployment gap skewness.

Table 6: Top 10 Countries that Appear Most Frequently in the High and Low Unemployment Gap Skewness Categories

Country	Proportion of Times in High Group	Country	Proportion of Times in Low Group
Australia	0.473	Norway	0.390
Canada	0.404	United States	0.295
Taiwan	0.253	Denmark	0.281
Switzerland	0.247	Philippines	0.281
Singapore	0.240	Japan	0.247
United States	0.212	New Zealand	0.240
Sweden	0.192	Mexico	0.205
United Kingdom	0.185	Brazil	0.199
Mexico	0.185	Hungary	0.192
Poland	0.185	Canada	0.185

Table 7: Pre-Crisis Carry Excess Return Summary Statistics, 1978Q1–2008Q2

	P_1	P_2	P_3	P_4	P_5	P_6
Mean Excess	-1.673	-0.445	0.188	1.229	1.726	7.417
Sharpe Ratio	-0.084	-0.023	0.011	0.069	0.096	0.323

Table 8: Pre-Crisis Two-Pass Estimation of Single-Factor Beta-Risk Model on Carry Excess Returns, 1978Q1–2008Q2

			Single-Factor Model				
HML Factor	λ	t-ratio	γ	t-ratio	R^2	Test-stat	p-val.
$sk_t(UE^{gap})$	0.484	**2.392**	4.211	1.464	0.928	1.248	0.940
$sk_t(\Delta UE)$	0.606	**2.180**	2.991	0.902	0.840	2.343	0.800
$sk_t(y^{gap})$	0.750	1.042	5.457	0.744	0.843	1.514	0.911
$sk_t(\Delta y)$	0.600	1.651*	4.654	1.311	0.410	3.476	0.627
$sk_t(q^{gap})$	0.258	0.972	3.621	1.612	0.017	8.955	0.111
$sk_t(\Delta q)$	1.155	1.550	7.099	1.899*	0.193	4.383	0.496
$sk_t(\Delta c)$	0.297	1.464	0.997	0.293	0.507	6.508	0.260
$sk_t(\pi)$	0.432	0.707	-4.545	-0.556	0.148	2.668	0.751
$\sigma_t(UE^{gap})$	4.176	1.419	3.233	0.807	0.262	5.235	0.388
$\sigma_t(\Delta UE)$	-0.041	-1.103	2.860	1.307	0.028	9.720	0.084
$\sigma_t(y^{gap})$	0.597	**2.325**	1.376	0.403	0.687	4.521	0.477
$\sigma_t(\Delta y)$	1.039	1.955*	0.070	0.025	0.637	6.417	0.268
$\sigma_t(q^{gap})$	-2.056	-1.429	2.332	1.135	0.059	8.705	0.121
$\sigma_t(\Delta q)$	4.469	1.104	0.848	0.376	0.073	9.432	0.093
$\sigma_t(\Delta c)$	1.012	**2.926**	-0.989	-0.332	0.771	5.594	0.348
$\sigma_t(\pi)$	1.395	1.134	1.264	0.526	0.142	9.415	0.094
$\mu_t(UE^{gap})$	3.386	1.771*	0.6	0.148	0.857	3.847	0.572
$\mu_t(\Delta UE)$	0.073	1.81*	-0.049	-0.015	0.571	8.01	0.156
$\mu_t(y^{gap})$	0.529	**2.129**	0.942	0.277	0.817	5.245	0.387
$\mu_t(\Delta y)$	1.57	0.867	-0.529	-0.133	0.198	3.965	0.554
$\mu_t(q^{gap})$	-2.264	-1.049	2.022	0.428	0.102	4.905	0.428
$\mu_t(\Delta q)$	2.436	1.204	-0.183	-0.066	0.139	6.535	0.258
$\mu_t(\Delta c)$	0.723	0.63	0.872	0.383	0.042	6.806	0.235
$\mu_t(\pi)$	-5.979	-1.496	3.661	0.741	0.257	6.65	0.248

Notes: See notes to Table 3.

Table 9: Pre-Crisis Two-Pass Estimation of Two-Factor Beta-Risk Model on Carry Excess Returns, 1978Q1–2008Q2

		Two-Factor Model (First HML Factor is $sk_t(UE^{gap})$)							
		2^{nd} HML							
λ_1	t-ratio	Factor	λ_2	t-ratio	γ	t-ratio	R^2	Test-stat	p-val.
0.436	**2.105**	$sk_t(\Delta UE)$	0.417	1.842*	3.659	1.248	0.903	1.252	0.940
0.476	**2.381**	$sk_t(y^{gap})$	0.309	0.726	4.341	1.268	0.929	0.938	0.967
0.484	**2.387**	$sk_t(\Delta y)$	0.126	0.329	4.134	1.269	0.929	0.894	0.971
0.488	**2.292**	$sk_t(q^{gap})$	-0.115	-0.230	3.669	0.852	0.933	1.052	0.958
0.480	**2.418**	$sk_t(\Delta q)$	-0.089	-0.121	4.299	0.954	0.928	1.178	0.947
0.663	1.950*	$sk_t(\Delta c)$	-0.081	-0.322	6.023	1.678*	0.990	0.427	0.995
0.574	**2.041**	$sk_t(\pi)$	-0.363	-0.662	10.683	1.077	0.998	0.037	1.000
0.387	**2.433**	$\sigma_t(UE^{gap})$	-0.505	-0.257	4.499	1.401	0.847	3.399	0.639
0.425	**2.426**	$\sigma_t(\Delta UE)$	-0.081	-1.295	4.278	1.290	0.822	3.007	0.699
0.300	**2.315**	$\sigma_t(y^{gap})$	0.225	1.213	3.643	1.206	0.873	3.372	0.643
0.368	**2.374**	$\sigma_t(\Delta y)$	0.348	0.616	3.444	1.140	0.759	4.475	0.483
0.539	**2.109**	$\sigma_t(q^{gap})$	-4.192	-1.452	3.525	1.104	0.896	1.340	0.931
0.394	**2.297**	$\sigma_t(\Delta q)$	-2.649	-0.642	4.768	1.358	0.791	3.363	0.644
0.236	1.839*	$\sigma_t(\Delta c)$	0.667	**2.226**	1.042	0.348	0.815	4.610	0.465
0.434	**2.431**	$\sigma_t(\pi)$	-0.917	-0.857	4.343	1.320	0.793	2.860	0.722
0.198	1.23	$\mu_t(UE^{gap})$	2.061	1.408	2.363	0.623	0.883	3.489	0.625
0.37	**2.312**	$\mu_t(\Delta UE)$	0.008	0.23	3.467	1.193	0.759	4.139	0.53
0.275	1.596	$\mu_t(y^{gap})$	0.26	0.879	3.347	1.024	0.855	3.928	0.56
0.533	1.907*	$\mu_t(\Delta y)$	-1.588	-0.958	6.495	1.219	0.836	1.985	0.851
0.41	**2.627**	$\mu_t(q^{gap})$	0.501	0.448	4.191	1.286	0.855	3.585	0.611
0.383	**2.524**	$\mu_t(\Delta q)$	-0.342	-0.25	3.94	1.354	0.762	4.281	0.51
0.443	**2.561**	$\mu_t(\Delta c)$	-0.905	-1.075	5.025	1.463	0.791	2.862	0.721
0.336	**1.981**	$\mu_t(\pi)$	-3.404	-1.109	4.137	1.009	0.838	3.126	0.68

Notes: See notes to Table 3.

Table 10: Mean Carry Excess Returns, 1978Q1–2014Q2

Six Quantiles	P_1	P_2	P_3	P_4	P_5	P_6
Mean Excess Return	2.362	2.321	1.650	4.014	3.789	8.796
Three Quantiles	P_1	P_2	P_3			
Mean Excess Return	2.342	2.832	6.293			

Table 11: Low- and High-Beta Countries, 1978Q1–2014Q2

		First Tertile			Third Tertile
Country	Beta	Excess Return	Country	Beta	Excess Return
Portugal	-30.045	0.957	Hungary	7.708	5.327
Greece	-27.133	1.298	Poland	8.890	3.916
Ireland	-13.331	-0.218	Europe	8.909	2.798
France	-13.135	4.968	Netherlands	11.697	2.739
Italy	-8.973	0.954	Chile	12.031	3.414
Philippines	-7.182	3.010	Germany	12.119	2.336
Israel	-5.711	1.580	Romania	13.557	10.941
Finland	-3.137	3.686	Indonesia	14.511	4.818
Belgium	-2.373	6.216	Austria	16.470	5.988
New Zealand	-2.073	5.581	Mexico	19.492	3.509
Canada	-1.762	1.728	Turkey	26.166	17.598
United Kingdom	-1.687	4.987	Colombia	34.303	18.066
South Korea	-0.852	3.059	Brazil	36.640	11.837

Table 12: Two-Pass Estimation of the Single-Factor Beta-Risk on Carry Excess Returns, HML $sk_t(UE^{gap})$, Developed Countries Only, 1978Q1–2014Q2

Single-Factor Model							
HML Factor	λ	t-ratio	γ	t-ratio	R^2	Test-stat	p-val.
$sk_t(UE^{gap})$	0.557	**2.193**	2.053	0.890	0.679	2.347	0.799

Table 13: Do Fama-French Factors Price Currency Excess Returns?

	A	B	C	D	E
λ_{mkt}	8.374	-1.428		-3.469	
	(1.640)	(-0.190)		(-0.591)	
λ_{SMB}	8.937	3.458	3.256		
	(**2.266**)	(0.692)	(1.216)		
λ_{value}	2.031	7.778			3.228
	(0.124)	(0.512)			(0.309)
$\lambda_{sk_t(UE^{gap})}$		0.572	0.460	0.799	0.648
		(1.127)	(**1.973**)	(**1.965**)	(1.935*)
γ	2.084	5.045	2.067	2.764	2.353
	(0.448)	(0.945)	(1.216)	(-0.591)	(0.309)
R^2	0.986	0.996	0.972	0.967	0.951
Test-stat	0.199	0.094	1.345	0.395	0.685
p-value	0.999	1.000	0.930	0.995	0.984

Notes: GMM t-ratios in parentheses. Bold indicates significance at the 5% level, '*' indicates significance at the 10% level.

Table 14: Does HML $sk_t\,(\Delta UE)$ Price Fama-French 25 Portfolios?

	A	B
λ_{mkt}	-3.265	
	(**-2.534**)	
λ_{SMB}	0.365	
	(0.854)	
λ_{value}	1.020	
	(1.777*)	
$\lambda_{sk_t(UE^{gap})}$	0.035	0.034
	(0.373)	(0.242)
γ	5.269	2.399
	(**4.308**)	(**3.909**)
R^2	0.647	0.006
Test-stat	23.132	23.626
p-value	0.512	0.483

Notes: GMM t-ratios in parentheses. Bold indicates significance at the 5% level, '*' indicates significance at the 10% level.

Table 15: SMM Estimates of the Global Risk Factor Process

	χ_g	ϕ_g	σ_g
Estimate	1.527	0.871	0.394
t-ratio	21.651	12.916	10.595

Notes: The moments used in the estimation include $E(z_{g,t})$, $E\left(z_{g,t}^2\right)$, $E(z_{g,t}z_{g,t-1})$, $E(z_{g,t}z_{g,t-2})$, $E\left(z_{g,t}^2 z_{g,t-1}^2\right)$, and $E\left(z_{g,t} z_{g,t-2}^2\right)$.

Table 16: Average SMM Parameter Estimates

	No Global Loadings		No Country-Specific Loadings		Unconstrained	
Parameter	Foreign	U.S.	Foreign	U.S.	Foreign	U.S.
ϕ	0.885	0.227	na	na	0.201	0.127
σ	0.228	0.963	na	na	0.127	0.096
θ	-1.041	0.085	-1.344	-0.833	-1.118	-0.721
ω	1.666	9.688	na	na	0.191	0.298
κ	na	na	na	na	0.163	0.151
δ	na	na	0.001	0.008	0.036	0.225
ρ	na	na	na	na	0.024	-0.039

Table 17: Excess Returns and Two-Pass Estimation of the Single-Factor Beta-Risk Model on Simulated Carry Excess Returns

A. Simulated Excess Return Summary Statistics	P_1	P_2	P_3	P_4	P_5	P_6
No Global Loadings						
Average returns by portfolios	-14.686	-12.983	-10.617	-6.891	-2.236	10.383
Sharpe	-0.725	-0.638	-0.518	-0.336	-0.106	0.529
Fama slope	0.055	0.057	0.061	0.062	0.074	-0.037
No Country-Specific Loadings						
Average returns by portfolios	-4.013	1.010	4.411	8.039	12.626	19.197
Sharpe	-1.687	0.554	1.300	1.489	1.572	1.615
Fama slope	-1.001	-0.999	-0.998	-0.999	-1.000	-0.999
Unconstrained						
Average returns by portfolios	-3.293	1.056	5.813	9.749	13.739	22.059
Sharpe	-1.402	0.705	1.917	2.018	2.001	2.411
Fama slope	-1.016	-0.974	-0.919	-0.969	-0.996	-1.097

B. Single-Factor Model Estimated on Simulated Observations	λ	t-ratio	γ	t-ratio	R^2	Test-stat	p-value
No Global Loadings	-2.743	-0.475	-2.472	-0.100	0.675	7.747	0.171
No Country-Specific Loadings	1.782	7.590	0.937	2.519	1.000	16.406	0.006
Unconstrained	2.462	9.798	2.422	4.566	0.980	15.326	0.009

Appendix A

Additional Notes on the Data

All interest rates are for 3-months maturity.

Australia: 73.1–86.1, 3 month T-bill rate. 86.2–14.2, 3-month interbank rate.

Austria: 91.2–98.4, EIBOR (Emirates Interbank Offer Rate, *Datastream*).

Belgium: 73.1–89.4, 3-month eurocurrency (Harris). 90.1–98.4, EIBOR.

Brazil: 04.1–14.2, Imputed from spot and forward rates (*Datastream*).

Canada: 73.1–96.1, 3-month eurocurrency. 96.2–14.2, 3-month T-bill rate.

Chile: 04.1–13.2, Imputed from spot and forward rates.

Colombia: 04.1–13.2, Imputed from spot and forward rates.

Czech Republic: 92.2–14.2, Interbank rate.

Denmark: 84.4–88.1, imputed from spot and forward rates. 88.2–14.2, Interbank rate.

Euro zone: 99.1–14.2, Interbank rate, Germany.

Finland: 87.1–98.4, EIBOR.

France: 73.1–96.1, 3-month eurocurrency. 96.2–98.4, EIBOR.

Germany: 73.1–96.1, 3-month eurocurrency. 96.2–98.4, EIBOR.

Greece: 94.2–98.4. Interbank.

Hungary: 95.3–14.2: Interbank.

Iceland: 95.3–00.1, Interbank mid-rate. 00.2–14.2, Reykjavik interbank offer rate.

India: 97.4–98.3, Imputed from spot and forward rates. 98.4–14.2 Interbank.

Indonesia: 96.1–14.2, Interbank rate.

Ireland: 84.1–98.4. Interbank.

Israel: 94.4–99.3, T-bill. 99.4–14.2, Interbank.

Italy: 73.1–96.1, 3-month eurocurrency. 96.2–98.4, EIBOR.

Japan: 73.1–96.1, 3-month eurocurrency. 96.2–14.2, Interbank.

Malaysia: 93.3–14.2, Interbank.

Mexico: 78.1–14.2, T-bill (FRED).

Netherlands: 73.1–96.1, 3-month eurocurrency. 96.2–98.4, EIBOR.

New Zealand: 74.1–13.4, Interbank (FRED).

Norway: 86.1–14.2. Interbank.

Philippines: 87.1–14.2 T-bill.

Poland: 94.4–14.2 Interbank.

Portugal: 96.4–98.4, Imputed from spot and forward.

Romania: 95.3–14.2. Interbank.

Singapore: 84.4–87.2, Imputed from spot and forward rates. 87.3–13.4, Interbank.

South Africa: 73.1–14.3. T-bill.

South Korea: 92.1–14.2. Interbank.

Spain: 88.3–98.4, Interbank.

Sweden: 84.4–86.4, Imputed from spot and forward rates. 87.1–14.3, Interbank.

Switzerland: 73.1–96.1, 3-month eurocurrency. 96.2–14.2, Interbank.

Taiwan: 82.2–14.2, Money market rates.

Thailand: 95.1–96.3, imputed from spot and forward rates. 96.5–14.2, Interbank.

Turkey: 96.4–06.4, imputed from spot and forward rates. 07.1–14.2, Interbank.

United Kingdom: 73.1–96.1, 3-month eurocurrency. 96.2–98.4, U.K. Interbank.

Appendix B

Two-Pass Regression Procedure and GMM Standard Errors

We have k factors, T time-series observations and n excess returns (assets). Vectors are underlined. Matrices are bolded. Scalars have no special designation. The objective is to estimate the k-factor 'beta-risk' model

$$E\left(r_{i,t}^{e}\right) = \underline{\beta}_i'\underline{\lambda} + \alpha_i, \tag{13}$$

where $\underline{\beta}_i$ is a k-dimensional vector of the factor betas for excess return i and $\underline{\lambda}$ is the $k-$dimensional vector of factor risk premia. The expectation is taken over t. The beta-risk model's answer to the question as to why average returns vary across assets is that returns with high betas (covariance with a factor) pay a high-risk premium ($\underline{\lambda}$). The cross-sectional test can be implemented with a two-pass procedure. Let $\underline{f}_t$ be the k-dimensional vector of the macro factors. In the first pass for each excess return $i = 1, ..., n$, estimate the factor betas in the time-series regression,

$$r_{i,t}^{e} = a_i + \underbrace{(\beta_{1,i}, ..., \beta_{k,i})}_{\underline{\beta}_i'} \begin{pmatrix} f_{1,t} \\ \vdots \\ f_{k,t} \end{pmatrix} + \epsilon_{i,t} = \underline{\tilde{\beta}}_i'\underline{F}_t + \epsilon_{i,t},$$

where

$$\underline{F}_t = \begin{pmatrix} 1 \\ \underline{f}_t \end{pmatrix}, \ \underline{f}_t = \begin{pmatrix} f_{1,t} \\ \vdots \\ f_{k,t} \end{pmatrix}, \ \underset{(k+1)\times 1}{\underline{\tilde{\beta}}_i} = \begin{pmatrix} a_i \\ \underline{\beta}_i \end{pmatrix}, \ \underset{(k\times 1)}{\underline{\beta}_i} = \begin{pmatrix} \beta_{1,i} \\ \vdots \\ \beta_{k,i} \end{pmatrix}.$$

In the second pass, we can run the cross-sectional regression of average returns $\bar{r}_i^e = (1/T)\sum_{t=1}^{T} r_{i,t}^e$, using the betas as data, to estimate the factor risk premia, $\underline{\lambda}$. If the excess return's covariance with the factor is systematic and undiversifiable, that covariance risk should be 'priced' into the return. The factor risk premium should not be zero. The second-pass regression run with a constant is

$$\bar{r}_i^e = \gamma + \underbrace{(\lambda_1, ..., \lambda_k)}_{\underline{\lambda}} \begin{pmatrix} \beta_{1,i} \\ \vdots \\ \beta_{k,i} \end{pmatrix} + \alpha_i = \gamma + \underline{\lambda}'\underline{\beta}_i + \alpha_i.$$

The α_i are the pricing errors. When the cross-sectional regression is run without a constant, set $\gamma = 0$.

$$\overline{r}_i^e = \gamma + \underline{\beta}_i' \underline{\lambda} + \alpha_i.$$

OLS standard errors give asymptotically incorrect inference because the βs are not data but are generated regressors. Cochrane (2005) describes a procedure to obtain GMM standard errors that delivers an asymptotically valid inference that is robust to the generated regressors problem and robust to heteroskedasticity and autocorrelation in the errors. Cochrane's strategy is to use the standard errors from a GMM estimation problem that exactly reproduces the two-stage regression point estimates. We will need the following notation:

$$\underset{(k\times k)}{\mathbf{\Sigma}_f} = E\left(\underline{f}_t - \underline{\mu}_f\right)\left(\underline{f}_t - \underline{\mu}_f\right)'$$

$$\underline{\epsilon}_t = (\epsilon_{1,t}, \epsilon_{2,t}, ..., \epsilon_{n,t})'$$

$$\underset{n\times n}{\mathbf{\Sigma}} = E\left(\underline{\epsilon}_t \underline{\epsilon}_t'\right)$$

$$\underset{n\times k}{\mathbf{B}} = \begin{pmatrix} \underline{\beta}_1' \\ \vdots \\ \underline{\beta}_n' \end{pmatrix} = \begin{pmatrix} \beta_{1,1} & \cdots & \beta_{k,1} \\ \vdots & & \vdots \\ \beta_{1,n} & \cdots & \beta_{k,n} \end{pmatrix}$$

$$\underset{k\times n}{\mathbf{A}} = \underset{k\times k}{(\mathbf{B}'\mathbf{B})^{-1}} \underset{k\times n}{\mathbf{B}'}$$

$$\underset{n\times n}{\mathbf{M}_\beta} = \mathbf{I}_n - \underset{n\times k}{\mathbf{B}} \underset{k\times k}{(\mathbf{B}'\mathbf{B})^{-1}} \underset{k\times n}{\mathbf{B}}'$$

$$\underset{n\times(k+1)}{\mathbf{X}} = (\underline{\iota}_n \ \mathbf{B}'), \text{ where } \underline{\iota}_n = \begin{pmatrix} 1 \\ \vdots \\ 1 \end{pmatrix} \leftarrow n'\text{th row}$$

$$\underset{(k+1)\times n}{\mathbf{C}} = \underset{(k+1)\times(k+1)}{(\mathbf{X}'\mathbf{X})}^{-1} \underset{(k+1)\times n}{\mathbf{X}'}$$

$$\underset{n\times n}{\mathbf{M}_X} = \mathbf{I}_n - \underset{n\times(k+1)}{\mathbf{X}} \underset{(k+1)\times(k+1)}{(\mathbf{X}'\mathbf{X})^{-1}} \underset{(k+1)\times n}{\mathbf{X}'}$$

$$\underset{(k+1)\times(k+1)}{\tilde{\mathbf{\Sigma}}_f} = \begin{pmatrix} \underset{\text{scalar}}{0} & \underset{1\times k}{\underline{0}} \\ \underset{k\times 1}{\underline{0}} & \underset{k\times k}{\mathbf{\Sigma}_f} \end{pmatrix}$$

Estimation without the constant. When estimating without the constant in the second-pass regression, the parameter vector is

$$\underset{[k(n+1)+k]\times 1}{\underline{\theta}} = \begin{pmatrix} \underline{\tilde{\beta}}_1 \\ \vdots \\ \underline{\tilde{\beta}}_n \\ \underline{\lambda} \end{pmatrix} = \begin{pmatrix} a_1 \\ \underline{\beta}_1 \\ \vdots \\ a_n \\ \underline{\beta}_n \\ \underline{\lambda} \end{pmatrix}.$$

Let the second moment matrix of the factors be

$$\underset{(k+1)\times(k+1)}{\mathbf{M}_F} = \frac{1}{T}\sum_{t=1}^{T} \underline{F}_t \underline{F}'_t.$$

The moment conditions are built off of the error vector,

$$\underset{n(k+2)\times 1}{\underline{u}_t\left(\theta\right)} = \begin{pmatrix} \underline{F}_t\left(r^e_{1,t} - \underline{F}'_t\underline{\tilde{\beta}}_1\right) \\ \vdots \\ \underline{F}_t\left(r^e_{n,t} - \underline{F}'_t\underline{\tilde{\beta}}_n\right) \\ r^e_{1,t} - \underline{\beta}'_1\underline{\lambda} \\ \vdots \\ r^e_{n,t} - \underline{\beta}'_n\underline{\lambda} \end{pmatrix} = \begin{array}{cc} \begin{pmatrix} \underline{F}_t\left(r^e_{1,t} - \underline{F}'_t\underline{\tilde{\beta}}_1\right) \\ \vdots \\ \underline{F}_t\left(r^e_{n,t} - \underline{F}'_t\underline{\tilde{\beta}}_n\right) \\ \underline{R}^e_t - \mathbf{B}\underline{\lambda} \end{pmatrix} & \begin{array}{r} \\ \\ \leftarrow \text{row } n\left(k+1\right) \\ \leftarrow \left(n\times 1\right) \end{array} \end{array}$$

where

$$\underline{r}^e_t = \begin{pmatrix} r^e_{1,t} \\ \vdots \\ r^e_{n,t} \end{pmatrix}.$$

Let

$$\underline{g}_T\left(\underline{\theta}\right) = \frac{1}{T}\sum_{t=1}^{T}\underline{u}_t\left(\theta\right)$$

$$\underset{[n(k+1)]\times[n(k+1)+k]}{\mathbf{d}_T} = \frac{\partial \mathbf{g}_T\left(\underline{\theta}\right)}{\partial\underline{\theta}'} = \begin{pmatrix} \underset{[n(k+1)]\times[n(k+1)]}{-\mathbf{I}_n\otimes\mathbf{M}_F} & \underset{[n(k+1)]\times k}{\mathbf{0}} \\ \underset{n\times[n(k+1)]}{-\mathbf{I}_n\otimes\begin{pmatrix}\underset{\text{scalar}}{0} & \underline{\lambda}'\end{pmatrix}} & \underset{n\times k}{-\mathbf{B}} \end{pmatrix}.$$

To replicate the estimates in the two-pass procedure, we need[13]

$$\underset{[n(k+1)+k]\times[n(k+2)]}{\mathbf{a}_T} = \begin{pmatrix} \mathbf{I}_{n(k+1)} & \underset{n(k+1)\times n}{\mathbf{0}} \\ \underset{k\times n(k+1)}{\mathbf{0}} & \underset{k\times n}{\mathbf{B}'} \end{pmatrix}, \tag{14}$$

not $\mathbf{d}_T\mathbf{S}_T^{-1}$. The coefficient covariance matrix we want is

$$\mathbf{V}_\theta = \frac{1}{T}\left(\mathbf{a}_T\mathbf{d}_T\right)^{-1}\left(\mathbf{a}_T\mathbf{S}_T\mathbf{a}_T'\right)\left[\left(\mathbf{a}_T\mathbf{d}_T\right)^{-1}\right]'. \tag{15}$$

To test if the pricing errors are zero, use the covariance matrix of the moment conditions,

$$\mathbf{V}_g = \frac{1}{T}\left(\mathbf{I}_{(n(k+1))} - \mathbf{d}_T\left(\mathbf{a}_T\mathbf{d}_T\right)^{-1}\mathbf{a}_T\right)\mathbf{S}_T\left(\mathbf{I}_{(n(k+2))} - \mathbf{d}_T\left(\mathbf{a}_T\mathbf{d}_T\right)^{-1}\mathbf{a}_T\right). \tag{16}$$

We want to get $\mathbf{V}_\theta$ and $\mathbf{V}_g$ by plugging in.

GMM standard errors when estimating with a constant. The cross-sectional regression is now

$$\frac{1}{T}\sum_{t=1}^{T} r_{i,t}^e = \gamma + \underline{\beta}_i'\underline{\lambda} + \alpha_i,$$

where γ is the constant. We have to add γ to the coefficient vector θ. Place it according to

$$\underset{(n+1)(k+1)\times 1}{\underline{\theta}} = \begin{pmatrix} a_1 \\ \underline{\beta}_1 \\ \vdots \\ a_n \\ \underline{\beta}_n \\ \gamma \\ \underline{\lambda} \end{pmatrix} = \begin{pmatrix} \underline{\tilde{\beta}}_1 \\ \vdots \\ \underline{\tilde{\beta}}_n \\ \gamma \\ \underline{\lambda} \end{pmatrix}.$$

[13] In the usual GMM problem, we minimize

$$\underline{g}_T\left(\underline{\theta}\right)'\mathbf{S}_T^{-1}\underline{g}_T\left(\underline{\theta}\right),$$

where

$$\mathbf{S}_T \overset{a.s.}{\to} \mathbf{S} = \mathbf{E}\left(\sum_{j=-\infty}^{\infty}\underline{u}_t\left(\underline{\theta}\right)\underline{u}_{t-j}\left(\theta\right)'\right).$$

We do Newey-West on $\underline{u}_t\left(\theta\right)$ to get $\mathbf{S}_T$. We will want to plug in our estimated λ and βs into $\mathbf{d}_T$. This problem chooses $\underline{\theta}$ to set

$$\mathbf{d}_T\mathbf{S}_T^{-1}\underline{g}_T\left(\underline{\theta}\right) = \underline{0}$$

and can be recast as having a weighting matrix on the moment conditions

$$\mathbf{a}_T\underline{g}_T\left(\underline{\theta}\right) = 0$$

where

$$\mathbf{a}_T = \mathbf{d}_T\mathbf{S}_T^{-1}$$

The covariance matrix of $\underline{\theta}$ for this problem is,

$$\mathbf{V}_\theta = \frac{1}{T}\left(\mathbf{d}_T\mathbf{S}_T\mathbf{d}_T\right)^{-1}$$

but this is not the covariance matrix for the two-pass estimation problem. The reason is that the last set of n moment conditions in $\underline{\mathbf{g}}_T\left(\underline{\theta}\right)$ isn't the cross-sectional regression estimated by least squares (which is $\mathbf{B}'\left(\frac{1}{T}\sum_{t=1}^{T}\underline{R}_t^e - \mathbf{B}\underline{\lambda}\right)$).

Define

$$\mathbf{X} = \begin{pmatrix} \underset{n\times 1}{\iota} & \underset{n\times k}{\mathbf{B}} \end{pmatrix}.$$

The error vector that defines the model is

$$\underline{u}_t\left(\underline{\theta}\right) = \begin{pmatrix} \underline{F}_t\left(r^e_{1,t} - \underline{F}'_t\underline{\tilde{\beta}}_1\right) \\ \vdots \\ \underline{F}_t\left(r^e_{n,t} - \underline{F}'_t\underline{\tilde{\beta}}_n\right) \\ r^e_{1,t} - \gamma - \underline{\beta}'_1\underline{\lambda} \\ \vdots \\ r^e_{n,t} - \gamma - \underline{\beta}'_n\underline{\lambda} \end{pmatrix} = \begin{pmatrix} \underline{F}_t\left(r^e_{1,t} - \underline{F}'_t\underline{\tilde{\beta}}_1\right) \\ \vdots \\ \underline{F}_t\left(r^e_{n,t} - \underline{F}'_t\underline{\tilde{\beta}}_n\right) \\ \underline{R}^e_t - \mathbf{X}\begin{pmatrix} \gamma \\ \underline{\lambda} \end{pmatrix} \end{pmatrix}.$$

Do Newey-West on $\underline{u}_t\left(\theta\right)$ to get $\mathbf{S}_T$. Use

$$\underset{[(n+1)(k+1)]\times[n(k+2)]}{\mathbf{a}_T} = \begin{pmatrix} \mathbf{I}_{n(k+1)} & \underset{[n(k+1)]\times n}{\mathbf{0}} \\ \underset{(k+1)\times[n(k+1)]}{\mathbf{0}} & \underset{(k+1)\times n}{\mathbf{X}'} \end{pmatrix}$$

$$\underset{[n(k+1)]\times[(k+1)(n+1)]}{\mathbf{d}_T} = \frac{\partial \mathbf{g}_T\left(\underline{\theta}\right)}{\partial\underline{\theta}'} = \begin{pmatrix} \underset{[n(k+1)]\times[n(k+1)]}{-\mathbf{I}_n \otimes \mathbf{M}_F} & \underset{[n(k+1)]\times(k+1)}{\mathbf{0}} \\ \underset{n\times[n(k+1)]}{-\mathbf{I}_n \otimes \begin{pmatrix} \underset{\text{scalar}}{0} & \underline{\lambda}' \end{pmatrix}} & \underset{n\times(k+1)}{-\mathbf{X}} \end{pmatrix}$$

to plug into (15) and (16).

We do not use GMM to estimate the model. We use the two-step procedure to get the point estimates for the betas and lambdas and plug those estimates into the GMM formulae to get standard errors.

Appendix C

Alternative Window Sizes

This appendix reports estimations of the beta model when the relevant moments are computed with windows of 16 and 24 quarters.

Table 18: Two-Pass Estimation of the Single-Factor Beta-Risk Model on Carry Excess Returns, 1977Q1–2014Q2

	Single-Factor Model						
HML Factor	λ	t-ratio	γ	t-ratio	R^2	Test-stat	p-val.
$sk_t(UE^{gap})$	0.530	**3.561**	3.186	1.295	0.969	2.427	0.787
$sk_t(\Delta UE)$	0.816	**2.264**	3.215	0.786	0.995	0.095	1.000
$sk_t(\Delta y)$	0.427	**2.580**	5.480	1.930*	0.677	5.141	0.399
$sk_t(y^{gap})$	0.382	**2.166**	7.108	**2.639**	0.172	7.179	0.208
$sk_t(\Delta c)$	0.592	**2.317**	7.672	**1.990**	0.630	2.513	0.775
$sk_t(\pi)$	-0.172	-1.314	5.188	**2.376**	0.027	9.919	0.078
$sk_t(q^{gap})$	0.876	1.797*	6.720	1.679*	0.250	3.823	0.575
$sk_t(\Delta q)$	0.769	**2.332**	7.185	**2.640**	0.773	4.416	0.491
$\sigma_t(UE^{gap})$	4.273	**2.373**	5.548	1.599	0.603	6.371	0.272
$\sigma_t(\Delta UE)$	0.013	0.470	3.857	**2.553**	0.002	11.348	0.045
$\sigma_t(\Delta y)$	2.093	**2.737**	1.216	0.341	0.967	1.117	0.953
$\sigma_t(y^{gap})$	1.058	**2.659**	1.514	0.389	0.927	2.330	0.802
$\sigma_t(\Delta c)$	2.079	**2.386**	-1.326	-0.282	0.890	1.730	0.885
$\sigma_t(\pi)$	4.615	**2.464**	2.265	0.740	0.846	2.787	0.733
$\sigma_t(q^{gap})$	-6.824	-1.203	5.959	1.031	0.320	1.951	0.856
$\sigma_t(\Delta q)$	9.611	**2.933**	3.728	1.340	0.923	3.068	0.690
$\mu_t(UE^{gap})$	4.242	**2.718**	4.632	1.065	0.818	3.330	0.649
$\mu_t(\Delta UE)$	0.121	**2.634**	3.888	0.826	0.812	2.755	0.738
$\mu_t(\Delta y)$	-1.999	**-2.465**	2.496	0.718	0.970	0.635	0.986
$\mu_t(y^{gap})$	0.981	**2.788**	3.840	1.011	0.796	3.096	0.685
$\mu_t(\Delta c)$	-1.453	**-3.327**	1.327	0.497	0.986	0.482	0.993
$\mu_t(\pi)$	9.620	**2.865**	3.016	1.246	0.913	2.998	0.700
$\mu_t(q^{gap})$	2.099	**2.142**	3.048	0.709	0.532	2.643	0.755
$\mu_t(\Delta q)$	4.726	1.842*	-0.373	-0.074	0.674	2.027	0.845

Notes: The raw data are quarterly (1973Q1 to 2014Q2) and, when available, are end-of-quarter and point-sampled. 16 quarters start-up to compute initial HML factors. Model estimated on returns from 1977Q1 to 2014Q2. Δy, y^{gap}, Δc, ΔUE, UE^{gap}, π, q^{gap}, and Δq represent the GDP growth rate, output gap, consumption growth rate, change in the unemployment rate, unemployment gap, inflation rate, real exchange rate gap, and real exchange rate depreciation, respectively. For each country (41 countries plus the euro area) and each macroeconomic variable (x), we compute the 'conditional' mean ($\mu_t(x)$), volatility ($\sigma_t(x)$) and skewness ($sk_t(x)$) using a 16-quarter window. To form the portfolio returns, we sort by the nominal interest rate (carry) for each country from low to high. The rank ordering is divided into six portfolios, into which the currency returns are assigned. P_6 is the portfolio of returns associated with the highest nominal interest rate countries and P_1 is the portfolio of returns associated with the lowest nominal interest rate countries. This table reports the two-pass procedure estimation results from a one-factor model. In the first pass, we run $N = 6$ individual time-series regressions of the excess returns on the K factors to estimate the factor 'betas,' $r^e_{i,t} = a_i + \sum_{k=1}^{K} \beta_{i,k} f^{HML}_{k,t} + \epsilon_{i,t}$, where $r^e_{i,t}$ is the excess return, $\beta_{i,k}$ is the factor beta and $f^{HML}_{k,t}$ is the high-minus-low (HML) macro risk factor. The factors considered include the high-minus-low (HML) values of the conditional mean, variance, and skewness of Δy, y^{gap}, Δc, ΔUE, UE^{gap}, π, q^{gap}, and Δq. Each HML value is equal to the average in the highest quartile minus the average in the lowest quartile. In the second pass, we run a single cross-sectional regression of the (time-series) mean excess returns on the estimated betas, $\bar{r}^e_i = \gamma + \sum_{k=1}^{K} \lambda_k \beta_{i,k} + \alpha_i$, where $\bar{r}^e_i$ is the average excess return, γ is the intercept, λ_k is the risk premia, and α_i is the pricing error. The table reports the price of risk (λ) and its associated t-ratio (using GMM standard errors), the estimated intercept (γ) and its associate t-ratio, R^2 and the Wald test on the pricing errors (Test-stat) and its associated p-value (p-val.). Bold indicates significance at the 5% level. '*' indicates significance at the 10% level.

Table 19: Two-Pass Estimation of the Two-Factor Beta-Risk Model on Carry Excess Returns, 1977Q1–2014Q2

Two-Factor Model (First HML Factor is $sk_t(UE^{gap})$)									
λ_1	t-ratio	2^{nd} HML Factor	λ_2	t-ratio	γ	t-ratio	R^2	Test-stat	p-val.
0.406	**1.988**	$sk_t(\Delta UE)$	0.752	1.664*	3.209	0.853	0.996	0.097	1.000
0.524	**2.982**	$sk_t(\Delta y)$	0.161	1.262	3.253	1.350	0.970	2.346	0.800
0.519	**3.750**	$sk_t(y^{gap})$	0.156	1.065	2.712	0.969	0.972	2.301	0.806
0.490	**2.836**	$sk_t(\Delta c)$	0.127	0.751	3.837	1.529	0.976	1.409	0.923
0.546	**3.685**	$sk_t(\pi)$	-0.156	-1.217	4.269	1.460	0.984	0.845	0.974
0.456	**2.884**	$sk_t(q^{gap})$	0.112	0.327	3.731	1.404	0.974	1.675	0.892
0.506	**2.933**	$sk_t(\Delta q)$	0.312	1.382	3.638	1.445	0.972	2.231	0.816
0.413	**3.005**	$\sigma_t(UE^{gap})$	0.354	0.342	3.527	1.560	0.976	2.863	0.721
0.481	**3.434**	$\sigma_t(\Delta UE)$	-0.013	-0.300	3.198	1.305	0.966	3.038	0.694
0.351	1.708*	$\sigma_t(\Delta y)$	1.246	1.337	2.208	0.780	0.984	1.304	0.935
0.395	**2.812**	$\sigma_t(y^{gap})$	0.548	1.811*	2.494	0.912	0.995	0.528	0.991
0.430	**2.917**	$\sigma_t(\Delta c)$	0.784	1.032	1.827	0.603	0.983	0.852	0.974
0.480	**3.419**	$\sigma_t(\pi)$	2.158	1.818*	2.985	1.284	0.969	2.878	0.719
0.487	**3.520**	$\sigma_t(q^{gap})$	0.011	0.006	2.965	1.065	0.967	3.017	0.697
0.409	**2.765**	$\sigma_t(\Delta q)$	4.963	1.454	3.294	1.480	0.974	2.918	0.713
0.445	**3.325**	$\mu_t(UE^{gap})$	1.315	1.209	3.357	1.365	0.971	2.565	0.767
0.392	**2.984**	$\mu_t(\Delta UE)$	0.039	1.497	3.311	1.203	0.986	0.964	0.965
0.422	**1.966**	$\mu_t(\Delta y)$	-1.366	-1.579	2.826	1.038	0.977	1.105	0.954
0.489	**2.777**	$\mu_t(y^{gap})$	0.312	1.051	3.268	1.379	0.974	2.276	0.810
0.316	0.945	$\mu_t(\Delta c)$	-1.388	-1.285	1.468	0.376	0.987	0.509	0.992
0.467	**2.953**	$\mu_t(\pi)$	5.232	1.791*	3.102	1.394	0.968	2.924	0.712
0.499	**3.671**	$\mu_t(q^{gap})$	-0.159	-0.323	3.240	1.171	0.990	0.432	0.994
0.484	**3.502**	$\mu_t(\Delta q)$	0.948	0.733	2.915	1.053	0.967	3.227	0.665

Notes: The raw data are quarterly (1973Q1 to 2014Q2) and, when available, are end-of-quarter and point-sampled. 16 quarters start-up to compute initial HML factors. Model estimated on returns from 1977Q1 to 2014Q2. Δy, y^{gap}, Δc, ΔUE, UE^{gap}, π, q^{gap}, and Δq represent the GDP growth rate, output gap, consumption growth rate, change in the unemployment rate, unemployment gap, inflation rate, real exchange rate gap, and real exchange rate depreciation, respectively. For each country (41 countries plus the euro area) and each macroeconomic variable (x), we compute the 'conditional' mean ($\mu_t(x)$), volatility ($\sigma_t(x)$) and skewness ($sk_t(x)$) using a 16-quarter window. To form the portfolio returns, we sort by the nominal interest rate (carry) for each country from low to high. The rank ordering is divided into six portfolios, into which the currency returns are assigned. P_6 is the portfolio of returns associated with the highest nominal interest rate countries and P_1 is the portfolio of returns associated with the lowest nominal interest rate countries. This table reports the two-pass procedure estimation results from a two-factor model where $sk_t(UE^{gap})$ is the maintained first factor. In the first pass, we run $N = 6$ individual time-series regressions of the excess returns on the K factors to estimate the factor 'betas,' $r^e_{i,t} = a_i + \sum_{k=1}^{K} \beta_{i,k} f^{HML}_{k,t} + \epsilon_{i,t}$, where $r^e_{i,t}$ is the excess return, $\beta_{i,k}$ is the factor beta and $f^{HML}_{k,t}$ is the high-minus-low (HML) macro risk factor. The factors considered include the high-minus-low (HML) values of the conditional mean, variance, and skewness of Δy, y^{gap}, Δc, ΔUE, UE^{gap}, π, q^{gap}, and Δq. Each HML value is equal to the average in the highest quartile minus the average in the lowest quartile. In the second pass, we run a single cross-sectional regression of the (time-series) mean excess returns on the estimated betas, $\bar{r}^e_i = \gamma + \sum_{k=1}^{K} \lambda_k \beta_{i,k} + \alpha_i$, where $\bar{r}^e_i$ is the average excess return, γ is the intercept, λ_k is the risk premia, and α_i is the pricing error. The table reports the price of risk (λ) and its associated t-ratio (using GMM standard errors), the estimated intercept (γ) and its associate t-ratio, R^2 and the Wald test on the pricing errors (Test-stat) and its associated p-value (p-val.). Bold indicates significance at the 5% level. '*' indicates significance at the 10% level.

Table 20: Two-Pass Estimation of the Single-Factor Beta-Risk Model on Carry Excess Returns, 1979Q1–2014Q2

	Single-Factor Model						
HML Factor	λ	t-ratio	γ	t-ratio	R^2	Test-stat	p-val.
$sk_t(UE^{gap})$	0.530	**3.561**	3.186	1.295	0.969	2.427	0.787
$sk_t(\Delta UE)$	0.816	**2.264**	3.215	0.786	0.995	0.095	1.000
$sk_t(\Delta y)$	0.427	**2.580**	5.480	1.930*	0.677	5.141	0.399
$sk_t(y^{gap})$	0.382	**2.166**	7.108	**2.639**	0.172	7.179	0.208
$sk_t(\Delta c)$	0.592	**2.317**	7.672	1.990*	0.630	2.513	0.775
$sk_t(\pi)$	-0.172	-1.314	5.188	**2.376**	0.027	9.919	0.078
$sk_t(q^{gap})$	0.876	1.797*	6.720	1.679*	0.250	3.823	0.575
$sk_t(\Delta q)$	0.769	**2.332**	7.185	**2.640**	0.773	4.416	0.491
$\sigma_t(UE^{gap})$	4.273	**2.373**	5.548	1.599	0.603	6.371	0.272
$\sigma_t(\Delta UE)$	0.013	0.470	3.857	**2.553**	0.002	11.348	0.045
$\sigma_t(\Delta y)$	2.093	**2.737**	1.216	0.341	0.967	1.117	0.953
$\sigma_t(y^{gap})$	1.058	**2.659**	1.514	0.389	0.927	2.330	0.802
$\sigma_t(\Delta c)$	2.079	**2.386**	-1.326	-0.282	0.890	1.730	0.885
$\sigma_t(\pi)$	4.615	**2.464**	2.265	0.740	0.846	2.787	0.733
$\sigma_t(q^{gap})$	-6.824	-1.203	5.959	1.031	0.320	1.951	0.856
$\sigma_t(\Delta q)$	9.611	**2.933**	3.728	1.340	0.923	3.068	0.690
$\mu_t(UE^{gap})$	4.242	**2.718**	4.632	1.065	0.818	3.330	0.649
$\mu_t(\Delta UE)$	0.121	**2.634**	3.888	0.826	0.812	2.755	0.738
$\mu_t(\Delta y)$	-1.999	**-2.465**	2.496	0.718	0.970	0.635	0.986
$\mu_t(y^{gap})$	0.981	**2.788**	3.840	1.011	0.796	3.096	0.685
$\mu_t(\Delta c)$	-1.453	**-3.327**	1.327	0.497	0.986	0.482	0.993
$\mu_t(\pi)$	9.620	**2.865**	3.016	1.246	0.913	2.998	0.700
$\mu_t(q^{gap})$	2.099	**2.142**	3.048	0.709	0.532	2.643	0.755
$\mu_t(\Delta q)$	4.726	1.842*	-0.373	-0.074	0.674	2.027	0.845

Notes: The raw data are quarterly (1973Q1 to 2014Q2) and, when available, are end-of-quarter and point-sampled. 24 quarters start-up to compute initial HML factors. Model estimated on returns from 1979Q1 to 2014Q2. Δy, y^{gap}, Δc, ΔUE, UE^{gap}, π, q^{gap}, and Δq represent the GDP growth rate, output gap, consumption growth rate, change in the unemployment rate, unemployment gap, inflation rate, real exchange rate gap, and real exchange rate depreciation, respectively. For each country (41 countries plus the euro area) and each macroeconomic variable (x), we compute the 'conditional' mean ($\mu_t(x)$), volatility ($\sigma_t(x)$) and skewness ($sk_t(x)$) using a 24-quarter window. To form the portfolio returns, we sort by the nominal interest rate (carry) for each country from low to high. The rank ordering is divided into six portfolios, into which the currency returns are assigned. P_6 is the portfolio of returns associated with the highest nominal interest rate countries and P_1 is the portfolio of returns associated with the lowest nominal interest rate countries. This table reports the two-pass procedure estimation results from a one-factor model. In the first pass, we run $N = 6$ individual time-series regressions of the excess returns on the K factors to estimate the factor 'betas,' $r^e_{i,t} = a_i + \sum_{k=1}^{K} \beta_{i,k} f^{HML}_{k,t} + \epsilon_{i,t}$, where $r^e_{i,t}$ is the excess return, $\beta_{i,k}$ is the factor beta and $f^{HML}_{k,t}$ is the high-minus-low (HML) macro risk factor. The factors considered include the high-minus-low (HML) values of the conditional mean, variance, and skewness of Δy, y^{gap}, Δc, ΔUE, UE^{gap}, π, q^{gap}, and Δq. Each HML value is equal to the average in the highest quartile minus the average in the lowest quartile. In the second pass, we run a single cross-sectional regression of the (time-series) mean excess returns on the estimated betas, $\bar{r}^e_i = \gamma + \sum_{k=1}^{K} \lambda_k \beta_{i,k} + \alpha_i$, where $\bar{r}^e_i$ is the average excess return, γ is the intercept, λ_k is the risk premia, and α_i is the pricing error. The table reports the price of risk (λ) and its associated t-ratio (using GMM standard errors), the estimated intercept (γ) and its associate t-ratio, R^2 and the Wald test on the pricing errors (Test-stat) and its associated p-value (p-val.). Bold indicates significance at the 5% level. '*' indicates significance at the 10% level.

Table 21: Two-Pass Estimation of the Two-Factor Beta-Risk Model on Carry Excess Returns, 1979Q1–2014Q2

Two-Factor Model (First HML Factor is $sk_t(UE^{gap})$)									
λ_1	t-ratio	2^{nd} HML Factor	λ_2	t-ratio	γ	t-ratio	R^2	Test-stat	p-val.
0.406	**1.988**	$sk_t(\Delta UE)$	0.752	1.664*	3.209	0.853	0.996	0.097	1.000
0.524	**2.982**	$sk_t(\Delta y)$	0.161	1.262	3.253	1.350	0.970	2.346	0.800
0.519	**3.750**	$sk_t(y^{gap})$	0.156	1.065	2.712	0.969	0.972	2.301	0.806
0.490	**2.836**	$sk_t(\Delta c)$	0.127	0.751	3.837	1.529	0.976	1.409	0.923
0.546	**3.685**	$sk_t(\pi)$	-0.156	-1.217	4.269	1.460	0.984	0.845	0.974
0.456	**2.884**	$sk_t(q^{gap})$	0.112	0.327	3.731	1.404	0.974	1.675	0.892
0.506	**2.933**	$sk_t(\Delta q)$	0.312	1.382	3.638	1.445	0.972	2.231	0.816
0.413	**3.005**	$\sigma_t(UE^{gap})$	0.354	0.342	3.527	1.560	0.976	2.863	0.721
0.481	**3.434**	$\sigma_t(\Delta UE)$	-0.013	-0.300	3.198	1.305	0.966	3.038	0.694
0.351	1.708*	$\sigma_t(\Delta y)$	1.246	1.337	2.208	0.780	0.984	1.304	0.935
0.395	**2.812**	$\sigma_t(y^{gap})$	0.548	1.811*	2.494	0.912	0.995	0.528	0.991
0.430	**2.917**	$\sigma_t(\Delta c)$	0.784	1.032	1.827	0.603	0.983	0.852	0.974
0.480	**3.419**	$\sigma_t(\pi)$	2.158	1.818*	2.985	1.284	0.969	2.878	0.719
0.487	**3.520**	$\sigma_t(q^{gap})$	0.011	0.006	2.965	1.065	0.967	3.017	0.697
0.409	**2.765**	$\sigma_t(\Delta q)$	4.963	1.454	3.294	1.480	0.974	2.918	0.713
0.445	**3.325**	$\mu_t(UE^{gap})$	1.315	1.209	3.357	1.365	0.971	2.565	0.767
0.392	**2.984**	$\mu_t(\Delta UE)$	0.039	1.497	3.311	1.203	0.986	0.964	0.965
0.422	**1.966**	$\mu_t(\Delta y)$	-1.366	-1.579	2.826	1.038	0.977	1.105	0.954
0.489	**2.777**	$\mu_t(y^{gap})$	0.312	1.051	3.268	1.379	0.974	2.276	0.810
0.316	0.945	$\mu_t(\Delta c)$	-1.388	-1.285	1.468	0.376	0.987	0.509	0.992
0.467	**2.953**	$\mu_t(\pi)$	5.232	1.791	3.102	1.394	0.968	2.924	0.712
0.499	**3.671**	$\mu_t(q^{gap})$	-0.159	-0.323	3.240	1.171	0.990	0.432	0.994
0.484	**3.502**	$\mu_t(\Delta q)$	0.948	0.733	2.915	1.053	0.967	3.227	0.665

Notes: The raw data are quarterly (1973Q1 to 2014Q2) and, when available, are end-of-quarter and point-sampled. 24 quarters start-up to compute initial HML factors. Model estimated on returns from 1979Q1 to 2014Q2. Δy, y^{gap}, Δc, ΔUE, UE^{gap}, π, q^{gap}, and Δq represent the GDP growth rate, output gap, consumption growth rate, change in the unemployment rate, unemployment gap, inflation rate, real exchange rate gap, and real exchange rate depreciation, respectively. For each country (41 countries plus the euro area) and each macroeconomic variable (x), we compute the 'conditional' mean ($\mu_t(x)$), volatility ($\sigma_t(x)$) and skewness ($sk_t(x)$) using a 24-quarter window. To form the portfolio returns, we sort by the nominal interest rate (carry) for each country from low to high. The rank ordering is divided into six portfolios, into which the currency returns are assigned. P_6 is the portfolio of returns associated with the highest nominal interest rate countries and P_1 is the portfolio of returns associated with the lowest nominal interest rate countries. This table reports the two-pass procedure estimation results from a two-factor model where $sk_t(UE^{gap})$ is the maintained first factor. In the first pass, we run $N = 6$ individual time-series regressions of the excess returns on the K factors to estimate the factor 'betas,' $r^e_{i,t} = a_i + \sum_{k=1}^{K} \beta_{i,k} f^{HML}_{k,t} + \epsilon_{i,t}$, where $r^e_{i,t}$ is the excess return, $\beta_{i,k}$ is the factor beta and $f^{HML}_{k,t}$ is the high-minus-low (HML) macro risk factor. The factors considered include the high-minus-low (HML) values of the conditional mean, variance, and skewness of Δy, y^{gap}, Δc, ΔUE, UE^{gap}, π, q^{gap}, and Δq. Each HML value is equal to the average in the highest quartile minus the average in the lowest quartile. In the second pass, we run a single cross-sectional regression of the (time-series) mean excess returns on the estimated betas, $\bar{r}^e_i = \gamma + \sum_{k=1}^{K} \lambda_k \beta_{i,k} + \alpha_i$, where $\bar{r}^e_i$ is the average excess return, γ is the intercept, λ_k is the risk premia, and α_i is the pricing error. The table reports the price of risk (λ) and its associated t-ratio (using GMM standard errors), the estimated intercept (γ) and its associate t-ratio, R^2 and the Wald test on the pricing errors (Test-stat) and its associated p-value (p-val.). Bold indicates significance at the 5% level. '*' indicates significance at the 10% level.