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Currency Returns and Systematic Risk

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Abstract

We investigate the relation between consumption growth risk and currency excess returns arising from carry trade investments. We find that high interest rate is positively related to consumption growth shocks, and therefore provide low returns when consumption growth is weakening. Using Seemingly Unrelated Regressions (SUR), we are able to explore the correlation between unobservable systematic shocks across a system of asset pricing equations. We report a marked improvement in the efficiency of the estimated quantities of risk when compared with the Ordinary Least Squares (OLS) results. We employ the Consumption Capital Asset Pricing Model (CCAPM) and show that changes in the US and individual foreign country output growth are fundamental sources of currency risk. We provide empirical evidence that the risk factors associated with US GDP and foreign country's GDP growth accounts for both the cross-section of average excess returns (portfolios) and individual currencies payoffs. Finally, we reveal that regional business cycle fluctuations (Europe, America, and Asia) convey relevant information for pricing country-level profits from carry trade investments.

Key-words: carry trade; interest rates; exchange rates; currency risk.

JEL Classification code: F31, F41, G12.

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1 Introduction

The paper investigates the implications of consumption growth risk factors to carry trade. This is an investment strategy comprised of borrowing in currencies with low interest rates and investing in currencies with high interest rates. Profits from carry trade stems from the failure of the Uncovered Interest Rate Parity (UIP) condition: high interest rate currencies tend to appreciate, whereas low interest rate currencies tend to depreciate. This is the *forward premium puzzle* an anomaly well documented by the literature (Fama, 1984; Evans and Lewis, 1995; Engel, 1996; Bing, 2004; Frankel and Poonawala, 2010). As emphasized by Menkhoff et al. (2012), if carry trade investments deliver low returns during "bad times" for investors, then their profits are compensation for investors' higher risk-exposure. The theoretical reason is that agents expect the prices of currencies offering higher returns to rise in order to compensate for the expected extra earnings. As predicted for any asset using Arbitrage Pricing Theory or the Consumption Capital Asset Pricing Model (CCAPM), excess returns are due to non-diversifiable, or systematic, currency risk premia.

Evidence on the fundamental sources of currency risk premia is controversial at best. There is no conclusive evidence that supports any linear factor model for currency risk. The fierce debate between Lustig and Verdelhan (2007) and Burnside (2011b), for example, illustrates the difficulties in investigating the restrictions on the SDF that are necessary to generate the *forward premium puzzle* along the lines put forward by Backus et al. (2002). The main reason is that the variance of the excess returns tend to be much higher than the variance of consumption growth factors. This leads to a low level of precision in equation by equation Ordinary Least Squares (OLS) using the linear factor model. The disaggregation of consumption between its durable and nondurable components is primarily a way of accounting for higher SDF variability.

In our empirical analysis, we consider that the riskness of this investment strategy is associated with the correlation between currency excess returns and the consumption growth of domestic and imported goods. We apply the Consumption Capital Asset Pricing Model (CCAPM), considering open economy consumption risk factors, to rationalize the high returns generated by this currency speculative investment as compensation for consumption growth risk. We drew insights from traditional models in which bond prices reflect agents desire to smooth out domestic and imported consumption (Lucas, 1982). By assuming a linear relationship between the Gross Domestic Product (GDP) and consumption in equilibrium, we were able to test whether their growth rate could price currency risk. The search for new factors is relevant for asset pricing and continuously pushes the financial literature further in many different directions (see, for example, (Menkhoff et al., 2012; Bansal and Shaliastovich, 2013; Lettau et al., 2014; Verdelhan, 2018; Kremens and Martin, 2019)). Investigating the fundamental sources of non-diversifiable risk is a promising avenue of research for solving asset pricing puzzles in general, (Cochrane, 2017) and particularly for carry-trade currency risk (Ferreira and Moore, 2015; Ready et al., 2017; Lee and Wang, 2019; Colacito et al., 2020).

The two risk factors associated respectively with the US consumption of domestic and imported goods are proxied by US and foreign country's GDP growth. We put forward and test the hypothesis that another source of systemic risk could explain currency excess returns. This is given by undiversifiable shocks to an unobservable common factor. Preference shocks, for example, could result in a cross sectional correlation between the error term in the various currency risk equations. We use the Seemingly Unrelated Regression (SUR) and OLS methods to estimate the Consumption Capital Asset Pricing Model (CCAPM) for various currency pairs against the US dollar, considering our two proposed risk factors. The proportion of significant parameters (betas) obtained in the first pass of the Fama and MacBeth (1973) procedure was significantly higher using SUR, when compared to OLS.

We then construct portfolios of currencies sorted by these factors and analyze the performance of the factors in pricing the cross-section of average currency excess returns (portfolios). Additionally, we also explore the linkage between systematic currency risk and the comovement of business cycles across countries. We apply factor analysis to extract latent factors from our dataset. We refer to them as regional business cycles. Then, we investigate if these latent factors summarize information relevant for pricing individual currency excess returns.

In summary, our main contribution relative to the existing literature is as follows. We show that our proposed risk factors, consistent with the Euler equation, are priced in currency markets. Instead of estimating equation by equation using OLS, we explore the structural correlation between shocks across equations. Our method is based on an equilibrium hypothesis and allows one to obtain results that are asymptotically more efficient than OLS system estimation. As a result, the SUR model provides the most efficient way of estimating the asset pricing model for individual currency excess returns.

We report that the risk factors associated with consumption growth of domestic and imported goods account for both the cross-section of average excess returns (portfolios) and individual currency payoffs: the higher the interest rate, the larger the loading on the risk factors. We also uncover a change in consumption growth risk of domestic and imported goods that has taken place since the outbreak of the Global Financial Crisis (GFC) of 2008. Finally, we reveal that regional business cycles fluctuations (Europe, America, and Asia) convey relevant information for pricing country-level profits from carry trade investments. The rest of the paper is organized as follows. A review of the literature is presented in the next section. In Section 3 we lay out our empirical methodology. Section 4 reports the data, the main results and provides several robustness checks. Lastly, in Section 5, we draw the main conclusions and final remarks.

2 Literature Review

Systematic risk assessment plays a fundamental role in asset pricing theories. Lintner-Sharpe's Capital Asset Pricing Model (Lintner, 1965; Sharpe, 1964) measures the risk of an asset by its covariance with stock market return (market beta). In contrast, the CCPAM (Lucas, 1978; Breeden, 1979) attempts to measure the risk of an asset by the covariance of its return with per capita consumption (consumption beta). The main distinction between the Capital Asset Pricing Model (CAPM) and the CCAPM rests on their theoretical foundations. The CAPM comes out of the portfolio selection model developed by Markowitz (1952). Alternatively, the CCAPM stems from the solution of the household intertemporal utility maximization problem derived from its consumption. Most of the literature investigating asset pricing in the FX market have focused on two main lines of research: the first is based on macroeconomic fundamentals; and the second rests mainly on the identification of correlations between both financial market indicators and the moments of exchange rate distributions with yields from FX market investments. Our paper is directly associated with the first approach, also known as Macro-Finance, which assesses the linkage between asset pricing and economic fluctuations (Cochrane, 2017).

Asset Pricing in the FX Market. An early contribution to the literature is the work of Lustig and Verdelhan (2007). They apply the closed economy theoretical model introduced by Yogo (2006) to evaluate the yields from carry trade. These authors build portfolios of currencies sorted by interest rate differentials and show that the UIP condition fails in the cross-sectional dimension. They are primarily concerned with explaining profits from carry trade through the CCAPM, using consumption growth of durable and nondurable goods as risk factors. Conversely, Burnside (2011c) argues that the consumption betas estimated by the CCAPM are statistically insignificant and/or economically too small to rationalize the high returns from carry trade portfolios. In a similar vein, Burnside (2011a) finds that traditional CAPM risk factors, the three factors of Fama and French (1992), and the standard CAPM augmented with industrial production and the US stock market volatility do not have sufficient explanatory power of carry trade returns. These studies are silent on the importance of open economy risk factors in pricing currency excess returns, which is the focus of this paper.

By applying factor analysis to a collection of time series comprised of the returns on currency portfolios, Lustig et al. (2011) extract two principal components that are able to capture most of the data variance. The level factor (labeled as RX factor) which essentially represents the average yield on portfolios, and the slope factor (labeled as HML factor) which is the difference between returns from six and one (soreted by interest rates in ascending order). Then, they propose an asset pricing model with both a specific and a global risk factor capable of replicating the findings of the level and slope factors, respectively. They also complement their empirical work by constructing an alternative proxy for the slope factor derived from the global stock market volatility. The authors find a negative relation between yields on carry trade and stock market volatility. High interest-rate bearing currencies tend to have low returns in moments of high stock market volatility. Building on the work of Lustig et al. (2011), a novel measure of global volatility risk that comes out of currency markets is developed by Menkhoff et al. (2012). Essentially, their empirical results corroborate the evidence found by Lustig et al. (2011). High interest rate currencies are negatively correlated with global currency market volatility, offering lower returns in times of unexpected high volatility. Verdelhan (2018) also employs principal component analysis on bilateral exchange rates to extract two common factor to compare with the "Carry" and "Dollar" factors constructed from portfolios of currencies. The "Carry" factor corresponds to the change in exchange rates between baskets of high and low interest rate currencies. The "Dollar" factor is the average change in the exchange rate between the US dollar and the other currencies. Importantly, the author finds that when the "Carry" factor is applied to pricing individual currency returns, a consistent result emerges: the higher the interest rate of the corresponding currency, the larger the value of the risk factor.

A growing number of papers have yielded alternative answers to the *forward premium puzzle* and the high average payoffs from carry trade. Bansal and Shaliastovich (2013) argue that the risk premium on these investments can be explained by the differential in expected inflation and consumption volatilities between countries. The contribution of Burnside et al. (2010) stems from revealing a connection between excess returns from carry trade and the“Peso Problem" present in currency markets.[2] They argue that both the payoffs from carry trade and the values of the price kernel depend on the occurrence of peso events. The average risk-adjusted payoffs over non-peso states are positive and larger than those in the peso states. From another perspective, Hassan (2013) argues that securities issued in currencies of economies that have a large fraction of the world wealth are better for hedging consumption risk. As a consequence, larger economies have lower nominal and real interest rates than smaller countries, leading to violations of the UIP condition.

Ferreira and Moore (2015) claim that households hold foreign bonds to insure against fluctuations in prices of imported goods (precautionary savings). Employing different approaches, Fratzscher et al. (2018) assess the connection between excess returns from carry trade and the pattern of central bank interventions in FX markets, Ready et al. (2017) exploit the linkage between payoffs from carry trade and the routes of international trade, Della Corte et al. (2019) investigate the connection between currency excess returns and the occurrence of severe fluctuations in sovereign credit default swaps, and Calomiris and Mamaysky (2019) explore the association between gains from carry trade and the international spillover of monetary policy, orchestrated by the world's leading central banks, especially after the outset of the GFC. Most of the literature relies on currency portfolios to study excess returns. They are built in order to average out idiosyncratic components and, thus, focus only on systematic risk. Unlike these studies, in the present paper, the focus is on the role of open economy risk factors in pricing excess returns for individual currencies. Working with a broad set of country-level data allows us to refine our understanding of systematic versus idiosyncratic variation in each currency pair, which is cumbersome when studying excess returns using a pool of currencies.

The most recent papers have set out to explore the relationship between gains from carry trade and cross country business cycles differences. The model developed by Colacito et al. (2018) analyze the interaction between profits from carry trade and the heterogeneous exposure of countries to global economic growth shocks. The innovative approach of Berg and Mark (2018) centers on the implications of differences across countries in the values of the skewness of the realized unemployment rate gap to price currency returns. Lee and Wang (2019) scrutinize the impact of currency market jumps (very high or low change in nominal exchange rates) on carry trade profits. They argue that more sensitive currencies to negative market jumps (currency depreciation) deliver higher expected excess returns and negative jump betas are negatively correlated with country-level GDP growth and positively correlated with nominal interest rates. Lee and Wang (2019) also find a change in magnitude and sign of their jump betas for Japan and Australia after the onset of the GFC. Based on the significant currency premium associated with negative market jumps, Lee and Wang (2019) conclude that negative jump betas incorporate relevant information about business cycle conditions. Jiang et al. (2020) stress the importance of the interrelation between carry trade profits and: 1) the US monetary policy; 2) the US dollar-denominated corporate debt variations; and 3) fluctuations in foreign output. In their model, changes in the US interest rate and exchange rates affects both currency excess returns and foreign business cycles. On the other hand, Colacito et al. (2020) shows that the relative strength of the business cycles of countries measured by the output gap is a key driver of carry trade returns. They claim that the strategy of buying currencies of countries which are in their business cycle peaks and selling currencies of economies which are nearer their cycle trough generates high excess returns. Our work complements this literature, in the sense that it gives explicit consideration to the role of international business cycles in shaping excess returns from carry trade.

[2]These authors use the term “Peso Problem" to denominate the effects on inference prompted by low-probability events that do not occur in the sample.

SUR in Asset Pricing Models. Our paper is also directly associated with the literature that explores the estimation of asset pricing models by SUR models. The literature that investigates asset pricing models in equity and bond markets has long been applying SUR to obtain more efficient estimates (see, among others, Harlow and Rao (1989); Flannery et al. (1997); Pastor and Stambaugh (2002); Zarei et al. (2019)). Fama (1984) also employed the SUR technique to investigate the *forward premium puzzle*. The author finds that the joint estimation substantially improves the precision of the estimates. On the other hand, Hodrick and Srivastava (1987) employ the international intertemporal asset pricing model of Lucas (1982) combined with SUR method to examine the pricing of currency future contracts. Evans and Lewis (1995) apply the SUR model to evaluate the connection between future spot exchange rates and current forward exchange rates. By combining multiple cointegrating regressions with the SUR model, Mark et al. (2005) argue that this estimation strategy is efficient when the equilibrium errors are correlated across equations. They estimate the same equation employed by Evans and Lewis (1995) and find a more compelling evidence of the existence of a cointegrating relation between future spot exchange rates and current forward exchange rates for Germany, Japan and the United Kingdom. The work of Frankel and Poonawala (2010) also focus on the *forward premium puzzle*. The authors employ the SUR model to investigate a set of fourteen emerging market and twenty advanced economy currencies. They detect a smaller bias in emerging countries when compared to the advanced counterparts and conclude that emerging markets probably have more easily-identified trends of currency depreciation than advanced economies. As far as we are aware, SUR models have not been used in the estimation of asset returns in the foreign exchange market. Therefore, our paper contributes to the literature by introducing the SUR approach in the analysis of carry trade returns.

3 Empirical Strategy

In this section, we introduce the nature of carry trade investments and their connection to interest rates and exchange rates of countries. Then, we present the SUR model in an asset pricing setting and show the potential benefits of its application for pricing foreign exchange assets. When moving from theory to empirics, we consider that the US plays the role of the domestic economy, while the other countries are foreign markets.

3.1 Carry Trade Investment

Uncovered Interest Rate Parity Condition. For ease of interpretation, it is helpful to start by introducing the essence of the carry trade investment opportunity. When the foreign interest rate is higher than its domestic counterpart, rational risk-neutral investors should expect the foreign currency to depreciate by the difference between the two nominal interest rates. Thus, borrowing at home and investing abroad should produce a zero excess return in terms of the domestic interest rate (Lustig and Verdelhan, 2007). This is the conventional UIP condition. Nevertheless, the empirical evidence suggests that investors systematically earn excess returns through investments in high interest-rate-bearing currencies. This strategy is profitable given the frequently observed violation of the UIP condition, as emphasized by several authors (see, for example, Hansen and Hodrick (1980); Fama (1984); Lustig and Verdelhan (2007); Burnside (2011c); Menkhoff et al. (2012); Kremens and Martin (2019)). The failure of the UIP, presumably attributable to an embedded risk factor, gives rise to a range of speculative opportunities. We focus our attention on carry trade investments, the most common strategy in FX markets.

The nominal excess return, $RX_t^{j,e}$, obtained from the purchase of a unit of j currency in the forward market in period $t-1$, and the consecutive sale on the spot market at t is given by:

$$RX_{t+1}^{j,e} \equiv f_t - s_{t+1}, \tag{1}$$

where f_t is the log of the forward exchange rate and s_t is the log of the spot exchange rate, denominated in units of country's j currency per US dollar. An increase in s_t indicates an appreciation of the US dollar. Notice that the excess return can also be defined as: $RX_{t+1}^{j,e} \equiv f_t - s_t - \Delta s_{t+1}$. Under normal circumstances, forward rates satisfy the Covered Interest Rate Parity (CIP) condition: $f_t - s_t \approx i_{jt} - i_t$, where i_{jt} and i_t denote the respective country j and the US risk-free nominal interest rates, paid by a bond with the same maturity of the currency forward contract (Lustig and Verdelhan, 2009). Hence, the currency excess return approximately equals the interest rate differential between the foreign and the US economy net of exchange rate depreciation:[3]

$$RX^{j,e}_{t+1} \approx i_{jt} - i_t + -\Delta s_{t+1}, \tag{2}$$

where $\Delta s_t \approx s_{t+1} - s_t$. Typically, carry trade consists of taking long positions in high nominal interest rate currencies and short positions in low nominal interest rate currencies. The positive payoff from this strategy is certain if, over the maturity of the investment, the depreciation of the high interest-rate currency is lower than the cross-currency nominal interest rate differential. Notice that we follow Lustig et al. (2011) and Menkhoff et al. (2012) and use discrete returns and not log returns (as in Equation (2)) in our asset pricing exercises. This is to satisfy the Euler equation which is defined in levels of returns. Additionally, we follow Lustig and Verdelhan (2007) and compute real instead of nominal excess returns as follows:

$$R^{j,e}_{t+1} \equiv \left\{ \left[(1+i_{jt}) \left(\frac{S_{t-1}}{S_t} \right) - (1+i_{t-1}) \right] \left(\frac{P_{t-1}}{P_t} \right) \right\}, \tag{3}$$

where $R^{j,e}_{t+1}$ is the real *ex post* currency excess return obtained by investors who borrow at the US nominal interest rate and purchase a bond issued by country j, considering that both trades are closed at t, with the same maturity; S_t denotes the exchange rate of country j in level; and P_t is the US Consumer Price Index (CPI).

Currency Risk in Individual Equations. As mentioned, the typical specification employs "closed economy" consumption factors to model currency risk. An example is given below:

$$R^{j,e}_t = \alpha + b_1 \Delta c^n_t + b_2 \Delta c^d_t + \eta_t, \tag{4}$$

where $R^{j,e}_t$ is the real excess return obtained by the domestic resident investing in currency j bonds; α, b_1 and b_2 are parameters; Δc^n_t and Δc^d_t are the growth rates of domestic consumption of nondurable and durable goods, respectively; η_t is a white noise error term with variance σ^2_η. **?**, Lustig and Verdelhan (2011) and Burnside (2011b) all estimated equation (4) in addition to the CAPM market factor, as in Yogo (2006). The main difficulty in linear regressions like (4) lies in the precision of OLS estimates. Consider, in particular, the variance of the parameter associated with the growth of nondurable consumption:

$$\mathrm{Var}(\hat{b}_1) = \frac{\sigma^2_w}{\sum_{t=1}^{T} (\Delta c^n_t - \mu_n)^2 (1 - R^2_j)}, \tag{5}$$

where R^2_j is the R-squared from regressing Δc^n_t on Δc^d_t and on a constant term; μ_n is the sample average of Δc^n_t.

Equation (5) illustrates the non-precision problem. The variance of excess returns tends to be much higher than the variance of consumption factors, implying that σ^2_η dominates the fraction above. A high correlation between consumption factors also increases $\mathrm{Var}(\hat{\alpha_1})$ at a rate proportional to the speed at which $R^2_j \to 1$. Averaging excess returns by constructing equally weighted portfolios as Lustig and Verdelhan (2007) and Lustig and Verdelhan (2011) did, could potentially ameliorate the non-precision problem. Even if portfolios smooth away idiosyncratic risks and reduce σ^2_η, there might still be correlation across shocks induced by undiversifiable shocks.

3.2 Asset Pricing in a SUR Framework

Currency Risk in a System of Equations with Correlated Shocks. We focus on consumption betas, leaving aside the estimation of factor prices. The reason is that the main controversy in the literature lies in the significance of the factor betas. Therefore, we start by assuming a linear representation similar to equation (4). Second, we substitute the closed economy factors for open economy ones, i.e., we consider that *ex-ante* currency excess returns are determined by the consumption growth of domestic and imported goods and services. Third, we assume that the consumption of domestic and foreign goods and services is

[3]Notice that while CIP approximately holds before the GFC (see, among others, Akram et al. (2008)), the departures from CIP have risen since the outbreak of the crisis (Andersen et al., 2019). In the latter case, the forward discount accounts for both interest rate differentials and CIP deviations (Colacito et al., 2020).

linearly related to the domestic and foreign GDP, respectively. One could justify these assumptions on the basis of the intertemporal CCAPM model, for instance. Risk sharing in the symmetric perfect equilibrium is characterized by a consumption level that is equally divided between domestic and imported endowments [see **?**]. Endowments of both economies thus become stochastic discount factors in international asset pricing equations. In other words, open-economy consumption Euler equations would ensure that:

$$\mathbb{E}_t\left(R_{t+1}^{j,e}M_{t+1}\right)=0,$$

$\mathbb{E}_t$ is the expectation operator term conditional on information available at time t and the stochastic discount factor is represented by $M_{t+1}=M_{t+1}\left(\Delta y_t+1,\Delta y_{jt+1}\right)$, which is assumed to be a function of two variables: domestic output growth and the growth of imported output from country j, respectively; thus Δ stands for the first difference. As previously stated, the latter hypothesis is based on an equilibrium condition. Although obtaining the log-linear model and the beta representation is straightforward under those assumptions, we take a shortcut and postulate the following systems of equations:

$$R_t^{j,e}=\beta_{0j}+\beta_{1j}\Delta y_t+\beta_{2j}\Delta y_{jt}+u_{jt}, \tag{6}$$

where β_{0j}, $\hat{\beta}_{1j}$ and $\hat{\beta}_{2j}$ are parameters; u_{jt} is an error term, the properties of which will be explained below; $j=1,2,\ldots,J$ indexes countries whereas $t=1,2,\ldots,T$ indexes time. Using $\boldsymbol{\beta}_j\equiv(\beta_{0j},\beta_{1j},\beta_{2j})^\top$ and the definitions:

$$\mathbf{R_t^{j,e}}\equiv\begin{pmatrix}R_1^{j,e}\\R_2^{j,e}\\\vdots\\R_T^{j,e}\end{pmatrix},\ \mathbf{X_{jt}}\equiv\begin{pmatrix}1&\Delta y_1&\Delta y_{j1}\\1&\Delta y_2&\Delta y_{j2}\\\vdots&\vdots&\vdots\\1&\Delta y_T&\Delta y_{jT}\end{pmatrix},\ \mathbf{u_{jt}}\equiv\begin{pmatrix}u_{j1}\\u_{j2}\\\vdots\\u_{jT}\end{pmatrix}, \tag{7}$$

one can stack all J equations, $\forall t$, in matrices, allowing us to write:

$$\mathbf{R_t}=\mathbf{X_t}\boldsymbol{\beta}+\mathbf{u_t},$$

where

$$\mathbf{R_t}\equiv\begin{pmatrix}\mathbf{R_t^{1,e}}\\\mathbf{R_t^{2,e}}\\\vdots\\\mathbf{R_t^{J,e}}\end{pmatrix},\ \mathbf{X_t}\equiv\begin{pmatrix}\mathbf{X_{1t}}&\mathbf{0}&\cdots&\mathbf{0}\\\mathbf{0}&\mathbf{X_{2t}}&\cdots&\mathbf{0}\\\vdots&\vdots&\ddots&\vdots\\\mathbf{0}&\mathbf{0}&\cdots&\mathbf{X_{Jt}}\end{pmatrix},\ \mathbf{u_t}\equiv\begin{pmatrix}\boldsymbol{u}_{1t}\\\boldsymbol{u}_{2t}\\\vdots\\\boldsymbol{u}_{Jt}\end{pmatrix},$$

and $\boldsymbol{\beta}\equiv(\boldsymbol{\beta}^1,\boldsymbol{\beta}^2,\ldots,\boldsymbol{\beta}^J)^\top$.

Our main assumption is:

$$u_{jt}=\epsilon_{jt}+\xi_t$$

where ϵ_{jt} are idiosyncratic country shocks, which are independent of each other and also independent of the systematic shocks, ξ_t. It follows that:

$$\mathbb{E}\left(u_{jt}u_{it}|\mathbf{X_t}\right)=\sigma_{ji}\qquad\text{if } j\neq i \text{ and } \sigma_j \text{ otherwise}, \tag{8}$$

$$\mathbb{E}\left(u_{jt}u_{js}|\mathbf{X_t}\right)=\sigma_{jj}\qquad\text{if } t=s \text{ and } 0 \text{ otherwise}, \tag{9}$$

where $i\in\{1,2,\ldots,J.\}$ and $s\in\{1,2,\ldots,T.\}$. It follows that $\sigma_{jj}=\mathrm{Var}(\epsilon_{jt})+\mathrm{Var}(\xi_t)$ and $\sigma_{ji}=\sigma_{ij}=\mathrm{Var}(\xi_t)$, for $i\neq j$. In addition to the hypothesis of no autocorrelation, represented by (9), we also ensure orthogonality of the error terms and conditional homoskedasticity by assuming, respectively:

$$\mathbb{E}(\mathbf{u_t}|\mathbf{X}\mathbf{t})=0, \tag{10}$$

$$\mathbb{E}\left(\mathbf{u_t}\mathbf{u_t}^\top|\mathbf{X_t}\right)=\Omega, \tag{11}$$

where $\mathbf{u_t}^\top = [\mathbf{u_{1t}}^\top \mathbf{u_{2t}}^\top \dots \mathbf{u_{Jt}}^\top]$, $\Omega = \Sigma \otimes I$, I is the $(T \times T)$ identity matrix and Σ is the variance-covariance error matrix

$$\Sigma = \begin{pmatrix} \sigma_{11} & \sigma_{12} & \cdots & \sigma_{1J} \\ \sigma_{21} & \sigma_{22} & \cdots & \sigma_{2J} \\ \vdots & \vdots & \ddots & \vdots \\ \sigma_{J1} & \sigma_{J2} & \cdots & \sigma_{JJ} \end{pmatrix}, \tag{12}$$

where $\otimes$ is the Kronecker product. As shown above, given common or systemic shocks, the correlation between the error terms is different from zero. Systemic shocks are undiversifiable and could be related to risk preference shocks, for instance. If excess returns are subject to both sources of shocks, a SUR estimation will be asymptotically more efficient than OLS.

Seemingly Unrelated Regressions. As the SUR technique assumes a common set of instruments, the use of GDP variables instead of consumption factors ensures consistency between theory and estimation. As can be seen from (6), Δy_t appears in all equations. Using GDP as a factor has the additional advantage of ensuring higher variability of the SDF. The reason is that income variability is higher than consumption's, tending to reduce the regression's estimated standard errors. Furthermore, as emphasized by Verdelhan (2018), in the international real business cycle literature, global and country-specific shocks in each SDF (home and foreign economies) are linked to fundamental macroeconomic variables, such as consumption growth or GDP growth. In these models, the importance of world shocks would depend on the comovement of output and consumption growth between countries. SDF shocks would be related to consumption/GDP growth shocks. Hence, a high share of comovement in consumption/GDP growth would imply a high share of currency systematic risk. The underlying idea of the empirical approach is not only relevant but also theoretically promising

> "...the volatility and predictability of stock, bond and foreign exchange returns can only be consistent with arbitrage-free markets if the expected return, i.e., the discount factor, is highly variable over time. The question then is whether theoretical models are able to generate such high variability in the discount factor." [Committee (2013), p.20]

Each j equation could consistently be estimated by OLS, however, SUR estimation ensures efficiency. The procedure is based on a General Least Squares (GLS) estimator, which takes into account the correlation between the error terms of the different equations (Zellner, 1962). The SUR estimator is:

$$\hat{\beta} = [\mathbf{X_t}^\top \Omega^{-1} \mathbf{X_t}]^{-1} \mathbf{X_t}^\top \Omega^{-1} \mathbf{R_t} = [\mathbf{X_t}^\top (\Sigma^{-1} \otimes \mathbf{I}) \mathbf{X_t}]^{-1} \mathbf{X_t}^\top (\Sigma^{-1} \otimes \mathbf{I}) \mathbf{R_t}. \tag{13}$$

The Σ matrix is not known, so it needs to be estimated. The two step procedure starts by using OLS residuals of the individual estimated equations in order to build a consistent estimator for Σ:

$$\hat{\sigma}_{ij} = s_{ij} = T^{-1} \sum_{t=1}^{T} (\hat{u}_{jt} \hat{u}_{it}) \tag{14}$$

$$S \equiv \begin{pmatrix} s_{11} & s_{12} & \cdots & s_{1n} \\ s_{21} & s_{22} & \cdots & s_{2n} \\ \vdots & \vdots & \ddots & \vdots \\ s_{J1} & s_{J2} & \cdots & s_{JJ} \end{pmatrix}. \tag{15}$$

Substitution of Σ for S gives the SUR estimator - a feasible GLS - which is analogous to the maximum likelihood estimator. SUR and OLS estimators will be asymptotically equal (see, for example, Hayashi (2000)) if: *i*) there is a common set of regressors in every equation j, or *ii*) there is no correlation between equations: $\sigma_{ij} = 0$, $\forall\, i \neq j$. The first possibility is ruled out since there is only one common regressor in every equation (Δy_t). Therefore, asymptotic equivalence between OLS and SUR would only happen if the J equations were not correlated. As we show next, our empirical application considers the decisions of a US representative agent. Hence, undiversifiable shocks impact her consumption/portfolio allocation decisions and affect excess returns of currency pairs against the US dollar.

4 Data and Results

Throughout this section, we provide empirical evidence of the efficiency gain obtained by applying the SUR instead of the OLS model to estimate the CCAPM with our proposed risk factors. Then, we explore further the residual structure of the SUR estimation through factor analysis. We find that fluctuations in regional business cycle seem to play an important role in explaining currency excess returns. Finally, we conduct an extensive series of robustness exercises to figurue out whether the main findings of our paper hinge on a particular aspect of the specification of our empirical model. These adjustments provided further evidence on the significant role of our proposed risk factors in the pricing of carry trade investments. We begin by describing our dataset and the sample period analyzed, followed by a preliminary discussion on currency excess returns and business cycle fluctuations to contextualize the main findings obtained.

4.1 Data and Excess Return Construction

Data Source. Data on nominal interest rates, exchange rates, GDP, and the consumer price index come from the International Financial Statistics (IFS) database published by the International Monetary Fund (IMF). In the absence of information on nominal interest rates and GDP at the IMF (IFS), we resorted to alternative sources to obtain the data.[4]

Countries and Period Selection. The period spans from 1999:Q1 to 2019:Q4 comprising a window of 42 quarters past the end of the most recent US recession (i.e., 2009:Q2, according to the NBER dating cycle committee). Quarterly data was chosen as an attempt to account for business cycle frequencies. We focus our attention on the interval starting in 1999 to contemplate a period marked by growing international commercial and financial openness. As outlined by Lustig and Verdelhan (2007), the restrictions imposed by the Euler equation on the joint distribution of exchange and interest rates are coherent only if foreign investors are allowed to purchase local assets. As noted by Lane and Ferretti (2003), the process of financial globalization, promoted by capital account liberalization, electronic trading, increasing flow of information across economies, and falling transaction costs, has led to a large expansion in cross-border asset trading. Besides, as pointed by Coeurdacier and Rey (2013), the reduction in barriers to international trade has also been especially notorious over the last decades. They present empirical evidence on the growth in cross-border financial diversification among investors. Therefore, we believe it is reasonable to regard our sample period as appropriate, given that most countries, particularly emerging economies, have been part of this unprecedented wave of international financial integration and trade openness since the end of the 1990s.

The set of economies selected for this study accounts for more than 90% of world GDP in US dollars of 2018[5], and for approximately 90% of bilateral foreign currency turnover in April 2019 (Bank for Internatinal Settlements, 2019). In our analysis, we consider a panel of 48 developed and developing countries. Some of these economies have pegged their exchange rates partly or completely to the US dollar at various points in time. These markets differ in the level of economic development, international financial integration, and currency liquidity, hence, there are significant cross-sectional differences in the data. Therefore, we view our sample as an appropriate set in which to test the predictions of our model.

Carry Trade Returns. In order to construct the excess return data, the following variables were extracted from either the IFS, OECD, or complementary sources: nominal interest rate of each country j; end of period exchange rates, measured as the domestic price of a US dollar, and the US consumer price index (CPI). Regarding interest rates, treasury bills were the most common rates chosen for a proxy of returns on short-term bonds. When those interest rates were not available, we used Money Market rates.

[4]Due to lack of data at the IMF (IFS), we collected data on the 3-month money market rate at the Organization for Economic Co-operation and Development (OECD) for the following countries: Canada (2017:M5 to 2019:M12), Japan (2017:M7 to 2019:M12), Sweden (2017:M6 to 2019:M12), Belgium (2017:M12 to 2019:M12), Finland (2018:M2 to 2019:M12), France (2017:M6 to 2019:M12), Ireland (2017:M4 to 2019:M12), Lithuania (2017:M6 to 2019:M12), Luxembourg (2017:M6 to 2019:M12), The Netherlands (2019:M7 to 2019:M12), Portugal (2017:M6 to 2019:M12), and Paraguay (2017:M4 to 2019:M12). We also used data on 3-month money market from the European Central Bank (ECB) for the following economies: Croatia (2014:M2 to 2019:M12) and India (1991:M1 to 2004:M12). Moreover, we complemented IMF (IFS) data on gross domestic product with information provided by the OECD for the following countries: Colombia (1991:Q1 to 2000:Q1), Indonesia (1999:Q1 to 2000:Q1), Russia (1999:Q01 to 2002Q4), and the United Kingdom (2017:Q4 to 2019:Q4).

[5]Based on information published by the IMF.

In the absence of the latter, we selected Government Bonds as the risk-free rate, and finally, in the last case we used Deposit Rates. This order aimed to match an inherent property of carry trade investments, the fact that they are characterized by investments in short-term risk-free assets. Monthly series of excess returns were built according to Equation (3).

We build monthly series of carry trade returns, covering the period from 1999:M1 to 2019:M12 based on Equation (13), where the domestic economy is represented by the US and the other countries are the foreign markets. Consequently, from the US investor's perspective, the profit arises from the purchase of a foreign treasury bill or an equivalent instrument in period t, followed by the sale of the asset and the conversion of the proceeds back into US dollars, in time $t+1$. Carry trade returns are annualized (we multiply the monthly excess returns by 12) and a simple quarterly mean was used to obtain average excess returns per quarter (we followed the literature to obtain the carry trade returns. See, among others, Burnside et al. (2010); Lustig et al. (2011); Corte et al. (2016)).[6] Regarding the explanatory variables, we used an index of real GDP from the IFS (2010=100) in to construct a proxy for quarterly GDP growth. We corrected the raw index of real GDP data with the multiplicative seasonally adjustment "X-13 Arima" for the following countries: Iceland, Turkey, Paraguay, Russia, and Indonesia. Other economies already have seasonal adjustment from the source.

4.2 Results

Carry Trade Returns and Business Cycles. Country-level carry trade returns vary considerably across countries as shown in Table 1. Overall, the averages of the annualized excess returns are higher for high-interest-rate economies (Australia, Brazil, Turkey, etc.) than for low-interest-rate markets (Belgium, Japan, Sweden, etc.). Currency excess returns are also very volatile. For example, the respective annualized average profits of 8.34% and 13.28% for Brazil and Turkey are matched with standard deviations of 36.07% and 34.37%, respectively. On the other hand, the respective annualized average excess returns of -0.96% and -0.23% for Japan and Sweden are accompanied by standard deviations of 21.47% and 22.13%. The table also displays quarterly information on the GDP growth rates for each country. Overall, Asian countries present the highest averages of GDP growth rates (India, Indonesia, Philippines, etc.).

For illustrative purposes, we plot in Figure 1 the joint dynamics of the US and foreign country's GDP growth together with its corresponding excess return over time. In general, Panel (a) and (b) reveal that carry trade investments in the Brazilian and Turkish currencies generate losses during the "bad" states of the US business cycle. On the other hand, Panel (d) shows the opposite for the Japanese yen, currency excess returns are positive and high in moments of low US GDP growth. The main point to draw from Panel (c) is that it appears that the joint dynamics of the US GDP growth and carry trade profits, derived from investments in risk-free German bonds, have changed somewhat after mid-2008. Until mid-2008 the German excess returns provided a hedge against "bad" states in the US business cycles. After mid-2008, carry trade excess returns turned out to be negative during moments of low/negative US GDP growth.

As can be seen in Figure 2, a closer look at the data on GDP growth and currency excess returns for this set of countries confirms the peculiar behavior found in Figure 1. The shaded areas represent US recessions according to the NBER (National Bureau of Economic Research) recession index. It helps clarify the comovement between currency excess returns and the US GDP growth during recessions. Precisely, the excess returns of the Brazilian and Turkish currencies were negative during the two periods of the US economic downturn. Carry trade investments in the Japanese yen work as a hedge for US investors during economic recessions. On the other hand, carry trade investments in the German euro generated profits during the 2001 economic deceleration and losses in the 2008 economic crisis.

OLS Models. Bonds from countries whose consumption growth betas are negative can be considered insurance against fluctuations in the marginal utility of the domestic representative agent Lustig and Verdelhan (2007). Precisely, the excess returns that are negatively correlated with the US business cycle imply that, in periods of poor economic performance in the US, foreign bonds pay higher returns than domestic bonds. Panel (a) in Figure 3 reports the estimates of the $\hat{\beta}_{1j}$ (US GDP growth) and $\hat{\beta}_{2j}$ (foreign country's GDP growth), for our "unconditional OLS model", obtained from the OLS regression of

[6] It is important to note that we removed the month of entry of the European countries into the eurozone, due to the change in the currency denomination. Therefore, the quarterly carry trade returns for the quarters affected by this change were computed considering only the other two corresponding months.

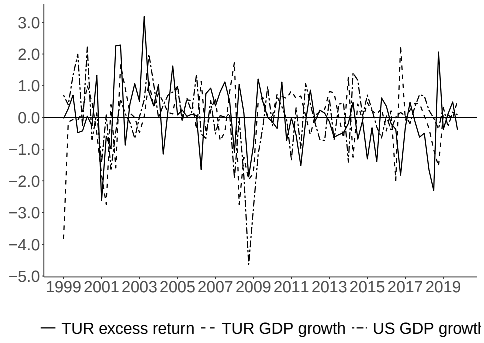
Table 1
Descriptive Statistics

The table reports descriptive statistics on carry trade returns and GDP growth rates for all countries. Means of the carry trade returns are annualized and reported in percentage points. GDP growth means are quarterly and reported in percentage points. Data are from IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.

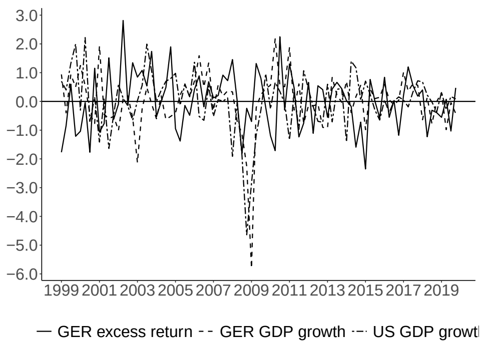
	Excess Returns				GDP Growth			
Series	**Mean**	**Std. Dev.**	**Min.**	**Max.**	**Mean**	**Std. Dev.**	**Min.**	**Max.**
Australia	3.47	23.76	-67.44	71.31	0.71	0.44	-0.48	1.71
Austria	1.49	20.33	-46.58	58.65	0.42	0.70	-2.07	1.99
Belgium	-0.09	20.24	-47.16	56.52	0.45	0.52	-2.17	1.50
Bolivia	1.51	4.17	-6.71	20.43	1.00	0.96	-2.58	3.15
Brazil	8.34	36.07	-89.93	91.10	0.59	1.14	-3.50	2.83
Bulgaria	0.05	20.48	-47.59	58.10	0.79	1.52	-5.44	8.66
Canada	1.51	16.35	-50.28	37.08	0.56	0.62	-2.26	1.81
Chile	0.93	21.67	-60.14	53.94	0.88	1.04	-4.28	3.50
Colombia	1.91	27.48	-71.75	70.36	0.89	0.91	-1.82	3.45
Croatia	1.30	20.00	-43.72	58.27	0.96	0.81	-2.40	2.81
Czech Republic	2.45	26.22	-66.42	79.87	0.71	0.83	-3.37	2.69
Denmark	0.59	20.33	-49.42	57.89	0.38	0.82	-2.36	2.98
Estonia	0.16	20.32	-46.66	53.79	1.00	2.34	-8.53	5.44
Finland	0.20	20.24	-47.00	56.66	0.45	1.26	-6.52	4.40
France	1.39	20.34	-46.44	58.53	0.36	0.47	-1.66	1.37
Germany	1.12	20.36	-46.71	58.44	0.35	0.87	-4.67	2.23
Greece	0.87	20.12	-46.76	57.17	0.04	1.59	-5.82	3.26
Hong Kong	-0.29	1.12	-3.95	3.12	0.87	1.37	-3.43	6.13
Hungary	3.91	27.75	-69.14	87.00	0.66	0.98	-4.34	2.25
Iceland	4.12	25.55	-88.70	66.71	0.76	2.50	-5.38	8.46
India	5.31	14.49	-28.97	38.61	1.70	0.99	-1.85	5.80
Indonesia	4.26	28.46	-77.30	123.94	1.28	0.50	-0.71	2.88
Ireland	0.08	20.23	-47.03	56.59	1.34	3.49	-4.77	23.23
Israel	2.96	15.38	-38.00	40.73	0.89	0.92	-1.07	4.25
Italy	0.38	20.17	-46.88	56.79	0.11	0.68	-2.78	1.44
Japan	-0.96	21.47	-57.07	57.88	0.19	1.01	-4.79	2.49
Korea	3.44	18.37	-46.45	45.52	1.08	0.97	-3.28	4.38
Lithuania	2.08	17.56	-38.74	49.83	0.96	1.86	-12.86	4.37
Luxembourg	1.14	20.32	-46.68	58.47	0.82	1.52	-3.18	5.06
Mexico	2.95	20.91	-78.46	45.83	0.51	0.97	-5.09	3.25
Netherlands	1.37	20.34	-46.60	58.57	0.41	0.67	-3.59	1.67
New Zealand	4.32	23.84	-53.77	64.43	0.76	0.84	-1.53	3.12
Norway	1.37	22.78	-67.06	74.72	0.44	1.04	-2.48	3.54
Paraguay	2.04	21.61	-76.74	61.59	0.76	2.70	-10.51	7.91
Philippines	1.61	11.70	-27.05	34.59	1.33	0.78	-2.33	3.42
Portugal	2.85	20.42	-44.91	58.68	0.25	0.78	-2.52	2.23
Romania	6.00	21.78	-55.78	68.84	0.95	1.41	-4.07	4.67
Russia	0.51	25.79	-120.24	43.78	0.92	1.32	-3.53	3.90
Singapore	0.84	10.63	-21.06	25.12	1.26	1.79	-2.87	6.22
Slovakia	4.88	21.07	-46.28	56.76	0.87	1.57	-7.04	4.44
Spain	0.44	20.23	-46.88	57.00	0.47	0.68	-2.61	1.60
Sweden	-0.23	22.13	-48.42	62.46	0.58	0.88	-3.55	2.67
Switzerland	2.10	19.57	-32.37	55.67	0.50	0.64	-2.80	2.43
Taiwan	0.06	10.62	-24.82	26.15	0.96	1.71	-4.48	8.42
Thailand	7.11	16.81	-44.25	39.31	0.96	1.68	-6.28	9.38
Turkey	13.28	34.37	-76.56	122.56	1.10	2.48	-8.39	6.69
United Kingdom	-0.07	18.43	-80.29	59.62	0.49	0.64	-2.25	3.01
Uruguay	9.10	22.26	-35.84	86.77	0.58	1.77	-5.81	5.35
United States	0	0	0	0	0.54	0.58	-2.16	1.83

a modified versions of Equation (6): $R_t^{j,e} = \beta_{0j} + \beta_{1j}\Delta y_t + u_{jt}$ and $R_t^{j,e} = \beta_{0j} + \beta_{2j}\Delta y_{jt} + u_{jt}$.

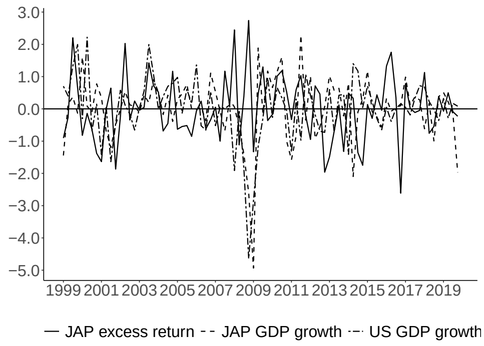
(a) Brazil



(b) Turkey



(c) Germany



(d) Japan

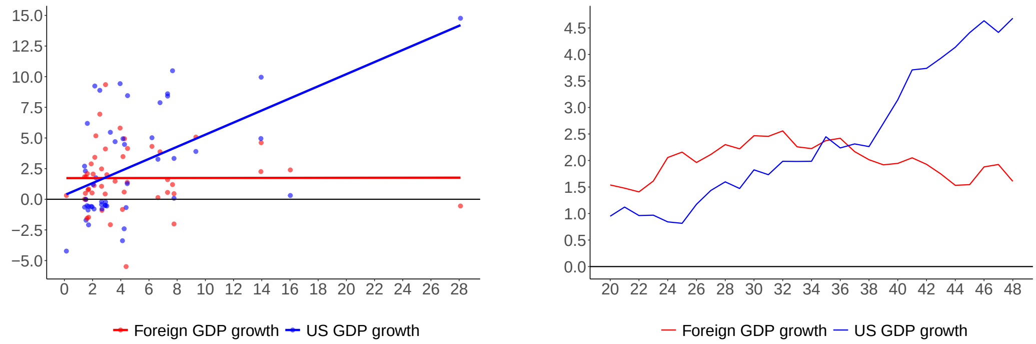
Figure 1: Types of Dynamic Behavior Between GDP Growth and Excess Returns. The figure shows, for illustration purposes, the growth rate of the US GDP along with the currency excess returns and the growth rate of GDP for Brazil (BRA), Turkey (TUR), Germany (GER), and Japan (JAP) respectively. All series are standardized. Means of the carry trade returns are annualized and reported in units of standard deviation. Means of GDP growth are quarterly and also reported in units of standard deviation. Data are from IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.

Panel (b) of Figure 3 shows the evolution of a 20-beta moving average (vertical axis) for the estimated $\hat{\beta}_{1j}$ and $\hat{\beta}_{2j}$. Before computing the moving average, we sorted the estimated beta values by the time-series averages of country's interest rates in ascending order. We end up with a vector $V^{\hat{\beta}_{1j}} = \hat{\beta}_{1j}^{1}, \hat{\beta}_{1j}^{2} .... \hat{\beta}_{1j}^{48}$; where the superscript represents the position of the beta values aligned according to the average interest rates. The first beta value belongs to the country with the lowest average interest rates. The last beta value corresponds to the economy with the highest average interest rates. The first term of the $\hat{\beta}_{1j}$ moving average, corresponds to the mean of the subset of the vector $V^{\hat{\beta}_{1j}}$ comprising the following values $\hat{\beta}_{1j}^{1}$, $\hat{\beta}_{1j}^{2} .... \hat{\beta}_{1j}^{20}$. The second term of the $\hat{\beta}_{1j}$ moving average is formed by the mean of the subset of the vector $V^{\hat{\beta}_{1j}}$ consisting of the following values $\hat{\beta}_{1j}^{2}$, $\hat{\beta}_{1j}^{3} .... \hat{\beta}_{1j}^{21}$. The last $\hat{\beta}_{1j}$ used to compute each term of the moving average is indicated in the horizontal axis. Similar considerations apply to the 20-beta moving average for the estimated $\hat{\beta}_{2j}$.

The results of Panel (a) in Figure 3 imply the existence of a positive relation between the $\hat{\beta}_{1j}$ values (vertical axis) and the means of interest rates of each country (horizontal axis). This result corroborates the findings of (Lustig and Verdelhan, 2007) for betas associated with the US consumption growth of nondurable and durable goods. Concerning $\hat{\beta}_{2j}$, it seems that there is no clear pattern of behavior. Panel (b) in the figure confirms the results presented for the beta data points. Likewise, a visual analysis of the panels in Figure 4 reinforces our conclusions from Figure 3. The only difference stems from the fact that the betas were obtained from the OLS regression of the “conditional OLS model" presented in Equation (6).

Figure 2: Excess Returns and US Business Cycles: Low vs High Average Interest Countries. The figure shows, for illustration purposes, the excess returns obtained from taking a short position in US bonds and a long position in bonds issued by four foreign markets. The left panel displays the evolution over time of profits from this investment for Germany (GER) and Japan (JAP), countries with low average interest rates. The right panel exhibits the excess returns for Brazil (BRA) and Turkey (TUR), economies with high average interest rates. The shaded areas are NBER recessions. Carry trade returns are annualized and reported in percentage points. Data are from IMF (IFS), NBER, and complementary sources. The sample period is 1999:Q1-2019:Q4.

In order to further explore the relation between the beta values and the average interest rates of each country, we run cross-sectional OLS regressions of the estimated $\hat{\beta}_{1j}$ and $\hat{\beta}_{2j}$ values on each of the following regressors: 1) the time-series average of each country interest rates; and, 2) the interest rate rank position of each country. We also added a constant term to every regression.

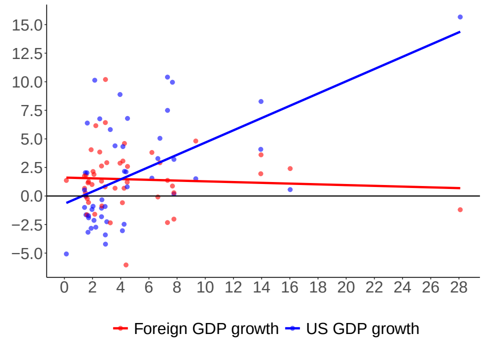


(a) $\hat{\beta}_{1j}$ and $\hat{\beta}_{2j}$

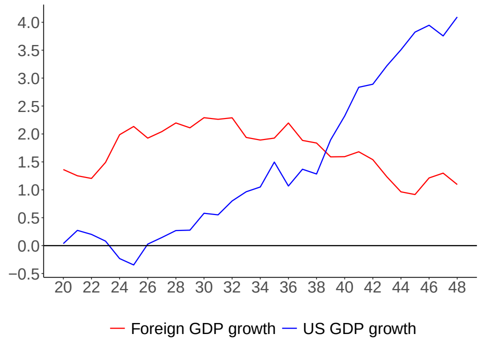
(b) $\hat{\beta}_{1j}$ and $\hat{\beta}_{2j}$ moving averages

Figure 3: Unconditional OLS Model: Estimated Beta Values. The figure shows the betas values obtained from the regression of the “unconditional OLS model". Panel (a) displays the means of interest rates of each country (horizontal axis) and the estimated $\hat{\beta}_{1j}$ values (vertical axis) obtained from the OLS regression of a modified version of Equation (6): $R_t^{j,e} = \beta_{0j} + \beta_{1j}\Delta y_t + u_{jt}$. Panel (a) also plots the estimated $\hat{\beta}_{2j}$ values obtained from the OLS regression of a modified version of Equation (6): $R_t^{j,e} = \beta_{0j} + \beta_{2j}\Delta y_{jt} + u_{jt}$. The Newey and West (1987) heteroskedasticity-consistent standard errors were used to compute the t-statistics of the estimates. Panel (b) exhibits the evolution of a 20-beta moving average (vertical axis) for both $\hat{\beta}_{1j}$ (US GDP growth) and $\hat{\beta}_{2j}$ (foreign country's GDP growth). Before computing the moving average, we sorted the estimated beta values by the time-series means of country's interest rates in ascending order (horizontal axis) to capture the effect of the increase in the rates on the beta values. Data are from IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.

Table 2 is a representative summary of the results. Here again, we find evidence of a positive relation between $\hat{\beta}_{1j}$ values and average interest rates. When considering the slope of the country's interest rates, both the "unconditional" and "conditional" models generate estimates that are positive and statistically significant coefficient at the 1% level (0.55 and 0.54, respectively). Similarly, an increase of one unit in the position of the excess returns in the ranking of average interest rates increases the beta values by 0.17 in both models, at the 1% confidence level. The results for the $\hat{\beta}_{2j}$ values, however, were statistically insignificant and/or slightly negative.



(a) $\hat{\beta}_{1j}$ and $\hat{\beta}_{2j}$



(b) $\hat{\beta}_{1j}$ and $\hat{\beta}_{2j}$ moving averages

Figure 4: Conditional OLS Model: Estimated Beta Values. The figure shows the beta values obtained from the regression of the "conditional OLS model". Panel (a) displays the means of interest rates of each country (horizontal axis) and the estimated $\hat{\beta}_{1j}$ and $\hat{\beta}_{2j}$ values (vertical axis) obtained from the OLS regression of Equation (6). The Newey and West (1987) heteroskedasticity-consistent standard errors were used to compute the t-statistics of the estimates. Panel (b) exhibits the evolution of a 20-beta moving average (vertical axis) for both $\hat{\beta}_{1j}$ (US GDP growth) and $\hat{\beta}_{2j}$ (foreign country's GDP growth). Before computing the moving average, we sorted the estimated beta values by the time-series means of country's interest rates in ascending order (horizontal axis) to capture the effect of the increase in the rates on the beta values. Data are from IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.

SUR Model. Results from the SUR estimation of the model of simultaneous equations defined in Equation (6) are presented in Panels (a) and (b) of Figure 5. As can be inferred from the visual analysis, there are no relevant differences between the magnitudes of the betas in comparison to the OLS models.[7]

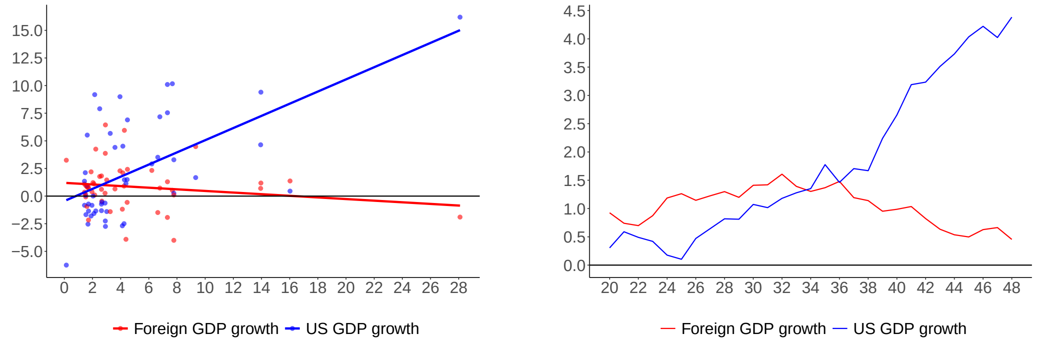
Table 2
Beta Regressions on the Average Interest Rate and on the Rank Position

The table presents the results of cross-section regressions of the estimated beta values ($\hat{\beta}_{1j}$ and $\hat{\beta}_{2j}$) on the average interest rate and on the rank position of interest rates of each country j, which are denoted by $\hat{\beta}_{jAVGIT}$ and $\hat{\beta}_{jRP}$, respectively. We construct the rank of interest rates by classifying countries by their rates in ascending order. The Newey and West (1987) heteroskedasticity-consistent standard errors were used to compute the t-statistics of the estimates. Note: *** denotes significance at 1%, ** denotes significance at 5%, * denotes significance at 10%. Data are from IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.

	Unconditional OLS		**Conditional OLS**		**SUR**	
	Intercept	**Slope**	**Intercept**	**Slope**	**Intercept**	**Slope**
$\hat{\beta}_{1jAVGIT}$	-0.45***	0.55***	-0.68	0.54***	0.33	0.49***
$\hat{\beta}_{1jRP}$	-1.99**	0.17***	-2.16**	0.17***	-1.24	0.16***
$\hat{\beta}_{2jAVGIT}$	1.20***	-0.07**	1.61***	-0.06	1.73***	0.00
$\hat{\beta}_{2jRP}$	1.17**	-0.01	1.38**	0.00	1.23**	0.02

[7] We also run the "unconditional OLS model", the "conditional OLS model" and the SUR models considering Germany as the only country from the eurozone (the results are not reported, but is available from the authors under request). The qualitative nature of the results remained unchanged.

Additionally, Table 2 confirms the existence of a positive relation between both the country's average interest rates and the rank position with the estimated $\hat{\beta}_{1j}s$ values obtained from the SUR regression.



(a) $\hat{\beta}_{1j}$ and $\hat{\beta}_{2j}$ (b) $\hat{\beta}_{1j}$ and $\hat{\beta}_{2j}$ moving averages

Figure 5: SUR Model: Estimated Beta Values. The figure shows the beta values obtained from the regression of the SUR model. Panel (a) displays the means of interest rates of each country (horizontal axis) and the estimated $\hat{\beta}_{1j}$ and $\hat{\beta}_{2j}$ values (vertical axis) obtained from the SUR regression of Equation (6). Graphical analysis and the correlogram of the residues were used to verify the presence of heterokesdasticity and autocorrelation. The *Q-statistic* developed by Box and Pierce (1970) was applied to test the joint hypothesis that all residual correlation coefficients, up to ten lags, were simultaneoulsy equal to zero. We found that only the residuals of the estimation of the equations associated with two countries showed signs of autocorrelation (Bolivia) and heterokesdacticity (Bolivia and Russia), at the 5% significance level. We ran the SUR model without these two countries and found results very similar to the complete model. Thus, we performed the following analysis considering our entire sample of countries. Panel (b) exhibits the evolution of a 20-beta moving average (vertical axis) for both $\hat{\beta}_{1j}$ (US GDP growth) and $\hat{\beta}_{2j}$ (foreign country's GDP growth). Before computing the moving average, we sorted the estimated beta values by the time-series means of country's interest rates in ascending order (horizontal axis) to capture the effect of the increase in the rates on the beta values. Data are from IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.

Table 3 looks at the individual and joint significance of the beta estimates of the "conditional OLS" and the SUR models. The key finding of our paper emerges from this table: the remarkable improvement in the precision provided by the SUR model. For example, around 6% and 20% of $\hat{\beta}_{1j}s$ and $\hat{\beta}_{2j}s$ coefficients are statistically significant at the 5% level in the "conditional OLS model". These figures increase to roughly 16% and 56% for the respective estimates in the SUR model. On the other hand, only about 29% of the beta estimates in the equation by equation OLS regression of the "conditional OLS model", were not jointly equal to zero (F-Statistic) at the 10% level. In comparison, the SUR model reached a level close to 71% of joint statistical significance. These results suggest that the proportion of significant parameters is higher than what would be expected by chance. They also imply that our choice of factors

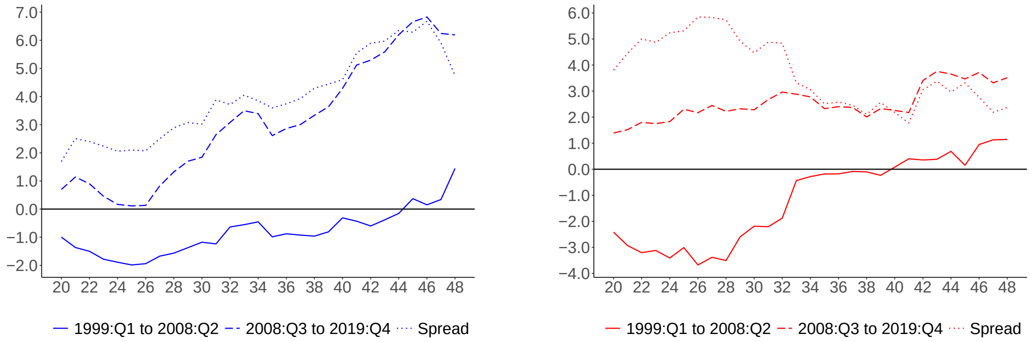
Table 3
Efficiency Gain from SUR Model in Comparison to the Conditional OLS Model

The table reports the percentage of slope coefficients, $\hat{\beta}_{1j}$ and $\hat{\beta}_{2j}$, which is statistically significant at 5%, 10%, and 15% level. The joint significance represents the F-test of the general significance of the model (joint hypotheses that both coefficients are equal to zero). The beta coefficients were obtained from the regression of the annualized excess returns of each currency pair on the US GDP growth rate and the corresponding country's GDP growth rate (Equation (6)) using the OLS and SUR methods. Annualized excess returns were computed based on Equation (3). Data are from IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.

	Conditional OLS			**SUR**		
	5%	**10%**	**15%**	**5%**	**10%**	**15%**
$\hat{\beta}_{1j}$	6.52	6.52	15.22	15.22	17.39	21.74
$\hat{\beta}_{2j}$	21.74	23.91	28.26	58.70	63.04	67.39
Joint Significance	19.57	28.26	32.61	65.22	69.57	71.74

is very promising. In what follows, we investigate further the apparent lack of the existence of a positive correlation between the $\hat{\beta}_{2j}$ values and average interest rates identified above.[8]

Change in Consumption Growth Risk. At first sight, the result obtained regarding the lack of positive correlation between the estimated $\hat{\beta}_{2j}s$ and the average interest rates may seem puzzling given the outcomes found in the literature.[9] However, Figure 6 allows us to unravel the puzzle of the estimated $\hat{\beta}_{2j}$ values. The figure exhibits the evolution of a 20-beta moving average for $\hat{\beta}_{1j}s$ (Panel (a)) and $\hat{\beta}_{2j}s$ (Panel (b)), obtained from the OLS estimation of the "conditional model" for two sample periods: from 1999:Q1 to 2008:Q2 and from 2008:Q3 to 2019:Q4. The "spread" is the difference between the second and the first sampling periods of the betas' moving average values.



(a) $\hat{\beta}_{1j}$ moving average

(b) $\hat{\beta}_{2j}$ moving average

Figure 6: Change in Consumption Growth Risk: Evolution and Spread of Estimated Coefficients Before and After the 2008-2009 Financial Crisis. The figure presents the evolution of a 20-beta moving average (vertical axis) for both $\hat{\beta}_{1j}$ (Panel (a)) and $\hat{\beta}_{2j}$ (Panel (b)) obtained from the OLS regression of the "conditional model" as determinted by Equation (6). Before computing the moving average, we sorted the estimates of the betas by the time-series means of the country's interest rates in ascending order to capture the effect of the increase in interest rates (horizontal axis) on the beta values. We generated the moving averages for two sample intervals: from 1999:Q1 to 2008:Q2 and from 2008:Q3 to 2019:Q4. The spread is the difference between the second and the first period values of the moving averages. Data are from IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.

Comparisons of the results in the two panels clarify the role of the GFC and the 2010-2012 Eurozone Crisis in shaping the behavior of average beta values:

1. The average beta values of GDP growth in the US and foreign economies are lower in the first period (before the onset of the GFC) when compared to the second. As betas are measures of systematic risk, we can confidently assert that during the interval from 2008:Q3 to 2019:Q4, carry trade investments became more risky;

2. Overall, the magnitude of the individual positive beta values is greater in the second period. The opposite is true for the first interval, the magnitude of the individual negative beta values is greater (in absolute terms) in the first period. In fact, the spread between the average beta values is always positive. The difference in magnitudes is directly responsible for the persistence of negative mean beta values of $\hat{\beta}_{1j}s$ found in the first interval. However, it is important to note that in both periods

[8] To take into account the effect of the GFC, we also estimated the "conditional OLS" and the SUR models with an additional regressor: a dummy variable with the value of one for the period from 2008:Q3 to 2019:Q4 and zero otherwise. Overall, we found very similar results when compared to the outcomes of our baseline model (these results are not reported bu are available from the authors upon request).

[9] Overall, the bulk of the works on carry trade returns emphasizes the fact that high-interest-rate currencies (developing and emerging economies) are more risky than low-interest currencies (advanced markets). Therefore, carry trade profits would be a compensation for investor´s higher risk-exposure (Lustig and Verdelhan, 2007; Burnside et al., 2010; Menkhoff et al., 2012).

we have positive and negative individual beta values: 1) for the $\hat{\beta}_{1j}s$, 36% and 47% of them are positive, respectively; and, 2) for the $\hat{\beta}_{2j}s$, 50% and 70% of them are positive, respectively;

3. The spread of the average $\hat{\beta}_{1j}$ values increases with the average interest rate, regardless of the level of the interest rate (the values are greater for countries with high interest rates and smaller for economies with low rates). On the other hand, the spread of the average $\hat{\beta}_{2j}$ values follows a different pattern, they are larger for countries with low rates and lower for economies in the long leg of interest rates. This reflects the change in signs of the $\hat{\beta}_{2j}s$ of low-interest-rate currencies from negative (during the sample period from 1999:Q1 to 2008:Q2) to positive (during the sample period from 2008:Q3 to 2019:Q4). The average $\hat{\beta}_{2j}$ values are all positives. This variation in the $\hat{\beta}_{2j}$ values explains its puzzling fluctuations found in the 20-beta moving average for the whole sample period, documented in Figures 3(c), 4(c) and 5(c). This asymmetric behavior can be understood by the distinct effects brought about by the GFC and the Eurozone Crisis on developing and advanced economies. In light of this analysis, we now turn to investigate the possible reasons underlying the change in the systematic risk associated with the consumption growth of imported goods (proxied by the GDP growth of foreign countries).

According to Lustig and Verdelhan (2007), the CCPAM can explain the variation in returns only if the consumption betas are small/negative for low interest rate economies and large/positive for high interest rate countries. They employ the result obtained by Backus et al. (2002) and derive the restrictions on the joint distribution of consumption growth in high and low-interest-rate currencies necessary to rationalize the consumption growth risk factors in the model. The authors end up with the following equation to determine the currency risk premium:[10]

$$crp_{jt+1} = std_t(\Delta c_{t+1})[std_t(\Delta c_{t+1}) - corr_t(\Delta c_{t+1}, \Delta c_{jt+1}) std_t(\Delta c_{jt+1})], \tag{16}$$

where crp_t is the currency risk premium of country's j currency; std_t and $corr_t$ stand for standard deviation and correlation, respectively; Δc_{t+1} and Δc_{jt+1} represent the respective US and foreign log consumption growth. As emphasized by Lustig and Verdelhan (2007), the right variation in currency risk premia could be achieved with a low (high) correlation of foreign consumption with the US consumption growth for high (low) interest rate currencies and more volatile consumption growth for low-interest-rate economies.

We split the forty-eight currencies into two baskets: the first, with the twenty-four countries with the lowest interest rates, and the second, with the twenty-four countries with the highest interest rates. We computed the standard deviation of each country's GDP growth and its correlation with the US GDP growth. Next, we calculated the average of the standard deviation and correlation of each basket. Then, we applied Equation (16), using discrete GDP growth instead of log consumption growth, to compute the standard deviation, correlation, and annualized currency risk premia for the two baskets, considering two intervals: from 1999:Q1 to 2008:Q2 and from 2008:Q3 to 2019:Q4. Finally, we computed the annualized average of the real currency excess return for each basket and sample period.

Overall, we found in the data the pattern of correlation required by the CCAPM to explain the variation in carry trade returns (a correlation of 0.19 and 0.47 for the first basket and 0.07 and 0.38 for the second basket, considering the first and second sampling periods, respectively). The standard deviation was greater for the first basket (1.03% and 1.25%, respectively) than for the second basket (1.21% and 1.12%, respectively), only in the second interval. Taken together this result implies a reduction in the average currency excess returns of approximately 92% (from 0.72% p.a to 0.06% p.a) for the first basket and 53% (from 0.94% p.a to 0.44% p.a) for the second basket, between the first and second periods. This outcome is in line with the decline detected, between the two intervals, in the annualized average of the real currency excess returns of approximately 128% and 77% for the first and second baskets, respectively.[11]

It is important to note that the second period was marked by a much higher increase in uncertainty about the evolution of the business cycle in advanced markets than in developing economies. Countries

[10]To derive the currency risk premium in terms of the US and foreign SDFs, Lustig and Verdelhan (2007) impose a set of restrictive assumptions on the model. What makes it difficult to compare the magnitude of the SDF-based risk premia with real currency excess returns. However, the formula is capable of pointing the direction followed by the risk premium over time.

[11]We also performed the same exercise, alternatively employing four and six baskets. Further, we selected 2008:Q1 and 2009:Q1 as the GFC outbreak dates. The outcomes we found mirrored those reported above (these results are not reported but are available from the authors upon request).

with more leveraged financial systems, stronger credit growth, a higher degree of financial and trade openness were more severely affected and those that export commodities were relatively less damaged by the GFC (see, among others, Rose and Spiegel (2011), Lane and Milesi-Ferretti (2011), and Berkmen et al. (2012)). Furceri and Mourougane (2012) show that the long-term effect of a financial crisis (measured as the decline in potential output) is larger in more developed low-interest-rate open economies. These countries' financial systems were hardest hit by the GFC, and the reduction in financial leverage and international trade prolonged the period of low GDP growth (Furceri and Zdzienicka, 2012; Lane, 2012). In a recent paper, Lane and Milesi-Ferretti (2018) highlight the role played by the GFC and the Eurozone Crisis in curbing the growth in cross-border capital flows to and from advanced economies, with a decrease in international banking activity, and an increase in the weight of emerging economies in world GDP. In short, these authors point to the advanced economies nature of the GFC and Eurozone Crisis. What, in turn, seems to be related to the pricing of assets in the foreign exchange market.[12]

We interpret these results as directly related to the degree of change in the average $\hat{\beta}_{2j}$ values of the low interest rate currencies in the second sample period. In general, Figure (6) reveals an upward trend in both average beta values, but the spread in the average $\hat{\beta}_{2j}$ values of low-interest-rate economies was the highest. In the second period, low-interest-rate developed economies were hit harder than developing countries by the GFC and the 2012-2013 Eurozone Crisis, increasing the degree of uncertainty associated with the condition of their business cycles. Which, in turn, triggered a relatively worse performance of carry trade investments, when compared to high-interest-rate developing countries. This movement can be rationalized by: 1) the higher level of volatility in the average rate of consumption (GDP) growth of developed countries; and 2) the higher level of correlation of consumption (GDP) growth between developed countries and the US economy that we observed in the data. The higher degree of volatility in the consumption of imported goods in the USA when compared to the consumption of domestic goods in the USA also plays a role in asset pricing (Ferreira and Moore, 2015). Precisely, consumption of imported goods is much more subject to shocks in business cycle fluctuations of foreign countries than the consumption of domestic goods. Therefore, it is natural to expect larger swings in the systematic risk associated with the growth in consumption of imported goods (foreign GDP) than the growth in consumption of domestic goods after periods of economic turbulence, as shown in Figure (6). This need not be the only mechanism to explain our empirical findings, but it is perhaps the most natural candidate.

To put our results into perspective, we compare them with Andrews et al. (2020) and Lee and Wang (2019).[13] The decline in currency risk premia experienced during the interval from 2008:Q3 to 2019:Q4 is also found by Andrews et al. (2020). They argue that the widespread shrink in short-term interest rates that has occurred since 2008 may explain the decline in carry trade profits. The authors also point out that exposure to an expected growth of global GDP is an important source of risk. Countries have heterogeneous exposure to this global expectation (economies with high interest rates (Australia) have very low exposures to the expected growth of global GDP, whereas countries with low interest rates (Japan) feature a substantially higher degree of exposure to this source of risk). On the other hand, Lee and Wang (2019) find that the negative jump betas are negatively correlated with country-level GDP growth. They compute a time series of the jump beta values for Australia and Japan, spanning the period from 1999:M1 to 2015:M12. The authors document a remarkable change in the sign and magnitude of the post-2008 beta coefficients, and that change remains stable until 2015:M12. In general, these outcomes corroborate the results presented in Figure(6). In light of these findings, we now turn to analyze the variance of the residuals obtained from our SUR model in order to better understand the relation between the movements of the business cycles across countries and carry trade profits.

Factor Analysis and Regional Business Cycles. We now turn to analyze the residuals generated by the SUR estimates. Our objective is to identify common forces capable of explaining fluctuations in carry trade returns considering our set of countries. In the spirit of Lustig et al. (2011) and Verdelhan (2018), we applied factor analysis to identify a set of common factors associated with carry trade returns and fluctuations in business cycles in specific regions of the world. We used the method of principal components to investigate the variance of the residuals of each equation estimated by the SUR model.

[12] Verdelhan (2018) points out that differences in the flow of capital between countries (especially portfolio investments) account for up to 53% of the differences in systematic risk across countries and that differences in variation in exports and imports explain 17%.

[13] Notice that the connection between country-level GDP and currency risk has also been investigated by the literature considering different settings (Hassan, 2013; Ready et al., 2017; Verdelhan, 2018).

Then, we extracted the common factors capable of explaining most of the data variance and perform a regression of the currency excess returns on these factors. In other words, the common factors obtained from the analysis of the variance of the residuals are used as factors in our asset pricing model.

There are two most widely used forms to estimate a factor model. The first method is principal component analysis, which does not require the normality assumption of the data and prior specification of the number of common factors. The second alternative, the Maximum Likelihood method, on the contrary, is based on the normal density function and requires pre-specification of the number of factors. Therefore, the first step in factor analysis is to determine whether the set of variables has a normal distribution. Both the Henzel et al. (2009) and Doornik and Hansen (2008) multivariate normality tests were employed and rejected the hypothesis of joint normal distribution, at the 10% significance level.

We applied the method of principal components to the series of residuals generated by the SUR estimation. We recovered the factor loads, commonality and proportion of the total variance attributed to each principal component after applying of the Varimax orthogonal rotation method (Kaiser, 1958). We retained three latent factors, which together account for most part of the total variance of the data.[14] All eigenvalues of the respective factors are greater than one, and their representation in a spree-plot reveals that the slope of the curve levels off after the third element (this result is not reported, but is available from the authors upon request). Therefore, the three factors seem to adequately capture the main patterns of correlation of the SUR residuals.

To evaluate the goodness-of-fit of the model, we computed the anti-image correlation matrix, which confirmed the sampling adequacy of the variables used in the factor analysis. The Bartlett test of sphericity rejected the hypothesis that the correlation matrix was an identity matrix, at the 1% significance level. The Kaiser-Meyer-Olkin (KMO) measure of sampling adequacy reached an overall value of 0.88. The individual KMO indices for the following economies were below 0.60: Bolivia (0.21), Hong Kong (0.09), Indonesia (0.41), Mexico (0.56) and Philippines (0.57). However, as we regard these countries as plausible targets to carry trade investments, we kept them within the data set. In addition, Cronbach's Alpha index confirmed the high reliability and internal consistency of the constructs.[15]

Carry trade returns involve a combination of two components: nominal interest rate differentials between countries and exchange rate depreciation. On the other hand, these components are inherently connected to business cycles. For example, if a foreign country's monetary authority raises its nominal short-term interest rate, aggregate demand (consumption and investment), prices and exchange rate may be affected. This, in turn, could result in changes in the carry trade profit of a long/short position held by a US investor in treasury bills issued by the foreign government. The example emphasizes the link between carry trade profits and foreign country business cycles.[16]

Therefore, we decompose the residual variance of the SUR estimation (the part of excess return not explained by our two risk factors - GDP growth in the US and the foreign country) into components that can be attributed to different latent factors and thus, we determine how much of the variance in that part of the excess return can be explained by the common factors associated with business cycles. A general picture of the latent dimensions extracted from the factor structure of the data is given in Table 4. We combined the factor loads of each variable into a set of discernible latent dimensions. The three factors turn out to have a straightforward interpretation: they capture business cycle fluctuations in large regions of the world. These are systematic risk factors capable of explaining part of the carry trade returns. The first factor ("European Carry") loads more heavily on European countries. This means that the carry trade returns for this set of countries (constructed according to Equation (13)) are partly explained by the comovements of business cycle variables (e.g, aggregate demand, interest, and exchange rates) among these economies.[17] Similar considerations apply to the other two factors ("American" and "Asian Carry").

[14]The three factors together account for approximately 70% of common variation in the data (47%, 13%, and 9%, respectively).Hair et al. (2009) suggest that the process of extraction of principal components should continue until reaching at least a proportion of 60% of the total variance of the data.

[15]Hair et al. (2009) suggest a value of at least 0.60 (exploratory research) or 0.70 (confirmatory research). The data set as a whole reached 0.98 and the respective dimensions 1 to 3 reached 0.99, 0.90, and 0.73 (standardized variables).

[16]There is a vast literature that investigates fluctuations in domestic and international business cycles (see, among others,Kose et al. (2003), Bernanke et al. (2005), Eickmeier (2007), Boivin and Giannoni (2008), and Bordo and Helbling (2010)). Factor models are typically used to find a small set of factors that can describe the business cycle fluctuations. In general, these authors assume that the latent dimensions extracted from large data set of economic indicators (GDP, industrial production, unemployment rate, price indices, etc.) are more appropriate to characterize fluctuations in business cycles than a single indicator (GDP).

[17]Importantly, despite the presence of several members of the eurozone, there are eleven economies out of twenty-six European countries that are outside of the membership.

Table 5 allows us to understand the role of the key elements mentioned above. The table shows the beta values obtained by the OLS regression of the carry trade returns on the three factors extracted by the principal component method. As we can see in the table, the f^1 factor ("European Carry") loads are much higher for European countries (ranging from values between 11.64 and 22.70). The interpretation of the other two factors, f^2 and f^3, ("American Carry" and "Asian Carry"), are essentially the same.[18]

Table 4
Regional Business Cycle Factors

The table specifies the latent dimensions (three common factors) obtained from the application of the method of principal components to the series of residuals generated by the SUR model. Data are from IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.

Factor 1 European Carry		**Factor 2 American Carry**		**Factor 3 Asia Carry**	
Austria	Italy	Australia	Mexico	Bolivia	Phillippines
Belgium	Lithuania	Brazil	New Zealand	Hong Kong	Singapore
Bulgaria	Luxembourg	Canada	Paraguay	Indonesia	Taiwan
Croatia	The Netherlands	Chile	Russia	Japan	Thailand
Czech Republic	Norway	Colombia	Turkey	Korea	
Denmark	Portugal	India	Uruguay		
Estonia	Romania	Israel			
Finland	Slovakia				
France	Sweden				
Germany	Switzerland				
Greece	United Kingdom				
Hungary	Iceland				
Ireland					

The results of Table 5 show that an important part of the carry trade profits can be explained by the three common factors identified in the data. The high level of the adjusted R^2 of the regressions and a large number of statistically significant coefficients at the 10% level confirm the relevance of the factors to price carry trade returns (except in the cases of Bolivia and Hong Kong). For example, if a US investor engages in a carry trade by taking a long position in treasury bills issued by the government of Bulgaria, her profit would depend mainly on the behavior of the business cycles (expressed in terms of movements in interest and exchange rates) from the set of European countries. This can be inferred by the beta value of the first factor and the adjusted R^2 of the regression (approximately 19.74 and 97%, respectively). Furthermore, the carry trade returns of this strategy would be relatively less affected by changes in business cycles in American and Asian countries, given the low beta values corresponding to these two factors (3.37 and 2.54, respectively). We interpret these findings as supportive of the view that the payoffs to carry trade investments partially reflects systematic risk factors underlying the comovement of interest and exchange rates among a specific set of countries (Europe, America, and Asia).[19]

[18] Bolivia and Hong Kong have pegged their currencies to the US dollar from 2011:Q1 to 2019:Q4 and from 1999:Q1 to 2019:Q4, respectively. As carry trade returns are highly affected by changes in exchange rates, both countries were left with very low θ values and adjusted R^2.

[19] Verdelhan (2018) claims that the principal components extracted from the correlation structure of bilateral exchange rates are difficult to interpret in any micro- or macrofinance model. However, by considering a broad set of countries, we could identify three distinct unobservable factors and link them to three distinct regions (Europe, America and Asia). Each common factor has a substantial degree of correlation with the corresponding average excess return of the countries associated with each the factor. Indeed, the correlation between the "European Carry" factor and the average excess returns of European countries is 0.96, the correlation between the "American Carry" factor and the average excess returns of American countries is 0.82, and the figure for the "Asian Carry" factor and the Asian economies is 0.87. A similar correlation pattern also appears when we compare exchange rates variations. This information can be considered by investors when building portfolio of currencies.

Table 5
Carry Trade Factors: Regional Business Cycles

The table reports OLS country-level results from the time-series regression of the carry trade returns for each country "j" on the three common factors: $R_t^{j,e} = \theta_0 + \theta_1 f_t{}^1 + \theta_2 f_t{}^2 + \theta_3 f_t{}^3 + \upsilon_t . f_t{}^1$, $f_t{}^2$, and $f_t{}^3$ correspond to the respective common factors retrieved from the factors structure of the data: "European Carry" , "American Carry", and "Asian Carry"; υ_t is a white noise. The factors were obtained from the application of the method of principal components to the series of residuals generated by the estimation of Equation (3) through the SUR model. All excess returns are annualized. R^2 is the adjusted R-squared. Note that *** denotes significance at 1%, ** denotes significance at 5%, and * denotes significance at 10%. The Newey and West (1987) heteroskedasticity-consistent standard errors were used to compute the t-statistics of the estimates. Data are from IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.

Country	θ_0	θ_1	θ_2	θ_3	R^2
Australia	3.46**	11.60***	13.81*	7.82*	67.32
Austria	1.48***	19.83***	3.18***	2.56***	99.22
Belgium	-0.08	19.74***	3.07***	2.47***	99.00
Bolivia	1.51*	-0.35	-0.45	0.74	1.51
Brazil	8.33***	4.78	25.53***	8.34***	55.61
Bulgaria	0.05	19.74***	3.37***	2.54***	97.21
Canada	1.51	6.05***	9.07***	3.25***	46.57
Chile	0.92	5.28**	10.85***	5.70**	35.65
Colombia	1.91	5.20**	14.52***	12.75***	51.30
Croatia	1.29**	18.23***	4.05***	3.68***	90.28
Czech Rep.	2.44**	22.70***	6.90***	4.47***	84.27
Denmark	0.58**	19.81***	2.94***	2.43***	98.47
Estonia	0.16	19.77***	2.84***	2.81***	98.60
Finland	0.19	19.74***	3.12***	2.47***	99.03
France	1.38***	19.84***	3.21***	2.55***	99.21
Germany	1.11***	19.86***	3.23***	2.52***	99.22
Greece	0.87**	19.49***	2.86***	3.05***	98.18
Hong Kong	-0.290**	-0.01	-0.22**	0.16	2.61
Hungary	3.91***	21.00***	8.90***	4.5**	69.12
Iceland	4.11	13.24***	3.62	0.77	26.31
India	5.30***	3.55***	6.99***	4.17***	35.27
Indonesia	4.25*	0.99	6.54**	15.32***	31.95
Ireland	0.08	19.73***	3.11***	2.46***	99.01
Israel	2.96**	5.30***	9.17***	2.84***	49.05
Italy	0.37	19.68***	3.014***	2.53***	99.00
Korea	3.43*	4.75***	4.76***	11.52***	51.06
Lithuania	2.08**	14.55***	4.94***	3.22***	79.28
Luxembourg	1.14***	19.81***	3.21***	2.59***	99.22
Mexico	2.94	-1.447	15.42***	0.92	53.41
Netherlands	1.37***	19.85***	3.17***	2.55***	99.22
New Zealand	4.32***	11.55***	12.07***	6.57***	55.12
Norway	1.37	16.78***	7.68***	5.08***	69.56
Paraguay	2.03	-0.04	13.66***	-2.22	38.82
Philippines	1.60	0.32	2.87***	5.98***	29.69
Portugal	2.85***	19.76***	3.07***	2.76***	97.76
Romania	5.99***	14.75***	8.46***	1.40	59.97
Russia	0.50	8.58***	12.01***	8.47***	41.47
Singapore	0.83	6.44***	2.37***	6.33***	76.37
Slovakia	4.87***	18.48***	5.76***	0.81	83.99
Spain	0.44	19.73***	3.04***	2.53***	99.02
Sweden	-0.23	18.44***	7.23***	4.49***	83.74
Switzerland	2.09**	16.46***	-0.03	5.58***	78.08
Thailand	7.10***	2.22	1.82	11.29***	46.14
Taiwan	0.06	3.56***	2.87***	6.96***	60.16
Turkey	13.27***	7.41**	19.01***	0.05	32.83
United Kingdom	-0.06	11.64***	6.94**	0.70	52.51
Uruguay	8.72***	6.39***	12.62***	3.83***	39.24

4.3 Robustness Checks

Alternative Risk Factors. We conducted an extensive series of robustness exercises to figure out whether the main findings of our paper hinge on a particular aspect of the model specification or the estimation strategy. We began by substituting out the growth rates of the US and foreign GDP for the growth rate of the US consumption and the growth rate of the US imports by country (as a proxy for the US consumption of imported goods). We considered four different series of seasonally adjusted real personal consumption expenditure provided by the US Bureau of Economic Analysis: durable goods, nondurable goods, total goods, and total goods and services. The results of the equation by equation OLS and SUR models, presented in Table 6 in Appendix A, led to results very similar to those found in the benchmark empirical specification.

Alternative Excess Returns. We then examined whether our results were robust to an alternative way of computing carry trade returns. To this end, we built excess returns based on Equation (1) following (Corte et al., 2016) and (Colacito et al., 2020).[20] The main insight that comes out of this exercise is to confirm the efficiency gain provided by the SUR model: Table 7 in Appendix B shows that the SUR model is capable of improving the efficiency of the estimation in asset pricing models.

Portfolio Analysis. In parallel, motivated by the SUR model results on the statistical significance of the $\hat{\beta}_{2j}s$, we also explored our open economy asset pricing model considering currency portfolios excess returns rather than carry trade profits at the country level, following the recent literature (Lustig et al., 2011; Menkhoff et al., 2012; Corte et al., 2016; Colacito et al., 2020). At the end of each quarter t, we allocate all currencies into six portfolios based on the country's GDP growth rates. Portfolios are rebalanced at the end of every quarter. They are ranked from low to high country-level foreign GDP growth rate: portfolio six corresponds to countries with the highest GDP growth rates, whereas portfolio one comprises countries with the lowest GDP growth rates. We compute portfolio returns as an equally weighted average of the currency excess returns within each portfolio. The total number of currencies in each portfolio is constant over time and is equal to eight. We construct our two candidate risk factors following Lustig et al. (2011): the average currency excess return (denoted RX) and the difference between the return on portfolio six and portfolio one (denoted HML), both at each point in time. We also apply the same methodology to construct portfolios and risk factors based on country 's interest rates. Hence, we refer to RX^{gdp} and HML^{gdp} for the factors constructed based on the individual foreign GDP growth rates, whereas we assign RX^{ir} and HML^{ir} for factors created based on the country's interest rates. Notice that the correlation between HML^{gdp} and HML^{ir} is very low (approximately, 0.17) and the correlation between HML^{gdp} and RX^{ir} is negative (approximately, -0.32). Hence, HML^{gdp} incorporates novel information for pricing currency excess returns.[21]

Figure 7 provides a visual summary of portfolio returns and the profits from two carry trade strategies: 1) the RX, where a US investor buys all foreign currencies; and 2) the HML, in which a US investor goes long in the highest foreign GDP growth rates/interest rates and short in the lowest foreign GDP growth rates/interest rates. The returns of these two strategies correspond to our proposed risk factors. Three key features are worth noting from the figure. First, overall carry trade returns increase from portfolio one to six. Second, the annualized average excess return for the whole set of countries (RX portfolio) is approximately 2.60%. Third, the spread between portfolio six and one (HML portfolio) is considerably higher when currencies are sorted by interest rates (approximately 5.55%) than by foreign GDP growth (approximately 1.82%).[22] Mostly, our results corroborate the findings of Lustig et al. (2011). They sort portfolios based on interest rates and also report: 1) the same pattern of increasing excess returns from

[20] As we used discrete returns and not log returns in our empirical exercises, we modified Equation (1) as follows: $RX^{j,e}_{t+1} = \left(\frac{S_{t+1}-F_t}{S_t}\right)\left(\frac{P_{t-1}}{P_t}\right)$. We collected daily spot and one-month US forward exchange rates from Thomson Reuters and followed the value date conventions in matching the forward with the appropriate spot rate (Bekaert and Hodrick, 1993). Due to lack of data, we restricted our analysis to the following set of countries: Australia, Canada, Denmark, Germany (euro), Philippines, Hungary, India, Indonesia, Japan, Mexico, Norway, New Zealand, Czech Republic, Russian, Singapore, Sweden, Switzerland, Thailand, Turkey, and the United Kingdom.

[21] Notice that the correlation between the HML^{gdp} risk factor and the three common factors identified previously, "European", "American", and "Asian" carries, are equal to -0.41, 0.06, and 0.00. Similarly, the correlation between the RX^{gdp} risk factor and the three common factors is equal to 0.83, 0.45, and 0.29.

[22] This finding is consistent with Menkhoff et al. (2012) and Corte et al. (2016). They also find lower values for their HML (4.11% and 4.40%) for currency excess returns sorted by alternative measures instead of interest rates.

portfolio one to six; 2) the annualized average excess return of approximately 1.90% (considering all portfolios), and; 3) the HML portfolio return of approximately 8.53% (for a set of thirty-five countries spanning the period from 1983:M11 to 2009:M12). Similar results are also documented by Colacito et al. (2020) for a set of twenty-seven economies, covering the period from 1983:M10 to 2016:M1, when they construct interest rates sorted currency portfolios: annualized excess returns ranging from -0.63% to 7.17%. Verdelhan (2018) reports excess returns between 1.31% and 7.14% in portfolios one and six (sorted based on the author's "dollar factor") and Corte et al. (2016) find excess returns between 0.92% and 5.32% for portfolios sorted simultaneously by country's net foreign asset position and external liabilities denominated in domestic currency.

We now turn to analyze the degree of covariance between currency excess returns and the risk factors described above. To estimate the portfolio betas, we ran time-series OLS regressions according to the following equation: $R_t^{j,e} = \gamma_0^{\iota} + \gamma_1^{\iota} RX_t^{\iota} + \gamma_2^{\iota} HML_t^{\iota} + \tau_t^{\iota}$. j denotes portfolio one to six; ι stands for gdp (for country-level GDP growth) or for ir (for interest rates). Table 8 reports the asset pricing results for the full sample period (Panel (c.1)) and two distinct intervals: from 1999:Q1 to 2008:Q2 (Panel (c.2)) and from 2008:Q3 to 2019:Q4 (Panel (c.3)). As emphasized by Lustig et al. (2011), since our factors are returns, no-arbitrage implies that the risk prices of these factors should equal their average excess returns (RX and HML portfolio returns). Therefore, by applying the portfolio betas presented in Panel (c.1) on the risk prices (RX and HML portfolio returns) we can back out the annualized returns for the six portfolios under the no-arbitrage condition. For example, for portfolio one we obtained 2.21% and 0.93% (sorted by the foreign GDP growth and interest rates, respectively), and for portfolio six we got 4.02% and 6.43% (sorted by the foreign GDP growth and interest rates, respectively). These values differ slightly from the portfolio returns previously reported and are in line with the outcome documented by Lustig et al. (2011).

Overall, the sign and magnitudes of the portfolio betas are very close to those found by Lustig et al. (2011). Moreover, as the RX and HML factors are orthogonal, the γ_2 of portfolio one should equal the γ_2 of portfolio six and all $\gamma_1 s$ should be equal to one. This is roughly what we find in our estimates and is also reported by Lustig et al. (2011). It is important to note that the beta values of portfolios sorted by the foreign country's GDP growth vary considerably from the first sample period (Panel (c.2)) to the second (Panel (c.3)). This change can be better understood by analyzing the responses of excess returns at the country level to our RX^{gdp} and HML^{gdp} risk factors.

Currency-pairs Analysis. To assess the performance of the RX^{gdp} and HML^{gdp} risk factors in pricing bilateral carry trade profits we follow the strategy employed by (Verdelhan, 2018). The author constructs two risk factors (the "carry" and the "dollar" factor) from portfolios of currencies. The "carry" factor captures the change in exchange rates between baskets of high and low-interest rate countries, whereas the "dollar" factor corresponds to the average change in the exchange rate considering all currencies (defined concerning the US dollar). (Verdelhan, 2018) considers a sample of developed and developing economies, spanning the period from 1983:M11 to 2010:M12, and runs equation by equation OLS regressions to estimate the responses of bilateral exchange rates to the two proposed factors.

Table 9 is a representative summary of the results of the following two equation by equation OLS regression: $R_t^{j,e} = \phi_0 + \phi_1 RX_t^{gdp} + \phi_2 HML_t^{gdp} + \chi_t$ ("conditional OLS model") and $R_t^{j,e} = \psi_0 + \psi_1 HML_t^{gdp} + \tau_t$ ("unconditional OLS model"). Overall, both factors appear highly significant. When the two factors are combined, the adjusted $R^2 s$ ranges (except Hong Kong and Bolivia) from around 6% (Paraguay) and 86% (Germany). When we considered only the HML^{gdp} factor, the adjusted $R^2 s$ ranges from -1% (Mexico) to 21% (Lithuania). (Verdelhan, 2018) finds very similar results: adjusted $R^2 s$ ranging from 20% to 90% (developed countries) and from 10% to 75% (developing countries) for the model with the two factors; and, with the "carry" factor, adjusted $R^2 s$ ranging from 0% to 23% (developed countries).

In order to investigate a possible change in behavior of the HML_t^{gdp} factor loads triggered by the GFC, we estimated the "unconditional OLS model" covering two periods: from 1999:Q1 to 2008:Q2 and from 2008:Q3 to 2019:Q4 (these results are not reported but are available from the authors upon request). These regressions produced outcomes very similar to those shown in Table 9, nevertheless, it is important to highlight three aspects of the results obtained: first, in the first period, we found considerably different estimates of ψ_1 (all statically significant at the 5% level) for Brazil, Mexico, and Japan (1.27, 0.66, and -1.27, respectively); second, in the second interval, we observed an increase in the HML^{gdp} factor loads for approximately 65% of the currencies; third, we sorted the HML^{gdp} factor loads by country's interest rates in ascending order and split them into two baskets containing twenty-four currencies. Then, we computed the percentage change for each country in the HML^{gdp} factor loads, between the first and the second periods, and calculated the simple mean of these values within each basket. We found an average increase of 16% for the low-interest-rate economies and a decline of 31% for the high-interest-rate countries in the HML^{gdp} factor loads. We repeated the same exercise considering four and six equally weighted baskets of currencies. We obtained essentially the same qualitative results.

The change in the HML^{gdp} loads, after the post-mid 2008 period, can also be seen from a country-level perspective. Panel (a) and (b) of Figure 8 reveal that the higher the interest rate, the larger the loading on the HML^{gdp} factor. Additionally, as we can see from Panel (b) and (c) of this figure, overall, the estimated $\psi_1 s$ values increase in the second period (from 1999:Q1 to 2008:Q2) when compared with the first period (from 2008:Q3 to 2019:Q4). Roughly, we found that most parts of the estimated $\psi_1 s$ values shift from ranging between -0.30 and 0.35 (first interval) to -0.25 and 0.45 (second interval). Essentially, the behavior identified for the HML^{gdp} loads mirrors that reported for the consumption growth of imported goods (proxied by individual country GDP growth). To sum up, we show that our proposed risk factors, consistent with the Euler equation, are priced in currency markets. The SUR model provided the most efficient way of estimating the asset pricing model for individual currency excess returns. We demonstrate that the HML^{gdp} factor accounts for both the cross-section of average excess returns (portfolios) and individual currency payoffs: the higher the interest rate, the larger the loading on the HML^{gdp} factor. We also confirmed the change in consumption growth risk of imported goods that has taken place since mid-2008.

5 Concluding Remarks

We complemented the literature in two main directions. First, we innovated by using GDP growth rates as risk factors for explaining bilateral and portfolio currency excess returns. These new factors, which can be derived from preferences that are consistent with an open economy model, allowed us to explore the structural correlation between undiversifiable shocks across the system of asset pricing equations. We were thus able to disentangle unobservable systematic risk shocks from the observable and idiosyncratic ones. Empirically, this was accomplished by employing *seemingly unrelated regressions* on the linear factor model. Our approach was shown to improve the efficiency of the estimators of the quantities of risk (betas). The percentage of jointly significant coefficients was much higher by comparison to the equation by equation OLS estimates. Overall our results suggest that the US and foreign country's GDP growth can price currency risk in a consistent way. Our paper thus presents favorable evidence to log-linear CCAPM models and arbitrage factor models for currency risk pricing.

References

Akram, Q. F., Rime, D., and Sarno, L. (2008). Arbitrage in the foreign exchange market: Turning on the microscope. *Journal of International Economics*, 76(2):237–253.

Andersen, L., Duffie, D., and Song, Y. (2019). Funding value adjustments. *The Journal of Finance*, 74(1):145–192.

Andrews, S., Colacito, R., Croce, M. M., and Gavazzoni, F. (2020). Concealed carry. *Available at SSRN*.

Backus, D. K., Foresi, S., and Telmer, C. I. (2002). Affine Term Structure Models and the Forward Premium Anomaly. *The Journal of Finance*, 56(1):279–304.

Bank for Internatinal Settlements, B. (2019). Foreign exchange turnover in April 2019: Preliminary global results (Triennial Central Bank Survey 2019). *Bank for International Settlements, Monetary and Economic Department.*

Bansal, R. and Shaliastovich, I. (2013). A Long-Run Risks Explanation of Predictability Puzzles in Bond and Currency Markets. *The Review of Financial Studies*, 26(1):1–32.

Bekaert, G. and Hodrick, R. J. (1993). On biases in the measurement of foreign exchange risk premiuns. *Journal of International Money and Finance*, 12:115–138.

Berg, K. A. and Mark, N. C. (2018). Global macro risks in currency excess returns. *Journal of Empirical Finance*, 45:300–315.

Berkmen, S. P., Gelos, G., Rennhack, R., and Walsh, J. P. (2012). The global financial crisis: Explaining cross-country differences in the output impact. *Journal of International Money and Finance*, 31:42–59.

Bernanke, B. S., Boivin, J., and Eliasz, P. (2005). Measuring the effects of monetary policy: a factor-augmented vector autoregressive (FAVAR) approach. *The Quarterly Journal of Economics*, 120(1):387–422.

Bing, H. (2004). Is the forward premium puzzle universal? *Applied Economics Letters*, 11(2):131–134.

Boivin, J. and Giannoni, M. (2008). Global forces and monetary policy effectiveness. Technical Report 13736, National Bureau of Economic Research.

Bordo, M. and Helbling, T. (2010). International Business Cycle Synchronization in Historical Perspective. Technical Report 16103, National Bureau of Economic Research.

Box, G. E. P. and Pierce, D. A. (1970). Distribution of residual autocorrelations in autoregressive-integrated moving average time series models. *Journal of the American Statistical Association*, 65(332):1509–1526.

Breeden, D. T. (1979). An Intertemporal Asset Pricing Model with Stochastic Consumption and Investment Opportunities. *Journal of Financial Economics*, 7:265–296.

Burnside, C. (2011a). Carry trade and risk. *NBER Working Paper*, 17278.

Burnside, C. (2011b). The cross section of foreign currency risk premia and consumption growth risk: Comment. *American Economic Review*, 101(7):3456–76.

Burnside, C. (2011c). The cross section of foreign currency risk premia and consumption growth risk: Comment. *American Economic Review*, 101(7):3456–3476.

Burnside, C., Eichenbaum, M., Kleshchelski, I., and Rebelo, S. (2010). Do peso problems explain the returns to the carry trade? *The Review of Financial Studies*, 24(3):853–891.

Calomiris, C. W. and Mamaysky, H. (2019). Monetary Policy and Exchange Rate Returns: Time-Varying Risk Regimes. *NBER Working Paper*, 25714.

Cochrane, J. H. (2017). Macro-finance. *Review of Finance*, 21(3):945–985.

Coeurdacier, N. and Rey, H. (2013). Home bias in open economy financial macroeconomics. *Journal of Economic Literature*, 51(1):63–115.

Colacito, R., Croce, M. M., Gavazzoni, F., and Ready, R. (2018). Currency risk factors in a recursive multicountry economy. *The Journal of Finance*, 73(6):2719–2756.

Colacito, R., Riddiough, S. J., and Sarno, L. (2020). Business cycles and currency returns. *Journal of Financial Economics*, 137(3):659–678.

Committee, N. P. (2013). Understanding Asset Prices. *Nobel Prize in Economics documents*, 1.

Corte, Pasquale, D., Riddiough, Steven, J., and Sarno, L. (2016). Currency Premia and Global Imbalances. *The Review of Financial Studies*, 29(8):2162–2193.

Della Corte, P., Jeanneret, A., and Patelli, E. (2019). A credit-based theory of the currency risk premium. *Available at SSRN 3413785*.

Doornik, J. A. and Hansen, H. (2008). An omnibus test for univariate and multivariate normality. *Oxford Bulletin of Economics and Statistics*, 70:927–939.

Eickmeier, S. (2007). Business cycle transmission from theUS to Germany - A structural factor approach. *European Economic Review*, 51(3):521–551.

Engel, C. (1996). The forward discount anomaly and the risk premium: a survey of recent evidence. *Journal of Empirical Finance*, 3:123–192.

Evans, M. D. D. and Lewis, K. K. (1995). Do Long- Term Swings in the Dollar Affect Estimates of the Risk Premia? *Review of Financial Studies*, 8(3):709–742.

Fama, E. F. (1984). Forward and spot exchange rates. *Journal of Monetary Economics*, 14:319–338.

Fama, E. F. and French, K. R. (1992). The cross-section of expected stock returns. *The Journal of Finance*, 47(2):427–465.

Fama, E. F. and MacBeth, J. D. (1973). Risk, return, and equilibrium: Empirical tests. *The Journal of Political Economy*, pages 607–636.

Ferreira, A. L. and Moore, M. J. (2015). Carry trade e risco cambial: um conto de dois fatores. *Revista Brasileira de Economia*, 69(4):429–449.

Flannery, M. J., Hameed, A. S., and Harjes, R. H. (1997). Asset pricing, time-varying risk premia and interest rate risk. *Journal of Banking & Finance*, 21:315–335.

Frankel, J. and Poonawala, J. (2010). The forward market in emerging currencies: Less biased than in major currencies. *Journal of International Money and Finance*, 29:585–598.

Fratzscher, M., Menkhoff, L., Sarno, L., Schmeling, M., and Stoehr, T. (2018). Systematic Intervention and Currency Risk Premia. *Available at SSRN 3119907*.

Furceri, D. and Mourougane, A. (2012). The effect of financial crises on potential output: New empirical evidence from OECD cuntries. *Journal of Macroeconomics*, 34:822–832.

Furceri, D. and Zdzienicka, A. (2012). Banking Crises and Short and Medium Term Output Losses in Emerging and Developing Countries: The Role of Structural and Policy Variables. *World Development*, 40(12):2369–2378.

Hair, J. F., Black, W. C., Babin, B. J., Anderson, R. E., and Tatham, R. L. (2009). *Análise multivariada de dados*. Bookman Editora.

Hansen, L. P. and Hodrick, R. J. (1980). Forward exchange rates as optimal predictors of future spot rates: An econometric analysis. *Journal of Political Economy*, 88(5):829–853.

Harlow, W. V. and Rao, R. K. S. (1989). Asset pricing in a Generalized Mean-Lower Partial Moment Framework: Theory and Evidence. *Journal of Financial and Quantitative Analysis*, 24(3):285–311.

Hassan, T. A. (2013). Country size, currency unions, and international asset returns. *The Journal of Finance*, 68(6):2269–2308.

Hayashi, F. (2000). Econometrics. 2000. *Princeton University Press. Section.*

Henzel, S., Hülsewig, O., Mayer, E., and Wollmershäuser, T. (2009). The price puzzle revisited: Can the cost channel explain a rise in inflation after a monetary policy shock? *Journal of Macroeconomics*, 31(2):268–289.

Hodrick, R. J. and Srivastava, S. (1987). Foreign currency futures. *Journal of International Economics*, 22:1–24.

Jiang, Z., Krishnamurthy, A., and Lustig, H. (2020). Dollar safety and the global financial cycle. *NBER Working Paper*, 27682.

Kaiser, H. F. (1958). The varimax criterion for analytic rotation in factor analysis. *Psychometrika*, 23(3):187–200.

Kose, M. A., Otrok, C., and Whiteman, C. H. (2003). International Business Cycles: World, Region, and Country-Specific Factors. *American Economic Review*, 93(4):1216–1239.

Kremens, L. and Martin, I. (2019). The Quanto Theory of Exchange Rates. *American Economic Review*, 109(3):810–843.

Lane, P. R. (2012). The european sovereign debt crisis. *Journal of Economic Perspectives*, 26(3):49–68.

Lane, P. R. and Ferretti, M. M. (2003). International Financial Integration. *IMF Working Paper*, (03/86).

Lane, P. R. and Milesi-Ferretti, G. M. (2011). The Cross-Country Incidence of the Global Crisis. *IMF Economic Review*, 39(1):77–110.

Lane, P. R. and Milesi-Ferretti, G. M. (2018). The External Wealth of Nations Revisited: International Financial Integration in the Aftermath of the Global Financial Crisis. *IMF Economic Review*, 66:189–222.

Lee, S. S. and Wang, M. (2019). The impact of jumps on carry trade returns. *Journal of Financial Economics*, 131:433–455.

Lettau, M., Maggiori, M., and Weber, M. (2014). Conditional risk premia in currency markets and other asset classes. *Journal of Financial Economics*, 114(2):197–225.

Lintner, J. (1965). The valuation of risk assets and the selection of risky investments in stock portfolios and capital budgets. volume 47, pages 13–37.

Lucas, R. E. (1978). Asset prices in an exchange economy. *Econometrica*, pages 1429–1445.

Lucas, R. J. (1982). Interest rates and currency prices in a two-country world. *Journal of Monetary Economics*, 10(3):335–359.

Lustig, H., Roussanov, N., and Verdelhan, A. (2011). Common risk factors in currency markets. *The Review of Financial Studies*, 24(11):3731–3777.

Lustig, H. and Verdelhan, A. (2007). The cross section of foreign currency risk premia and consumption growth risk. *American Economic Review*, 97(1):89–117.

Lustig, H. and Verdelhan, A. (2009). Comment on "Carry Trades and Currency Crashes". In *NBER Macroeconomics Annual*, volume 23, pages 361–384. University of Chicago Press.

Lustig, H. and Verdelhan, A. (2011). The cross-section of foreign currency risk premia and consumption growth risk: Reply. *American Economic Review*, 101(7):3477–3500.

Mark, N. C., Ogaki, M., and Sul, D. (2005). Dynamic Semmingly Unrelated Cointegrating Regressions. *Review of Financial Studies*, 72:797–820.

Markowitz, H. (1952). Portfolio selection. *The Journal of Finance*, 7(1):77–91.

Menkhoff, L., Sarno, L., Schmeling, M., and Schrimpf, A. (2012). Carry trades and global foreign exchange volatility. *The Journal of Finance*, 67(2):681–718.

Newey, W. K. and West, K. D. (1987). Hypothesis Testing With Efficient Method of Moments Estimation. *International Economic Review*, 28(3):777–787.

Pastor, L. and Stambaugh, R. F. (2002). Mutual fund performance and seemingly unrelated assets. *Journal of Financial Economics*, 63:315–349.

Ready, R., Roussanov, N., and Ward, C. (2017). Commodity trade and the carry trade: A tale of two countries. *The Journal of Finance*, 72(6):2629–2684.

Rose, A. M. and Spiegel, M. M. (2011). Cross-country causes and consequences of the crisis: An update. *European Economic Review*, 55:309–324.

Sharpe, W. F. (1964). Capital asset prices: A theory of market equilibrium under conditions of risk. *The Journal of Finance*, 19(3):425–442.

Verdelhan, A. (2018). The Share of Systematic Variation in Bilateral Exchange Rates. *Journal of Finance*, 73(1):375–418.

Yogo, M. (2006). A consumption-based explanation of expected stock returns. *The Journal of Finance*, 61(2):539–580.

Zarei, A., Ariff, M., and Bhatti, I. (2019). The impact of exchange rates on stock market returns: new evidence from seven free-floating currencies. *The European Journal of Finance*, 25(14):1277–1288.

Zellner, A. (1962). An efficient method of estimating seemingly unrelated regressions and tests for aggregate bias. *Journal of the American Statistical Association*, 57:348–368.

A Domestic and Imported Goods Factors

Table 6
Efficiency Gain from SUR Model in Comparison to the Conditional OLS Model: Domestic and Imported Goods Factors

The table reports the percentage of slope coefficients, $\hat{\beta}_{1j}$ and $\hat{\beta}_{2j}$, which is statistically significant at 5%, 10%, and 15% level. The joint significance represents the F-test of the general significance of the model (joint hypotheses that both coefficients are equal to zero). The beta coefficients were obtained from the regression of the annualized excess returns of each currency pair on the US GDP growth rate and the corresponding country's GDP growth rate (Equation (6)) using the OLS and SUR methods. Annualized excess returns were computed based on Equation (3). Data are from IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.

			Durable Goods			
	Conditional OLS			**SUR**		
	5%	**10%**	**15%**	**5%**	**10%**	**15%**
$\hat{\beta}_{1j}$	4,76	14,29	23,81	11,90	16,67	26,19
$\hat{\beta}_{1j}$	23,81	30,95	45,24	54,76	64,29	64,29
Joint Significance	19,05	28,57	33,33	57,14	66,67	69,05
			Nondurable Goods			
	Conditional OLS			**SUR**		
	5%	**10%**	**15%**	**5%**	**10%**	**15%**
$\hat{\beta}_{1j}$	2,38	4,76	7,14	0,00	4,76	7,14
$\hat{\beta}_{1j}$	19,05	30,95	40,48	59,52	64,29	69,05
Joint Significance	9,52	14,29	21,43	47,62	59,52	61,90
			Total Durable and Nondurable Goods			
	Conditional OLS			**SUR**		
	5%	**10%**	**15%**	**5%**	**10%**	**15%**
$\hat{\beta}_{1j}$	11,90	14,29	16,67	9,52	16,67	19,05
$\hat{\beta}_{1j}$	19,05	30,95	45,24	57,14	64,29	66,67
Joint Significance	14,29	21,43	33,33	54,76	61,90	69,05
			Total Goods and Total Services			
	Conditional OLS			**SUR**		
	5%	**10%**	**15%**	**5%**	**10%**	**15%**
$\hat{\beta}_{1j}$	0,00	0,00	2,38	0,00	2,38	4,76
$\hat{\beta}_{1j}$	23,81	28,57	40,48	57,14	64,29	66,67
Joint Significance	9,52	14,29	21,43	50,00	54,76	57,14

B Forward and Spot Exchange Rates

Table 7
Efficiency Gain of the SUR Model Compared to the Conditional OLS Model: Forward and Spot Exchange Rates

The table reports the percentage of slope coefficients, $\hat{\beta}_{1j}$ and $\hat{\beta}_{2j}$, statitically significant at 5%, 10%, and 15% level. The joint signficance stands for the F-test of the overall significance of the model (joint hypotheses that both coefficients are equal to zero). The beta coefficients were obtained from the regression of the annualized excess returns of each currency pair on the US GDP growth rate and the corresponding country GDP growth rate (Equation (6)) by the OLS and SUR methods. We used daily spot and one-month US forward exchange rates to compute the annualized excess returns as follows: $RX^{j,e}_{t+1} = \left(\frac{S_{t+1}-F_t}{S_t}\right)\left(\frac{P_{t-1}}{P_t}\right)$. Data are from Thomson Reuters, IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.

	Conditional OLS			**SUR**		
	5%	**10%**	**15%**	**5%**	**10%**	**15%**
$\hat{\beta}_{1j}$	0.00	7.69	7.69	0.00	0.00	0.00
$\hat{\beta}_{2j}$	0.00	7.69	7.69	7.69	15.38	23.08
Joint Significance	0.00	0.00	0.00	7.69	7.69	7.69

C Currency Portfolio Analysis

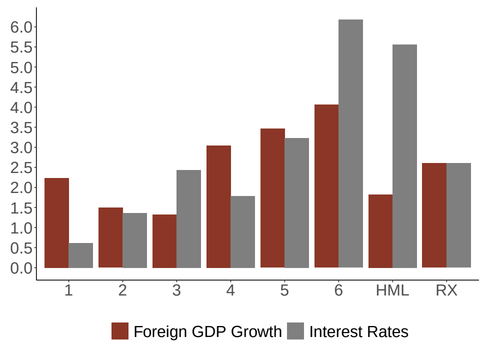


Figure 7: Currency Portfolio Excess Returns. The figure shows the excess returns of portfolio sorted by country's GDP growth rate and interest rates. They are ranked from low to high country-level GDP growth/interest rates: portfolio six corresponds to countries with the highest GDP growth/interest rates, whereas portfolio one comprises countries with the lowest GDP growth/interest rates. We compute portfolio returns as an equally weighted average of the currency excess returns within each portfolio. The total number of currencies in each portfolio is constant over time and is equal to eight. The portfolios are rebalanced at the end of each quarter. RX and HML correspond to portfolios formed by the average currency excess return and the difference between the return on portfolio six and portfolio one, respectively. All returns are annualized. Data are from IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.

Table 8
Currency Portfolio Betas: Foreign GDP Growth and Interest Rates

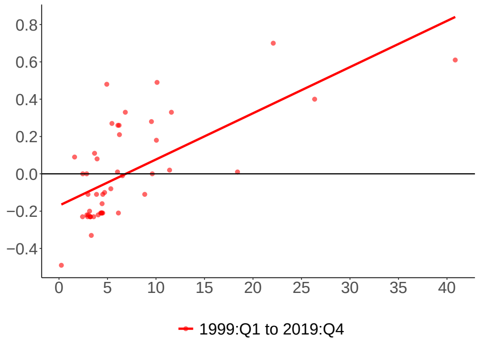
The table reports OLS country-level results from the time-series regression of the carry trade returns for each portfolio "p" on the two risk factors sorted by either country GDP growth rates or interest rates: $R_t^{p,e} = \gamma_0^\iota + \gamma_1^\iota RX_t^\iota + \gamma_2^\iota HML_t^\iota + \tau_t$. ι stands for country-level GDP growth or interest rates; τ_t is a white noise. All excess returns are annualized. R^2 is the adjusted R-squared. Note that *** denotes significance at 1%, ** denotes significance at 5%, and * denotes significance at 10%. The Newey and West (1987) heteroskedasticity-consistent standard errors were used to compute the t-statistics of the estimates. Data are from IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.

Panel (c.1): Sample Period: from 1999:Q1 to 2019:Q4								
Portfolio	γ_0^{gdp}	γ_1^{gdp}	γ_2^{gdp}	R^2	γ_0^{ir}	γ_1^{ir}	γ_2^{ir}	R^2
1	0.77**	0.90***	-0.49***	94.42	-0.26	1.02***	-0.31***	96.25
2	-1.24**	1.11***	-0.08*	91.51	-0.20	1.04***	-0.20***	93.00
3	-1.30***	1.07***	-0.09**	93.33	0.43	1.02***	-0.12**	93.21
4	0.34	1.02***	0.01	88.66	-0.91*	0.98***	0.02	90.00
5	0.66	0.97***	0.14**	85.83	1.22	0.90***	-0.06	87.54
6	0.77**	0.90***	0.50***	91.60	-0.26	1.02***	0.68***	96.29
Panel (c.2): Sample Period: from 1999:Q1 to 2008:Q2								
Portfolio	γ_0^{gdp}	γ_1^{gdp}	γ_2^{gdp}	R^2	γ_0^{ir}	γ_1^{ir}	γ_2^{ir}	R^2
1	0.08	0.87***	-0.37***	88.50	-0.78	1.03***	-0.31***	94.41
2	-1.11	1.20***	-0.06	90.48	-1.35	1.15***	-0.13	91.30
3	-1.61*	1.11***	-0.09	88.65	1.26	0.92***	-0.12**	90.80
4	1.43	0.93***	-0.13	81.40	-0.72	0.99***	-0.031	90.63
5	1.12	1.00***	0.054	79.77	2.38**	0.86***	-0.08	84.22
6	0.08	0.87***	0.62***	85.92	-0.78	1.03***	0.68***	94.53
Panel (c.3): Sample Period: from 2008:Q3 to 2019:Q4								
Portfolio	γ_0^{gdp}	γ_1^{gdp}	γ_2^{gdp}	R^2	γ_0^{ir}	γ_1^{ir}	γ_2^{ir}	R^2
1	1.73***	0.93***	-0.56***	97.78	0.04	1.02***	-0.31***	97.21
2	-1.79***	1.04***	-0.09	91.69	0.29	0.98***	-0.30***	94.85
3	-1.21**	1.05***	-0.09**	95.70	0.31	1.08***	-0.13*	94.29
4	-0.40	1.074***	0.11	92.37	-1.27	0.97***	0.11	89.05
5	-0.05	0.94***	0.20***	89.45	0.57	0.90***	-0.05	88.16
6	1.73***	0.93***	0.43***	96.41	0.04	1.02***	0.68***	97.13

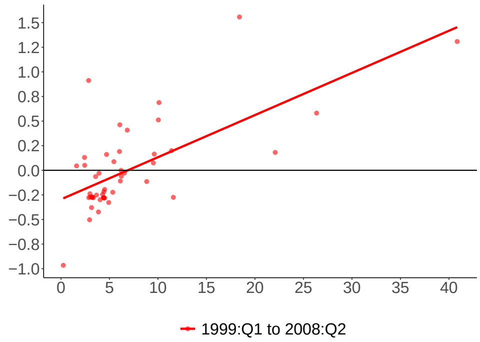
Table 9
Country-level Betas: Foreign GDP Growth

The table reports OLS country-level results from the time-series regression of the carry trade returns for each country "j" on the two risk factors: $R_t^{j,e} = \phi_0 + \phi_1 RX_t^{gdp} + \phi_2 HML_t^{gdp} + \chi_t$ (“conditional OLS model"). χ_t is a white noise. The table also shows the OLS country-level results from the time-series regression of the carry trade returns for each country "j" on the HML_t^{gdp} factor only: $R_t^{j,e} = \psi_0 + \psi_1 HML_t^{gdp} + \tau_t$ (“unconditional OLS model"). τ_t is a white noise. All excess returns are annualized. R^2 and $R^{2,gdp}$ are the adjusted R-squared for the conditional and unconditional models, respectively. Note that *** denotes significance at 1%, ** denotes significance at 5%, and * denotes significance at 10%. The Newey and West (1987) heteroskedasticity-consistent standard errors were used to compute the t-statistics of the estimates. Data are from IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.

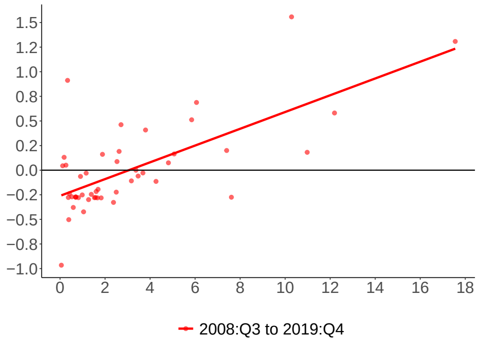
Country	ϕ_0	ϕ_1	ϕ_2	R^2	ψ_0	ψ_1	$R^{2,gdp}$
Australia	-1.25	1.44***	0.48**	66.87	3.73	-0.19	-0.50
Austria	-0.92	1.24***	-0.21**	86.22	3.38*	-0.79***	14.90
Belgium	-2.43***	1.22***	-0.23***	85.74	1.82	-0.81***	15.52
Bolivia	1.52*	-0.02	0.01	-1.40	1.43*	0.02	-0.80
Brazil	4.28	1.40***	0.70*	26.96	9.12**	0.04	-1.21
Bulgaria	-2.35**	1.23***	-0.22**	84.20	1.93	-0.80***	14.84
Canada	-0.80	0.78***	0.11	42.30	1.90	-0.24	1.09
Chile	-1.68	0.84***	0.26	25.99	1.22	-0.13	0.84
Colombia	-1.76	1.13***	0.33	29.78	2.16	-0.19	-0.73
Croatia	-0.98	1.20***	-0.16*	83.84	3.17*	-0.73***	13.17
Czech Rep.	-0.58	1.56***	-0.11	81.67	4.82*	-0.85***	10.38
Denmark	-1.79**	1.23***	-0.23**	84.83	2.46	-0.81***	15.38
Estonia	-2.21**	1.23***	-0.22**	85.02	2.04	-0.80***	15.04
Finland	-2.14**	1.22***	-0.23***	86.01	2.11	-0.81***	15.58
France	-1.03	1.24***	-0.21**	86.26	3.27	-0.79***	14.81
Germany	-1.30	1.24***	-0.21**	86.36	3.01	-0.80***	14.93
Greece	-1.58*	1.23***	-0.22**	85.60	2.67	-0.80***	15.32
Hong Kong	-0.28**	-0.00	-0.00	-2.02	-0.29**	0.00	-1.07
Hungary	0.03	1.61***	-0.01	69.24	5.62**	-0.77***	6.68
Iceland	1.41	1.00***	0.18	27.73	4.88	-0.28	0.00
India	3.04**	0.61***	0.33**	31.35	5.17***	0.04	-1.12
Indonesia	1.51	0.72***	0.49*	9.86	4.02	0.14	-0.95
Ireland	-2.24***	1.22***	-0.23***	86.01	2.00	-0.81***	15.66
Israel	1.22	0.66***	-0.10	40.64	3.52**	-0.42**	6.47
Italy	-1.95**	1.22***	-0.23***	85.72	2.28	-0.80***	15.51
Japan	-0.76	0.36***	-0.49*	12.77	0.48	-0.66***	8.55
South Korea	1.15	0.74***	0.21	28.76	3.73*	-0.13	-0.71
Lithuania	0.00	1.01***	-0.33***	81.99	3.52**	-0.81***	20.47
Luxembourg	-1.27	1.24***	-0.21**	86.16	3.02	-0.79***	14.82
Mexico	0.59	0.54**	0.28	10.39	2.46	0.03	-1.21
Netherlands	-1.03	1.24***	-0.21**	86.21	3.26	-0.80***	14.93
New Zealand	0.45	1.29***	0.27	54.29	4.93*	-0.32	0.68
Norway	-1.76	1.32***	-0.11	71.45	2.82	-0.74**	9.43
Paraguay	0.84	0.42*	0.00	5.54	2.31	-0.19	-0.41
Philippines	0.18	0.31***	0.26**	12.27	1.28	0.12	-016
Portugal	0.48	1.23***	-0.21**	84.72	4.76**	-0.79***	14.85
Romania	2.99*	1.18***	0.01	59.06	7.10***	-0.53*	4.94
Russia	-1.27	1.00***	-0.11	32.31	2.18	-0.58	4.13
Singapore	-0.42	0.57***	0.00	62.17	1.57	-0.26**	5.51
Slovakia	2.44*	1.20***	-0.08	73.13	6.63***	-0.64***	8.94
Spain	-1.89**	1.22***	-0.23**	85.80	2.35	-0.81***	15.55
Sweden	-3.52***	1.40***	-0.11	85.00	1.32	-0.77***	11.22
Switzerland	0.28	0.99***	-0.20*	60.39	3.73*	-0.67***	11.09
Thailand	6.01**	0.43***	0.09	10.84	7.53***	-0.10	-0.81
Taiwan	-1.00	0.44***	0.02	33.91	0.53	-0.18	2.00
Turkey	8.81***	1.28***	0.61**	23.53	13.26***	0.01	-0.12
United Kingdom	-2.58	0.96***	0.08	52.81	0.73	-0.37*	2.90
Uruguay	5.59***	0.95***	0.40**	30.37	8.90***	-0.04	-1.18



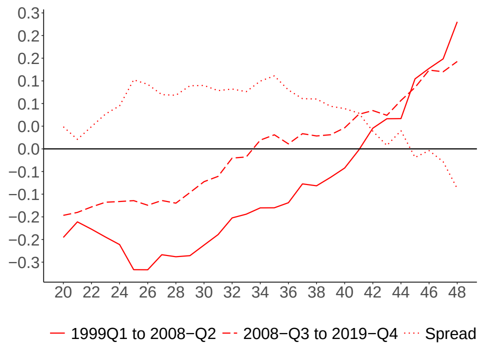
(a) Country-Level HML Beta Full Sample



(b) Country-Level HML Beta Before 2008-2009 Financial Crisis



(c) Country-Level HML Beta After 2008-2009 Financial Crisis



(d) Change in Consumption Growth Risk

Figure 8: Currency Portfolio Factors: Change in Country-Level HML Beta and in Consumption Growth Risk. The figure show the betas values obtained from the OLS regression of the time series of carry trade returns for each country "j" on the two risk factors: $R_t^{j,e} = \phi_0 + \phi_1 RX_t^{gdp} + \phi_2 HML_t^{gdp} + \chi_t$ ("conditional OLS model"). χ_t is a white noise. The horizontal axis contains the means of the interest rates of each country and the vertical axis the estimated ϕ_2 values. Panel (d) displays the evolution of a 20-ϕ_2 value moving average (vertical axis). Before computing the moving average, we sorted the ϕ_2 estimates by the country's time series averages of interest rates in ascending order to capture the effect of the increase in interest rates (horizontal axis) on the ϕ_2 values. We generated the moving averages for two sample intervals: from 1999:Q1 to 2008:Q2 and from 2008:Q3 to 2019:Q4. The spread is the difference between the second and the first period values of the moving averages. Data are from IMF (IFS) and complementary sources. The sample period is 1999:Q1-2019:Q4.