Uncertainty and Deviations from Uncovered Interest Rate Parity

Adilzhan Ismailov[†,‡] Barbara Rossi[*‡]

(Univ. Pompeu Fabra) (ICREA-Univ. Pompeu Fabra, Barcelona GSE, and CREI)

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Abstract: It is well-known that uncovered interest rate parity does not hold empirically, especially at short horizons. But is it really so? We conjecture that uncovered interest rate parity is more likely to hold in low uncertainty environments, relative to high uncertainty ones, since arbitrage opportunity gains become more uncertain in a highly unpredictable environment, thus blurring the relationship between exchange rates and interest rate differentials. In this paper, we first provide a new exchange rate uncertainty index, that measures how unpredictable exchange rates are relative to their historical past. Then we use the new measure of uncertainty to provide empirical evidence that uncovered interest rate parity does hold in five industrialized countries vis-a'-vis the US dollar at times when uncertainty is not exceptionally high, and breaks down during periods of high uncertainty.

[*]Corresponding author. E-mail: barbara.rossi@upf.edu. [†]E-mail: adilzhan.ismailov@upf.edu. [‡]Dept. of Economics, Univ. Pompeu Fabra, c. Ramon Trias Fargas 25-27, Mercè Rodoreda bldg., 08005 Barcelona, Spain.

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1 Introduction

A well-known empirical fact in international finance is that uncovered interest rate parity (UIRP) does not hold, especially at short horizons. UIRP states that, in the absence of arbitrage opportunities, the returns from investments in two countries should be equalized, once they are converted into the same currency; the implication is that interest rate differentials should predict bilateral nominal exchange rate appreciations or depreciations. UIRP is an important building block of most international macroeconomic models, and the lack of its validity is of such importance to deserve the term "UIRP puzzle". Another puzzling empirical fact about UIRP is that, not only the coefficients do not have the values predicted by the theory, but also that they are unstable over time. This paper offers an explanation to *both* these puzzles by arguing that uncertainty is one of the reasons explaining the empirical invalidity of the UIRP; that the coefficients in UIRP regressions are more likely to be close to the values predicted by UIRP at times of low uncertainty; and that their time variation is, at least partly, due to the fact that UIRP holds when uncertainty is low but does not when uncertainty is high. As we discuss further below, a large body of literature argues that the UIRP puzzle is not really a puzzle since it can be explained by time-varying risk premia. Our empirical results are consistent with this literature, as we argue that, for example, high uncertainty can be related to rare disasters, which can theoretically generate the time-varying risk premia we observe in the data. Our paper, however, has the advantage of providing both an empirical analysis as well as an empirically observable proxy that can explain deviations from UIRP.

More in detail, this paper makes two main contributions. First, it proposes a new measure of exchange rate uncertainty.[1] The novelty is not in the methodology to construct the new index, which is based on Rossi and Sekhposyan (2015), but rather its application to measure exchange rate uncertainty. To our knowledge, this is the first paper to propose *an index of exchange rate uncertainty*. We measure uncertainty at a point in time by the likelihood of observing the realized exchange rate forecast error at that point

[1] The time series of uncertainty indices are available at: barbararossi.eu/data

in time, relative to the historical distribution of exchange rate forecast errors. Since the uncertainty measure is based on forecast errors, it clearly depends on the model used to forecast exchange rates. To minimize the dependence of our empirical results on the choice of a specific model, we use Consensus survey forecasts, which have the favorable feature of being survey-based and timely incorporating a large amount of information. These survey forecasts have been used recently by Ozturk and Sheng (2016) to measure macroeconomic uncertainty; instead, we use them to construct an index of exchange rate uncertainty.

The second contribution is to make a step towards understanding why UIRP does not empirically fit the data. In fact, typical estimates of the slope are either negative or zero or too large to be reconciled with the theory (Froot and Thaler, 1990); UIRP also fails to produce competitive out-of-sample forecasts relative to the random walk (Meese and Rogoff, 1983a,b; 1988; Cheung et al., 2005; Alquist and Chinn, 2008) – see Rossi (2013) for a recent survey. Several possible explanations have been put forward in the literature. An important potential explanation is the presence of time-varying risk premia (Fama, 1984; Li et al., 2011). Other explanations include: imprecise standard errors (Baillie and Bollerslev, 2000; Rossi, 2007); small samples (Chinn and Meredith, 2004; Chinn and Quayyum, 2013; and Chen and Tsang, 2013); and rare disasters, such as currency crashes (Brunnermeier et al., 2009; Farhi and Gabaix, 2016).[2] In this paper we investigate an alternative explanation for the UIRP puzzle, namely the fact that *the uncovered interest rate parity might not hold in highly uncertain environments*, while it is more likely to hold when uncertainty is low. In fact, when uncertainty is high, investors might postpone their investment decisions, and thus create deviations from what is expected in the absence of arbitrage opportunities. Our result does not depend on the measure of uncertainty we use: in fact, the result is robust to using other measures of uncertainty, as we demonstrate in the paper. In addition, as we show, deviations from UIRP cannot be explained solely by differences in monetary policy: while it is true that for some countries (such as

[2] Avdiev et al. (2016) document instead large deviations from *covered* interest rate parity during the recent financial crisis, which they attribute to the lack of banks' ability to take on additional leverage.

Switzerland and the European Union – EU thereafter) UIRP is more likely to hold during the zero-lower bound period, the result is not true for all the countries in our sample. Furthermore, our results have direct implications for the risk premium: in fact, as we discuss, the risk premium is correlated with interest rate differentials in periods of high uncertainty, but not significantly correlated in periods of low uncertainty.

On the one hand, our main results focus on an uncertainty index based on survey forecasts, which has the advantage of not depending on a specific forecasting model; however, on the other hand, exchange rate survey forecasts are available only for a few countries, which limits the scope of the analysis. In order to extend the sample of countries, we construct an exchange rate uncertainty index based on the random walk, thus making our index suitable for big data. Among forecasting models of exchange rate determination, the random walk is a difficult benchmark to beat (Rossi, 2013). We show that our results for the main countries in our sample (Canada, the EU, Japan, Switzerland and the UK) are robust no matter whether we use surveys or the random walk to construct an uncertainty index. More importantly, we show that the UIRP puzzle is alleviated in low uncertainty environments for several of the additional countries that the extension to random walk forecast errors allows us to consider (Australia, Sweden, Denmark). For some other countries, although low uncertainty typically moves the coefficient in the right direction, it does not fully resolve the puzzle (South Africa and New Zealand); however, the latter are "commodity countries" (Chen and Rogoff, 2003; Chen et al., 2010), for which commodity prices might play a role in determining exchange rate fluctuations, which we abstract from.

This paper is related to several recent strands in the literature. The first strand is the empirical literature on the UIRP puzzle. While it is uncontroversial that the UIRP does not hold at short horizons, Chinn and Meredith (2004), Lothian and Wu (2011) and Chinn and Quayyum (2013) find more empirical evidence in favor of UIRP at longer horizons.[3] In particular, Chinn and Meredith (2004) argue that the lack of empirical evidence in favor of UIRP is due to small samples, and find that UIRP holds at longer

[3] Note that monetary models of exchange rates are more likely to hold at long horizons as well (Mark, 1995).

horizons (above one year) in the longer sample of data they have available. Lothian and Wu (2011) examine historical data from 1800 to 1999, and find that the UIRP regression slope is positive for the longest sample, and the strong negative relation found in the literature is a feature of the late 1970s and the 1980s. Finally, Chinn and Quayyum (2013) extend the analysis in Chinn and Meredith (2004) by a decade and find that the results in the latter are robust; however, the evidence is slightly weaker, potentially because the longer sample includes the zero-lower bound period. In this paper, differently from the contributions listed above, we focus instead on the lack of empirical validity of UIRP in the short run, which still remains a puzzle in the literature, and argue that uncertainty plays a potentially important role in explaining the puzzle.

Our paper is also related to the recent literature that has developed theoretical models to explain the UIRP puzzle. Two possible explanations for the lack of empirical validity of the UIRP are the presence of time-varying risk premia and expectational errors (Lewis, 1995). For example, Fama (1984) attributes the lack of empirical validity of the UIRP to time-varying risk premia. His paper shows that, in order to fit the empirical evidence, the implied risk premia of a country must be negatively correlated with its expected rate of depreciation and have greater variance. However, asset pricing models had not been able to produce risk premia with these properties, hence the term "puzzle". There are several possible theoretical explanations for time-varying risk premia, among which the most recent include Brunnermeier at al. (2009) and Farhi and Gabaix (2016). Brunnermeier et al. (2009) look at currency crashes and carry trades, where traders borrow low-interest-rate currencies and lend high-interest currencies. One of their findings is that higher levels of the VIX and TED spread predict higher future returns on the carry trade, implying larger UIRP violations. Farhi and Gabaix (2016) link time-varying risk premia in currency markets to rare but extreme disasters; since both the probability of these disasters as well as each country's exposure to them is time varying, the model can potentially generate the lack of UIRP, as relatively riskier countries end up with a higher interest rate to compensate investors in case the disaster happens. However, their evidence is limited to a calibration analysis showing that the theoretical predictions of the models are consistent with empirical puzzles (such as UIRP), as opposed to demonstrating empirically the link in the data. The reason is that rare disasters realize sporadically in the data, and thus it is difficult to find empirical evidence in favor of their model.[4] Our empirical results provide potential empirical support in favor of Farhi and Gabaix (2016) in the following sense. An unexpected rare disaster that realizes in the data will increase our uncertainty index; conversely, even a situation where agents expect a rare disaster that does not realize in the data will increase our uncertainty index, as the expectations will be different from the realization. Thus, at times of rare disasters, uncertainty goes up and it is more likely that the UIRP does not hold, while, during normal times, uncertainty decreases and it is more likely that the UIRP holds, consistently with our empirical results. However, our uncertainty index more broadly captures not only rare disasters but also any deviation between agents' expectations of exchange rate fluctuations and their realizations. In addition, our robustness results to using the VIX as a measure of uncertainty are consistent with Brunnermeier et al. (2009).[5]

The third strand is the literature on uncertainty. Several recent papers have analyzed the effects of uncertainty on the macroeconomy; for example, Bloom (2009), among others, has measured uncertainty as the volatility in financial markets. In this paper, we use survey forecasts to measure uncertainty,

[4] Other theoretical explanations of the lack of empirical validity of the UIRP include Colacito and Croce (2011), Verdelhan (2010) and Bacchetta and Van Wincoop (2010). On the one hand, Colacito and Croce (2011) consider long-run risks models as a potential explanation of several exchange rate puzzles, including UIRP, where the long run risk is related to a small predictable component in consumption growth. On the other hand, Verdelhan (2010) shows that habit models with time-varying risk aversion and procyclical real interest rates can also theoretically generate time-varying risk premia in currency markets. However, Verdelhan (2010) shows that the exchange rates series simulated by his calibrated model are too volatile and too much correlated with consumption growth shocks. Similarly, Bacchetta and Van Wincoop (2010) discuss and calibrate a theoretical model that attributes deviations from UIRP to infrequent portfolio decisions.

[5] In unreported results we investigated whether the failure of UIRP is more likely to be caused by expectation errors or by risk premia using Froot and Frankel's (1989) decomposition. The failure seems more likely to derive from expectation error for Switzerland and from risk premia for Canada, Japan and the UK; in the case of Europe, both are equally likely.

similarly to Ozturk and Sheng (2016), who use survey forecasts to measure global and country-specific macroeconomic uncertainty, and Rossi et al. (2016), who use survey density forecasts to understand the sources of macroeconomic uncertainty. However, differently from them, we focus on exchange rate uncertainty. The literature on the relationship between exchange rates and uncertainty is, instead, more limited. Berg and Mark (2016) and Mueller, Tahbaz-Salehi and Vedolin (2016), for example, study the relationship between trading strategies in exchange rate markets and uncertainty. The former study the exposure of carry-trade currency excess returns to global fundamental macroeconomic risk. Their measure of global macroeconomic uncertainty, defined as the cross-country high-minus-low conditional skewness of the unemployment gap, is a factor priced in currency excess returns. Mueller et al. (2016) instead study whether trading strategies of going short on one currency and long on other currencies exhibits significantly larger excess returns on FOMC announcement days, and find that the excess returns are higher the higher is uncertainty about monetary policy. Menkoff et al. (2012) propose a new risk factor capable of explaining the cross section of excess returns: the global foreign exchange volatility risk; they find that high interest rate currencies are negatively related to global foreign exchange volatility, and thus deliver low returns when volatility is unexpectedly high, at times when low interest rate currencies provide positive returns. Belke and Kronen (2015) analyze the role of uncertainty in explaining exchange rate bands of inaction and their effects on exports. Similarly to these contributions, our paper also studies the effects of uncertainty in exchange rate markets, but focuses instead on explaining the UIRP puzzle, as opposed to explaining larger excess returns in cross section carry-trade strategies or fluctuations in exports.

This paper is organized as follows. The next section describes the data used in this study and Section 3 discusses the exchange rate uncertainty index that we use. Section 4 revisits the empirical evidence on UIRP in our sample, while Section 5 investigates whether deviations from UIRP can be explained by uncertainty. Section 6 performs robustness analyses using other uncertainty indices, while Section 7 discusses results for a larger set of countries using uncertainty indices based on random walk forecast errors.

Section 8 concludes.

2 The Data

We collect monthly data spanning 1993:M11 to 2015:M1 on exchange rates, three-month Euro LIBOR rates, and the uncertainty measure(s). In our benchmark results, we focus on industrialized countries, and consider five currency pairs: the Swiss franc, the Canadian dollar, the British pound, the Japanese yen, and the Euro against the US dollar. We focus on exchange rates for industrialized countries for which the survey expectations necessary to construct our uncertainty index are available. Robustness results for additional countries are discussed in Section 7. The period has been chosen based on the availability of the uncertainty index. In fact, the data on our uncertainty measure start in 1993:M11 and end in 2015:M1 for all currencies except the Euro (for the Euro it begins on 2001:M7) – see below for more details on the uncertainty measure. The data on the exchange rates for the five currency pairs are from WM/Reuters. The exchange rates are values of the national currencies relative to one US dollar. For the interest rates we collect monthly data on three-month Euro LIBOR rates for the respective five countries and the United States. The data are from the Financial Times. All data have been collected via Datastream. More details (including mnemonics) are provided in Table 1, which also includes a description of the additional data we use in the robustness analysis to the larger set of countries.

INSERT TABLE 1 HERE

3 The Exchange Rate Uncertainty Index

Regarding uncertainty, several methodologies and strategies to construct uncertainty indices are available. Bloom (2009) proposes to measure macroeconomic uncertainty using the volatility in stock prices, while Baker et al. (2016) propose a measure of macroeconomic policy uncertainty. Since we are interested in exchange rate uncertainty, their measures are not the most appropriate. Jurado et al. (2015) and Ludvigson et al. (2015) propose to measure uncertainty as the time-varying volatility of forecast errors in predicting macroeconomic and financial variables, while Scotti (2016) measures uncertainty as macroeconomic news announcements. The uncertainty series that we construct are similar in spirit to Jurado et al. (2015) but they are obtained using the methodology in Rossi and Sekhposyan (2015). Rossi and Sekhposyan's (2015) uncertainty index is constructed by comparing the realized forecast error of the target variable with the unconditional forecast error distribution of the same variable. The intuition is that, if the observed realization of the forecast error is in the tails of the distribution, then the realization was very difficult to predict; thus, such an environment is deemed very uncertain. One of the advantages of the Rossi and Sekhposyan (2015) index is that it allows for asymmetry: in other words, it can separately distinguish between uncertainty due to unexpectedly high and low exchange rates – an important feature that is not shared by uncertainty indices based on the volatility of forecast errors.[6]

We construct the exchange rate uncertainty index based on fixed-horizon forecast errors from surveys conducted by Consensus Economics.[7] The uncertainty index is monthly and the forecast horizon is three months; therefore, the interest rate differential is based on three-month interest rates. Let the bilateral nominal exchange rate between a country and the US at time t be denoted by S_t and let $s_t = \ln(S_t)$. Furthermore, let the $h-$step-ahead forecast error for the rate of growth of the exchange rate between time t and time $t+h$ be denoted by $e_{t+h} = (s_{t+h} - s_t) - E_t(s_{t+h} - s_t)$, and its unconditional forecast error distribution be denoted by $p(e)$. Rossi and Sekhposyan's (2015) index is based on the cumulative density of forecast errors evaluated at the realized forecast error, e_{t+h}: $U_{t+h} = \int_{-\infty}^{e_{t+h}} p(e)\, de$. A large value of the index indicates a realization of the exchange rate that is very different from the expected value. In particular, a realized value much bigger (smaller) than the expected value, which is 0.5, measures a positive

[6]We perform a robustness analysis to using alternative uncertainty indices in Section 6.

[7]We use the average forecasts from a sample of approximately 250 professional forecasters.

(negative) “shock”. The overall exchange rate uncertainty index that does not distinguish between positive and negative “shocks” is:[8]

$$U_{t+h}^{*} = \frac{1}{2} + \left| U_{t+h} - \frac{1}{2} \right| .$$

Values of U_{t+h}^{*} close to unity indicate high uncertainty, while values close to 0.5 indicate low uncertainty.

Figure 1 plots the exchange rate uncertainty indices for the countries in our sample. The time series fluctuations of the uncertainty indices are consistent with several events that affected these countries over time. For example, focusing on the EU, the two periods of high uncertainty during the latest financial crisis are clearly visible; they are related to the two recent recessions in the Euro-area: the first from 2008:Q1 to 2009:Q2 and the second from 2011:Q3 to 2013:Q1. In particular, the Euro debt crisis shows up as an upward trend in uncertainty in the EU since mid-2011. A similar pattern affects the UK during the same period. Note also the upward trend in uncertainty visible in Canada during the recent US financial crisis starting in 2007. Finally, another notable event taking place in 2006 is Bank of Japan raising interest rates for the first time in several years, which might have caused the drastic increase in uncertainty around mid-2006.

INSERT FIGURE 1 HERE

4 Revisiting Uncovered Interest Rate Parity

Uncovered interest rate parity (UIRP) states that, in a world of perfect foresight and a nominal bilateral exchange rate S_t, investors can buy $1/S_t$ units of foreign bonds using one unit of the home currency, where S_t denotes the price of foreign currency in terms of home currency. Suppose the foreign bond pays one unit plus the foreign interest rate between time t and $(t+h)$, i_{t+h}^{*}, where h is the horizon of the investment. At the end of the period, the foreign return can be converted back into the home

[8] A Not-for-Publication Appendix investigates the effects of asymmetries in uncertainty.

currency with a value of $S_{t+h}\left[\left(1+i^*_{t+h}\right)/S_t\right]$ in expectation. In the absence of transaction costs, by no-arbitrage this return must be in expectation equal to the return of the home bond, $(1+i_{t+h})$. Therefore, $\left(1+i^*_{t+h}\right)E_t\left(S_{t+h}/S_t\right)=(1+i_{t+h})$, where $E_t(.)$ denotes the expectation at time t. By taking logarithms and ignoring Jensen's inequality, the uncovered interest rate parity equation follows directly:

$$E_t\left(s_{t+h}-s_t\right)=\alpha+\beta\left(i_{t+h}-i^*_{t+h}\right), \tag{1}$$

where the UIRP parameters α and β have the theoretical values: $\alpha=0$ and $\beta=1$.

Overall, the empirical evidence is not favorable to UIRP – see Rossi (2013) for a recent survey. It is well-known that the constant, α, is different from zero, and the slope, β, is either negative or close to zero, or sometimes positive and very large in magnitude. Similarly, the empirical evidence is equally not supportive of UIRP in out-of-sample forecast evaluation; in fact, it is also well-known, since the early work by Meese and Rogoff (1983a,b; 1988), that eq. (1) does not forecast exchange rates out-of-sample better than the random walk. The same result was reinforced by Cheung et al., (2005), Alquist and Chinn (2008) and Chinn and Quayyum (2013). Slightly more positive findings have been reported by Clark and West (2006) at short-horizons; however, as Rossi (2013) pointed out, the reason for the positive findings in Clark and West (2006) are mainly due to the use of an alternative test of predictive ability.[9]

We start by confirming the existing findings in the literature, namely that UIRP does not hold in the data. Panel A in Table 2 estimates regression (1) in our sample, and shows that, for several countries, β is very small, and in the case of Switzerland, Canada and Japan, it is negative and statistically significantly different from one. Only for the EU and the UK the slope is positive and statistically indistinguishable from its theoretical value under the UIRP. The constant instead is small and insignificantly different from zero for most countries.[10]

[9] One could potentially consider forecasting real exchange rates using real interest rates; however, the survey forecasts are for the nominal, not the real, exchange rate – which nevertheless is what is considered in the aforementioned literature.

[10] The 95% confidence intervals reported in parentheses in this paper are based on a Newey and West (1987) HAC estimator

Our results are similar to those in the literature, except that our estimates are slightly smaller than those reported in the earlier literature. For instance, Chinn and Quayyum (2013) use quarterly data spanning 1975:Q1-2011:Q4 for the same set of currency pairs, and they find slope estimates ranging from -1.85 to -2.25 with the exception of the Canadian dollar, whose slope is -0.17. However, a detailed analysis reveals that the large negative values are driven by sample selection. Firstly, the rolling-window estimates which we report later in the paper show that the slope coefficients have been increasing over time: our sample is shorter than, e.g., Chinn and Quayyum (2013), and in particular it omits the Seventies and the Eighties; the latter are decades with large deviations from UIRP according to Lothian and Wu (2011).[11] Secondly, if we consider the sample up to 2011:M10, that is, omitting the last 4 years to better match the sample used in Chinn and Quayyum (2013), the estimates become negative for four countries out of five and the negative coefficients are larger in magnitude in absolute value (see Table 2, Panel B).

A comparison of the results in the two panels in Table 2 also points out another important empirical feature of UIRP: the well-known fact that the UIRP parameters are unstable over time. For example, note how the slope coefficient for the Euro data turns from positive to negative depending on the sample, and how its magnitude varies in Japanese data. Rossi (2006) investigated the instability of the parameters in exchange rate monetary models (that is, models that explain exchange rate fluctuations using output, money and interest rate differentials) and found ample evidence of instabilities based on conventional tests of parameter instability. Furthermore, she argued that the empirical rejections of the monetary exchange rate model could be due to parameter instabilities; in fact, by using alternative and more powerful tests that evaluate Granger-causality robust to instabilities, she found that monetary models' predictors helped forecasting exchange rates *at some point in time*. However, she did not consider the UIRP in her analysis, so it is important to investigate whether UIRP fails in the data regardless of the presence of instabilities

for the covariance matrix, using a truncation lag equal to two.

[11] Our sample is shorter since it is determined by the availability of the uncertainty index.

in the data, a question we explore in the rest of this section.

We first investigate the stability of the UIRP parameters over time by plotting their estimates in rolling windows over ten years of data in the top panel in Figures 2(a-e). The figures confirm the presence of instabilities throughout the sample that we consider. For Canada, the value of the constant is small throughout the sample, but the slope value changes significantly from negative to positive. The slope changes drastically for the EU as well, ranging from values close to zero at the beginning of the sample to almost four towards the end of the sample. In the case of Japan, the coefficient is close to zero for almost all the sample except the beginning and the end. Switzerland and the UK are two other countries where the slope changes drastically from negative to large and positive values. For the latter country, the constant also is very unstable, taking both positive and negative values depending on the sample period.

We investigate more formally whether instabilities affect UIRP in Table 3(a-c). We consider the following regression:

$$E_t\left(s_{t+h}-s_t\right)=\alpha_t+\beta_t\left(i_{t+h}-i^*_{t+h}\right), \tag{2}$$

where the constant, or the slope parameter, or potentially both, might be time-varying. Absence of time variation manifests itself in constant parameters, that is: $\alpha_t=\alpha$ and/or $\beta_t=\beta$. We test parameter stability using a battery of tests, including Andrews' (1993) Quandt Likelihood Ratio test (QLR), Andrews and Ploberger's (1994) Exponential-Wald (Exp-W), as well as Nyblom's (1989) test. The tests differ depending on the type of instability they allow for; in particular, Andrews (1993) and Andrews and Ploberger (1994) allow for a one-time structural change, while Nyblom (1989) considers smoother and more frequent changes.

Table 3(a) reports results for testing the joint stability in both the constant and the slope parameters. It is clear that the stability is overwhelmingly rejected, with p-values that are zero in all cases. We then investigate whether the instability is more pronounced in the constant or in the slope. Table 3(b) reports tests of stability on the constant. The table shows that the constant is unstable for most countries except the UK. Table 3(c) reports tests of stability on the slope; the table shows that the slope is unstable for all countries, including the UK.

Since the parameters are time-varying, the UIRP tests presented in Table 2 are invalid, as they assume stability in the parameters. Therefore, we complement the analysis with tests that are robust to parameter instabilities. In particular, we implement the Exp-W*, Mean-W*, Nyblom* and QLR* tests proposed by Rossi (2005), which are valid to test the UIRP conditions that $\alpha_t = 0$ and $\beta_t = 1$ even in the presence of time-variation in the parameters.[12] Tables 4(a-c) show that the results in Table 2 are robust. In particular, Table 4(a) shows that the both parameters are significantly different from the values predicted by the UIRP; Tables 4(b-c) report results for the constant and the slope separately, and show that the rejections are mostly due to the fact that the slope is different from unity, especially for Canada, the UK and Japan.[13]

The analysis in this section shows that the coefficients estimated in UIRP regressions are very unstable over time and that UIRP does not hold in the data, regardless of the presence of instabilities. However, the analysis does not shed light on why there are time-varying deviations from UIRP. The next section will tackle this important question.

INSERT TABLES 2, 3, 4 AND FIGURE 2 HERE

[12] The difference among the Exp-W*, Mean-W*, QLR* and Nyblom* tests is, again, that they focus on different types of instabilities. In particular, the first three focus on the case of a one-time structural change while Nyblom* allows smoother and more frequent changes.

[13] Note that, in Table 4(b), the Exp-W* test does not reject for some countries while the Mean-W*, Nyblom* and QLR* tests reject. The reason why the tests disagree is because they consider different types of instabilities: the Nyblom* test, for example, has more power when parameters are smoothly time-varying.

5 Can Uncertainty Explain UIRP Deviations?

The previous section has confirmed the existence of two important puzzles in the empirical literature in international finance: *UIRP coefficients are both different from their theoretical values and unstable over time*. This paper tries to offer an explanation to *both* these puzzles by arguing that uncertainty is one of the reasons explaining the empirical invalidity of the UIRP; that the coefficients in UIRP regressions are more likely to be close to the values predicted by UIRP in times when uncertainty is low; and that their time variation is, at least partly, due to the fact that UIRP holds when uncertainty is low but does not when uncertainty is high.

As discussed in the introduction, a typical explanation for the UIRP puzzle is the existence of time-varying risk premia; but what generates time-varying risk premia? The most recent theoretical explanations include rare disasters (Farhi and Gabaix, 2016, and Brunnermeier et al., 2009), habits (Verdelhan, 2010) or long run risks related to a small predictable component in consumption growth (Colacito and Croce, 2011). Our empirical results provide potential empirical support in favor of Farhi and Gabaix (2016) in the following sense. An unexpected rare disaster that realizes in the data increases our uncertainty index; conversely, even a situation where agents expect a rare disaster and it does not realize in the data will show up as an increase in our uncertainty index, as the expectations will be different from the realization. Thus, at times of rare disasters, uncertainty goes up and it is more unlikely that the UIRP does not hold, while, during normal times, uncertainty decreases and it is more likely that UIRP holds, consistently with our empirical results. However, our uncertainty index includes not only rare disasters but also any deviation between agents' expectations of exchange rate fluctuations and their realizations.

A visual analysis of the relationship between uncertainty and the rolling estimates of the UIRP parameters is presented in Figure 2. The top panels in Figure 2 show the rolling estimates of the parameters while the bottom panels display the uncertainty index for each country; the bottom panels plot the exchange rate uncertainty index, U^*_{t+h}. The figure shows that there is correlation between uncertainty and the UIRP coefficients for most countries: when uncertainty is substantially high, there are more deviations from UIRP, both in terms of deviations of α from zero as well as deviations of β from unity. For example, the case of Switzerland (depicted in Figure 2d) is emblematic: the negative values of the slope and the constant are clearly visible at the beginning of the sample, and that is also when uncertainty is the highest. Similarly, in the case of the UK and Canada (depicted in Figures 2e and 2a, respectively), the slope approaches unity around 2005-2008, which is exactly when uncertainty is the lowest, and very different from unity both at the beginning (when the slope is negative) and towards the end of the sample (when the slope is positive and large), when uncertainty is the highest. For the EU, depicted in Figure 2(b), uncertainty is high for most of the sample we consider. Finally, in the case of Japan (depicted in Figure 2c) too, both the slope and the intercept are negative at the beginning of the sample, when the uncertainty is often at high levels.

To investigate more formally whether uncertainty can explain the UIRP puzzle, we estimate the following regression:

$$E_t\left(s_{t+h}-s_t\right)=\alpha_1\left(1-d_t\right)+\beta_1\left(1-d_t\right)\left(i_{t+h}-i^*_{t+h}\right)+\alpha_2 d_t+\beta_2 d_t\left(i_{t+h}-i^*_{t+h}\right), \qquad (3)$$

where d_t is a dummy variable equal to one if the uncertainty is exceptionally high. Since the uncertainty indices are quite volatile, we smooth them using the same rolling window that we used to estimate the parameters in the UIRP regression, equal to ten years of data. Time periods of high uncertainty are identified by situations in which uncertainty (U^*_{t+h}) is in the upper quartile of its distribution, i.e. we identify high uncertainty periods with sub-samples with the 25% highest values of uncertainty.

Table 5 reports the estimates of eq. (3). The table shows that the empirical evidence in favor of UIRP is weakest in periods where uncertainty is exceptionally high, and substantially stronger in periods where uncertainty is around normal values. More in detail, we note that, in the case of Switzerland, both values of α_2 and β_2 are negative and large in absolute value; since α_2 and β_2 are the constant and slope of the UIRP in periods of high uncertainty, the regression results confirm the existence of large deviations from UIRP when uncertainty is exceptionally high. However, in periods of low uncertainty, both α_1 and β_1 are closer to their theoretical values, and insignificantly different from them. Japan is another case where the slope switches from negative values (and significantly different from unity) during periods of high uncertainty, to positive values close to unity (and statistically insignificantly different from unity). In Canada, again, the slope is negative and close to zero in periods of high uncertainty, while it becomes positive and closer to unity in periods of low uncertainty; the constant also gets closer to its theoretical value of zero in periods of low uncertainty. In the case of the EU and the UK, the uncertainty state also drives the slope coefficient closer to its theoretical value; in all cases, the point estimates are more precisely estimated in periods of low uncertainty.

Note that our results have direct implications for the risk premium. In fact, let $R_{t+h,t} \equiv (s_{t+h} - s_t) - i_{t+h} - i^*_{t+h}$ denote the risk premium. The regression: $E_t R_{t+h} = \alpha_1 (1 - d_t) + \beta_1 (1 - d_t) \left(i_{t+h} - i^*_{t+h}\right) + \alpha_2 d_t + \beta_2 d_t \left(i_{t+h} - i^*_{t+h}\right)$ yields exactly the same coefficients α_1 and α_2 (and their confidence intervals) as the regression in eq. (3), and the slope coefficients β_1 and β_2 are exactly the same as the estimated slope coefficients we report in eq. (3) minus one (and similarly for their confidence intervals). Thus, the results in eq. (3) directly tell us that risk premia are more correlated with interest rate differentials during periods of high uncertainty than during low uncertainty, and significantly so for Switzerland and Japan. Notice that risk premia are never significantly correlated to interest rate differentials during periods of low uncertainty for any of the countries.

Finally, we investigate whether uncertainty can help explaining UIRP deviations directly by estimating the following regression:

$$E_t \left(s_{t+h} - s_t\right) = \alpha + \beta \left(i_{t+h} - i^*_{t+h}\right) + \gamma U^*_{t+h}, \tag{4}$$

and testing whether γ is significantly different from zero using the tests robust to instabilities. The results are reported in Table 6. Indeed, the table shows that uncertainty does significantly help in explaining deviations from UIRP for all countries.

It is interesting to investigate whether time-variation in the UIRP can be explained by differences in monetary policy alone. Table 7 estimates the UIRP in the sub-sample of the zero-lower bound in the US (December 2008 to December 2014), a time period where the interest rate was close to zero and, hence, the traditional monetary policy prescription of lowering interest rates in the presence of the recession was infeasible. By comparing Table 7 with Table 2 it is clear that, although for Switzerland and the EU the estimates of UIRP coefficients during the zero lower bound period are closer to their theoretical value than during the full sample, the same result does not hold for Canada, Japan and the UK.

INSERT TABLES 5, 6, 7

6 The Effects of Global Uncertainty

In the previous sections, we focused attention on indices that measure uncertainty in bilateral exchange rates, which is a relevant measure for our purposes since it proxies exchange rate uncertainty in financial markets. The uncertainty index we used was based on Rossi and Sekhposyan's (2015) methodology, whose advantage is that it can be easily tailored to measure uncertainty in any variable subject to the minimal requirement of availability of time series of forecast errors. Given the bilateral nature of the exchange rate data we used, the indices may include both global as well as country-specific idiosyncratic uncertainty. But which one is more relevant for explaining deviations from UIRP: global uncertainty or country-specific idiosyncratic uncertainty in financial markets? We attempt to answer this question in this section.

We construct an index of global uncertainty in financial markets by taking the common component of the Rossi and Sekhposyan (2015) uncertainty indices for the currency pairs we consider in Section 3,[14]

[14] The common component is measured by the first principal component estimated with a factor model from all the bilateral exchange rate uncertainty indices.

which captures global uncertainty in exchange rate financial markets, cleaned from any idiosyncratic or country-specific component. There are also many other uncertainty indices available in the literature that one could alternatively use, such as: the VIX (Bloom, 2009); the Jurado et al. (2015) macroeconomic uncertainty index; the Ludvigson et al. (2015) financial uncertainty index; and the Baker et al. (2016) economic policy uncertainty index. These alternative uncertainty indices are available mainly for the U.S. and can be thought of as a measure of global macroeconomic and/or political uncertainty given the prominent role of the U.S. on the international scene. We also consider the Menkoff et al. (2012) global foreign exhange volatility risk measure. Figure 3 depicts all the global uncertainty indices – they are very correlated in the sample we focus on.

We estimate eq. 3 using each one of these indices as a measure of global uncertainty in exchange rates, the macroeconomy or financial markets. The results are reported in Table 8(A-F). For all countries, in the case of the VIX, the Jurado et al. (2015) and the Ludvigson et al. (2015) uncertainty indices, the estimate of the slope coefficient on the interest rate differential gets closer to the theoretical value of unity during periods of low uncertainty while the coefficient can be quite different from its theoretical value in periods of high uncertainty.[15] So, in most cases, what matters is the global uncertainty. Results are similar for the Menkoff et al. (2012) global foreign exchange volatility risk measure. The only exception is the Baker et al. (2016) measure for the case of Japan; the index predicts a negative slope for Japan during the periods of low uncertainty and a positive slope when uncertainty is high; however, the Baker et al. (2016) index captures economic policy uncertainty in the US, which contains information above and beyond global uncertainty in financial markets, including market reforms etc., and in some cases relevant only for US internal purposes, and thus may have little power to explain the UIRP in a country like Japan.

[15] The standard errors are quite large in periods of high uncertainty; so the confidence intervals typically contain the theoretical value of unity even in periods of high uncertainty, although the point estimate is typically further away from its theoretical value.

By comparing Panel E in Table 8 (where we use the principal component from our cross-section of bilateral exchange rate uncertainty indices) and Table 5 (where we use our country-specific bilateral exchange rate uncertainty index), we note that the principal component is not as effective in explaining time-varying UIRP deviations as the country-specific uncertainty indices. Thus, not only global shocks in international financial markets are important, but also country-specific idiosyncratic uncertainty shocks.

INSERT TABLE 8 AND FIGURE 3 HERE

7 Exploring A Larger Set of Countries

The exchange rate uncertainty index described in Section 3 is based on survey forecast errors. On the one hand, using survey forecasts is desirable since it ensures that, if one is willing to make the realistic assumption that forecasters use all the available information when making their forecasts (including soft information from news), then the largest possible information set is used when constructing forecast errors; in addition, the forecasts do not depend on any specific theoretical model of exchange rate fluctuations. On the other hand, survey-based uncertainty indices have the disadvantage that they can only be constructed when survey forecasts are available, which may substantially limit the set of countries that a researcher can analyze. However, if a researcher is interested in measuring uncertainty in countries where survey forecast errors are not available, it is still possible to construct an uncertainty index based on models' forecasts.

In this section, we construct exchange rate uncertainty indices based on random walk forecasts. Since Meese and Rogoff (1983a,b), the random walk model has been considered the best benchmark when forecasting exchange rates (Rossi 2013), and hence it is a good candidate for generating the uncertainty index. The random walk model sets $E(s_{t+h} - s_t) = 0$; the forecast errors, $s_{t+h} - s_t$, can then be used to construct the uncertainty index U^*_{t+h} as in Section 3. We calculate the overall uncertainty index and study UIRP in times of high and low uncertainty.

We start by considering the same set of countries that we considered in Section 5 to verify the robustness of the results. The results, reported in Table 9, support the main findings in Section 5: the empirical evidence in favor of UIRP is weakest in periods where uncertainty is exceptionally high, and substantially stronger in periods where uncertainty is around normal values. For instance, the coefficient on the interest rate differential is positive and closer to unity when uncertainty is low for Switzerland, Canada and Japan, while it is negative or zero when uncertainty is high. In periods of low uncertainty, the slope coefficients of all countries get closer to their theoretical value (equal to one) relative to periods of high uncertainty.

We then extend our results to other countries for which survey forecasts and/or other uncertainty indices are not available. In particular, we extend our dataset to include Australia, Sweden, South Africa, Norway, New Zealand and Denmark; as before, the bilateral exchange rates are against the US dollar. This subset of countries includes both commodity and non-commodity currencies, both emerging and developed markets, and currencies of various degrees of historical volatility.

Firstly, Panel A in Table 10 revisits the empirical evidence for the UIRP relationship for these counties in the full sample. For all countries the point estimate of the coefficient on the interest rate differential is far from one, and for all countries except New Zealand we reject that it equals unity. In other words, the UIRP is violated for this set of countries as well.

We then calculate the uncertainty measure based on random walk forecast errors to investigate whether high uncertainty can explain the deviations from the UIRP. The results are reported in Table 10, Panels B-C. For all countries except Norway, the estimate of the slope in periods of low uncertainty is closer to the theoretical value than when uncertainty is high. Thus, the UIRP puzzle is alleviated in low uncertainty environments for several of the additional countries that the extension to random walk forecast errors allows us to consider (Australia, Sweden and Denmark). For some other countries, although low uncertainty typically moves the coefficient in the right direction, it does not fully resolve the puzzle (South Africa and New Zealand); however, the latter (and Norway, for which the puzzle is not resolved) are "commodity countries", for which commodity prices might play a role in determining exchange rate fluctuations, which we abstract from.

INSERT TABLE 10 HERE

8 Conclusions

This paper has investigated whether uncertainty can explain the short-run deviations from UIRP that we empirically observe in the data. We have found that deviations from UIRP are stronger in periods of high uncertainty, while UIRP tends to hold in periods of low uncertainty. While it is well-known that deviations from UIRP are large and time-varying, this is the first paper that provides an economic rationale for both the UIRP puzzle and the presence of time variation in UIRP coefficient estimates by linking UIRP deviations to uncertainty. The result is robust to using various measures of economic uncertainty as well as uncertainty indices based on random walk forecasts. Our empirical results are consistent with the existence of time-varying risk premia potentially linked to rare disasters.

Additional analyses that could be carried out in the future include investigating whether similar results hold at long horizons; however, the UIRP puzzle is really a puzzle at short horizons, which is what we focused on in this paper.

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Tables

Table 1. Data Description

Country	Period	Code	Description
		Exchange rates:	
Switzerland	1994M1:2015M1	SWISSF$	SWISS FRANC TO US $ - EXCH. RATE
Canada	1994M1:2015M1	CNDOLL$	CANADIAN $ TO US $ - EXCH. RATE
United Kingdom	1993M11:2015M1	UKDOLLR	UK £ TO US $ - EXCH. RATE
Japan	1993M11:2015M1	JAPAYE$	JAPANESE YEN TO US $ - EXCH. RATE
EU	2001M7:2015M1	EUDOLLR	EURO TO US $ - EXCH. RATE
South Africa	1997M4:2016M10	COMRAN$	SOUTH AFRICA RAND TO US $ - EXCH. RATE
Australia	1997M4:2016M10	AUSTDOI	AUSTRALIAN $ TO US $ - EXCH. RATE
Norway	1997M4:2016M10	NORKRO$	NORWEGIAN KRONE TO US $ - EXCH. RATE
Sweden	1997M4:2016M10	SWEKRO$	SWEDISH KRONA TO US $ - EXCH. RATE
Denmark	1997M4:2016M10	DANISH$	DANISH KRONE TO US $ - EXCH. RATE
New Zealand	1997M4:2016M10	NZDOLLI	NEW ZEALAND $ TO US $ - EXCH. RATE
		Interest rates:	
Switzerland	1993M11:2015M1	ECSWF3M	Euro LIBOR 3-month rate, middle rate
Canada	1993M11:2015M1	ECCAD3M	Euro LIBOR 3-month rate, middle rate
United Kingdom	1993M11:2015M1	ECUKP3M	Euro LIBOR 3-month rate, middle rate
Japan	1993M11:2015M1	ECJAP3M	Euro LIBOR 3-month rate, middle rate
EU	2001M7:2015M1	ECEUR3M	Euro LIBOR 3-month rate, middle rate
United States	1993M11:2015M1	ECUSD3M	Euro LIBOR 3-month rate, middle rate
South Africa	1997M4:2016M10	ECSAR3M	Euro LIBOR 3-month rate, middle rate
Australia	1997M4:2016M10	ECAUD3M	Euro LIBOR 3-month rate, middle rate
Norway	1997M4:2016M10	ECNOR3M	Euro LIBOR 3-month rate, middle rate
Sweden	1997M4:2016M10	ECSWE3M	Euro LIBOR 3-month rate, middle rate
Denmark	1997M4:2016M10	ECDKN3M	Euro LIBOR 3-month rate, middle rate
New Zealand	1997M4:2016M10	ECNZD3M	Euro LIBOR 3-month rate, middle rate

Note to Table 1. The table reports mnemonics and descriptions for our data. All interest rates are "middle rates". All exchange rate data are from WM/Reuters, while all interest rate data are from FT/Reuters.

Table 2. Traditional UIRP Regressions

	Panel A. Full Sample		Panel B. Sub-sample ending in 2011	
Country:	α	β	α	β
Switzerland	-0.01	-0.59	-0.023	-0.817
	(-0.028;-0.002)	(-1.09;-0.100)	(-0.039;-0.007)	(-1.382;-0.252)
EU	-0.007	0.391	-0.004	-0.351
	(-0.016;0.002)	(-0.576;1.358)	(-0.016;0.007)	(-1.178;0.476)
Canada	-0.001	-0.196	-0.003	-0.383
	(-0.007;0.004)	(-0.706;0.312)	(-0.010;0.003)	(-0.906;0.140)
UK	-0.004	0.378	-0.005	0.410
	(-0.012;0.004)	(-0.513;1.271)	(-0.014;0.004)	(-0.502;1.324)
Japan	-0.002	-0.118	-0.023	-0.585
	(-0.015;0.011)	(-0.533;0.296)	(-0.036;-0.010)	(-0.988;-0.181)

Note to the table. The table reports estimates of UIRP regressions (and 95% confidence intervals in parentheses) in the full sample as well as a sub-sample ending in 2011.

Table 3(a). Instability Tests: Joint Test on α and β

Country:		QLR	Exp-W	Nyblom
Switzerland	Test statistic	39.08	15.11	3.55
	P-value	0	0	0
EU	Test statistic	35.69	13.98	3.27
	P-value	0	0	0
Canada	Test statistic	24.54	9.44	2.44
	P-value	0	0	0
UK	Test statistic	50.09	19.9	1.98
	P-value	0	0	0
Japan	Test statistic	44.52	18.1	4.50
	P-value	0	0	0

Note to the table. The table reports joint tests of parameter instabilities on the two UIRP regression coefficients.

Table 3(b). Instability Tests: Test on the Constant (α)

Country:		QLR	Exp-W	Nyblom
Switzerland	Test statistic	23.73	7.377	0.671
	P-value	0	0	0.151
EU	Test statistic	34.06	13.52	1.978
	P-value	0	0	0
Canada	Test statistic	16.40	4.763	0.862
	P-value	0	0	0.076
UK	Test statistic	3.150	0.544	0.171
	P-value	0.809	0.828	0.847
Japan	Test statistic	51.40	21.00	1.577
	P-value	0	0	0

Note to the table. The table reports tests of parameter instabilities on the constant coefficient in the UIRP regressions.

Table 3(c). Instability Tests: Test on the Slope (β)

Country:		QLR	Exp-W	Nyblom
Switzerland	Test statistic	26.81	8.74	1.746
	P-value	0	0	0
EU	Test statistic	45.34	18.88	3.459
	P-value	0	0	0
Canada	Test statistic	27.28	10.68	2.416
	P-value	0	0	0
UK	Test statistic	26.44	8.54	1.06
	P-value	0	0	0.036
Japan	Test statistic	26.66	8.92	1.176
	P-value	0	0	0.023

Note to the table. The table reports tests of parameter instabilities on the slope coefficient in the UIRP regressions.

Table 4(a). Granger-causality Tests: Joint Test on α and β

Country:		Exp-W*	Mean-W*	Nyblom*	QLR*
Switzerland	Test statistic	68.91	121.76	31.93	146.45
	P-value	0	0	0	0
EU	Test statistic	23.09	26.052	6.023	54.09
	P-value	0	0	0	0
Canada	Test statistic	57.23	89.393	16.28	120.36
	P-value	0	0	0	0
UK	Test statistic	44.90	48.059	8.28	98.60
	P-value	0	0	0	0
Japan	Test statistic	77.34	129.908	31.9604	163.7371
	P-value	0	0	0	0

Note to the table. The table reports tests of UIRP robust to parameter instabilities. The tests are performed jointly on both the constant and the slope in the UIRP regressions.

Table 4(b). Granger-causality Tests: Test on the Constant (α)

Country:		Exp-W*	Mean-W*	Nyblom*	QLR*
Switzerland	Test statistic	11.19	16.55	3.52	29.48
	P-value	0	0	0.021	0
EU	Test statistic	11.32	13.16	3.02	29.36
	P-value	0	0	0.034	0
Canada	Test statistic	3.948	4.095	0.677	14.143
	P-value	0.123	0.404	0.550	0.051
UK	Test statistic	1.117	1.810	0.938	4.437
	P-value	0.822	0.847	0.404	0.820
Japan	Test statistic	17.31	6.837	1.749	43.652
	P-value	0	0.121	0.153	0

Note to the table. The table reports tests of UIRP robust to parameter instabilities. The tests are performed on the constant coefficient in the UIRP regressions.

Table 4(c). Granger-causality Tests: Test on the Slope (β)

Country:		Exp-W*	Mean-W*	Nyblom*	QLR*
Switzerland	Test statistic	50.58	85.23	36.43	110.4
	P-value	0	0	0	0
EU	Test statistic	19.34	18.55	3.33	46.76
	P-value	0	0	0.02	0
Canada	Test statistic	55.25	86.58	16.13	115.6
	P-value	0	0	0	0
UK	Test statistic	19.81	18.93	4.45	48.26
	P-value	0	0	0	0
Japan	Test statistic	42.5	60.26	24.18	93.75
	P-value	0	0	0	0

Note to the table. The table reports tests of UIRP robust to parameter instabilities. The tests are performed on the slope coefficient in the UIRP regressions.

Table 5: UIRP and Exchange Rate Uncertainty

	Low Uncertainty		High Uncertainty	
Country	α_1	β_1	α_2	β_2
Switzerland	0.001	0.469	-0.034	-9.389
	(-0.017;0.019)	(-0.274;1.213)	(-0.074;0.006)	(-19.342;0.564)
EU	-0.001	1.918	-0.012	3.445
	(-0.015;0.013)	(0.188;3.649)	(-0.081;0.056)	(-3.518;10.407)
Canada	-0.005	1.632	-0.009	-0.114
	(-0.015;0.005)	(0.525;2.738)	(-0.041;0.024)	(-4.606;4.379)
UK	-0.007	0.332	-0.033	6.951
	(-0.017;0.003)	(-0.485;1.150)	(-0.067;0.000)	(4.754;9.147)
Japan	0.009	0.739	-0.002	-0.331
	(-0.007;0.025)	(0.089;1.390)	(-0.030;0.026)	(-1.186;0.523)

Note to the table. The table reports parameter estimates in eq. (3), where the measure of uncertainty is overall exchange rate uncertainty (95% confidence intervals in parentheses).

Table 6: Does Uncertainty Granger-cause Exchange Rates?

Country:		Exp-W*	Mean-W*	Nyblom*	QLR*
Switzerland	Test statistic	68.91	121.7	31.93	146.4
	P-value	0	0	0	0
EU	Test statistic	23.09	26.05	6.02	54.09
	P-value	0	0	0	0
Canada	Test statistic	57.23	89.39	16.28	120.3
	P-value	0	0	0	0
UK	Test statistic	44.90	48.05	8.28	98.60
	P-value	0	0	0	0
Japan	Test statistic	77.34	129.9	31.96	163.7
	P-value	0	0	0	0

Note to the table. The table reports results for test statistics robust to parameter instabilities. The statistics test whether uncertainty is a significant predictor in UIRP regressions in eq. (4).

Table 7: UIRP Regressions during the Zero-Lower Bound

Country	α_1	β_1
Switzerland	-0.004	1.047
	(-0.026;0.018)	(-3.206;5.299)
EU	-0.002	1.684
	(-0.019;0.014)	(-0.54;3.909)
Canada	-0.025	3.555
	(-0.054;0.004)	(-0.246;7.355)
UK	-0.004	0.000
	(-0.027;0.019)	(-5.01;5.01)
Japan	0.007	-2.003
	(-0.009;0.023)	(-5.427;1.42)

Note to the table. The table reports estimates (and 95% confidence intervals in parentheses) of UIRP regressions in the zero lower bound sub-sample for the US, estimated to last between December 2008 and December 2014.

Table 8 (Panel A): UIRP and the VIX

Country	Low Uncertainty α_1	β_1	High Uncertainty α_2	β_2
Switzerland	0.011	0.761	-0.009	7.045
	(-0.006;0.029)	(0.015;1.508)	(-0.034;0.019)	(0.539;13.550)
EU	-0.001	1.867	-0.003	3.340
	(-0.016;0.014)	(0.361;3.373)	(-0.019;0.012)	(1.441;5.238)
Canada	0.002	1.601	-0.040	3.145
	(-0.010;0.014)	(0.459;2.742)	(-0.067;-0.012)	(-0.646;6.935)
UK	-0.005	1.188	-0.013	0.309
	(-0.016;0.005)	(-0.202;2.578)	(-0.061;0.034)	(-9.284;9.901)
Japan	0.023	0.759	-0.005	5.556
	(0.002;0.044)	(0.131;1.386)	(-0.019;0.008)	(1.876;9.237)

Table 8 (Panel B): UIRP and the Jurado et al. (2015) Macroeconomic Uncertainty Index

Country	Low Uncertainty α_1	β_1	High Uncertainty α_2	β_2
Switzerland	0.007	0.637	-0.008	4.915
	(-0.009;0.023)	(-0.076;1.349)	(-0.039;0.023)	(-3.608;13.437)
EU	-0.002	1.864	-0.004	2.854
	(-0.016;0.011)	(0.176;3.551)	(-0.061;0.053)	(-3.281;8.526)
Canada	-0.002	1.623	-0.024	2.280
	(-0.015;0.011)	(0.454;2.792)	(-0.042;-0.005)	(-1.004;5.563)
UK	-0.011	1.351	-0.016	4.119
	(-0.024;0.002)	(-0.048;2.749)	(-0.053;0.021)	(-3.152;11.389)
Japan	0.026	0.846	-0.010	6.330
	(0.006;0.046)	(0.234;1.458)	(-0.023;0.004)	(1.116;11.544)

Table 8 (Panel C): UIRP and Ludvigson et al.'s (2016) Financial Uncertainty

	Low Uncertainty		High Uncertainty	
Country	α_1	β_1	α_2	β_2
Switzerland	0.005	0.451	-0.022	0.204
	(-0.014;0.025)	(-0.452;1.355)	(-0.049;0.005)	(-0.821;1.228)
EU	-0.002	1.851	-0.002	3.267
	(-0.017;0.013)	(0.309;3.392)	(-0.019;0.014)	(1.559;4.975)
Canada	0.001	1.236	-0.026	0.455
	(-0.013;0.014)	(-0.157;2.629)	(-0.042;-0.011)	(-1.055;1.965)
UK	-0.004	1.094	-0.022	2.423
	(-0.015;0.007)	(-0.334;2.522)	(-0.043;-0.001)	(-1.826;6.672)
Japan	0.019	0.726	-0.016	-0.372
	(0.001;0.037)	(0.085;1.368)	(-0.037;0.004)	(-1.033;0.289)

Table 8 (Panel D): UIRP and the Economic Policy Uncertainty Index

	Low Uncertainty		High Uncertainty	
Country	α_1	β_1	α_2	β_2
Switzerland	-0.008	0.139	0.017	2.848
	(-0.030;0.014)	(-0.732;1.009)	(-0.029;0.063)	(-9.289;14.985)
EU	-0.001	1.941	0.001	5.117
	(-0.016;0.014)	(0.560;3.322)	(-0.015;0.016)	(-2.002;12.237)
Canada	-0.009	1.082	-0.014	3.345
	(-0.022;0.004)	(-0.223;2.338)	(-0.102;-0.074)	(-6.923;13.613)
UK	-0.013	1.417	-0.005	1.818
	(-0.026;0.001)	(-0.010;2.84405)	(-0.030;-0.021)	(-4.356;7.991)
Japan	-0.011	-0.104	0.046	5.214
	(-0.026;0.004)	(-0.592;0.384)	(0.006;0.085)	(-9.746;20.173)

Table 8 (Panel E): UIRP and the Principal Component from the Bilateral Exchange Rate Uncertainty Indices

	Low Uncertainty		High Uncertainty	
Country	α_1	β_1	α_2	β_2
Switzerland	0.001	0.443	-0.009	-1.052
	(-0.01;0.02)	(-0.27;1.15)	(-0.05;0.03)	(-12.7;10.6)
EU	-0.004	2.276	0.003	1.201
	(-0.02;0.01)	(0.53;4.02)	(-0.02;0.03)	(-0.54;2.94)
Canada	-0.003	1.802	-0.01	0.627
	(-0.02;0.01)	(0.61;2.99)	(-0.04;0.02)	(-3.35;4.60)
UK	-0.012	1.368	-0.011	3.06
	(-0.02;0.00)	(-0.03;2.77)	(-0.05;0.03)	(-4.62;10.7)
Japan	0.023	0.765	-0.008	4.577
	(0.00;0.04)	(0.14;1.39)	(-0.02;0.01)	(-0.66;9.81)

Table 8 (Panel F): UIRP and the Global Foreign Exchange Volatility Risk

	Low Uncertainty		High Uncertainty	
Country	α_1	β_1	α_2	β_2
Switzerland	-0.019	-1.777	-0.014	-1.257
	(-0.03;-0.01)	(-2.57;-0.99)	(-0.04;0.01)	(-3.64;1.13)
EU	-0.006	0.319	0.001	3.951
	(-0.02;0.01)	(-0.51;1.14)	(-0.03;0.03)	(1.25;6.66)
Canada	-0.012	-0.5	0.006	-0.755
	(-0.02;-0.01)	(-1.56;0.56)	(-0.01;0.02)	(-2.27;0.76)
UK	-0.008	1.283	-0.023	0.774
	(-0.02;0.01)	(0.47;2.09)	(-0.04;-0.01)	(-0.20;1.75)
Japan	0	0.307	0.005	-0.582
	(-0.01;0.01)	(-0.52;1.14)	(-0.01;0.02)	(-1.99;0.83)

Notes to the table. The table reports parameter estimates (and 95% confidence intervals in parentheses) in eq. (3), where the measures of uncertainty are the VIX (Panel A), the Jurado et al. (2015) Macroeconomic Uncertainty Index (Panel B), the Ludvigson et al. (2016) Financial Uncertainty Index (Panel C), the Baker, Bloom and Davis (2016) Economic Policy Uncertainty Index (Panel D), our Global Uncertainy Index (Panel E) and the Menkoff et al. (2012) Global Foreign Exchange Volatility Risk (Panel F).

Table 9: UIRP and Overall Uncertainty Based on the Random Walk

	Low Uncertainty		High Uncertainty	
Country	α_1	β_1	α_2	β_2
Switzerland	0	1.091	-0.01	-0.22
	(-0.017;0.018)	(-0.034;2.217)	(-0.061;0.042)	(-1.848;1.409)
EU	-0.003	2.279	-0.004	3.091
	(-0.018;0.012)	(0.672;3.887)	(-0.02;0.011)	(-1.205;7.387)
Canada	-0.006	1.534	0.002	0.12
	(-0.019;0.008)	(0.093;2.975)	(-0.03;0.034)	(-3.498;3.739)
UK	-0.012	1.745	-0.012	3.06
	(-0.026;0.002)	(0.286;3.203)	(-0.039;0.014)	(-2.356;8.476)
Japan	0.01	0.895	-0.002	-0.355
	(-0.007;0.026)	(0.212;1.577)	(-0.035;0.031)	(-1.171;0.461)

Notes to the table. The table reports parameter estimates in eq. (3), where the measure of uncertainty is the overall exchange rate uncertainty index constructed based on random walk forecast errors.

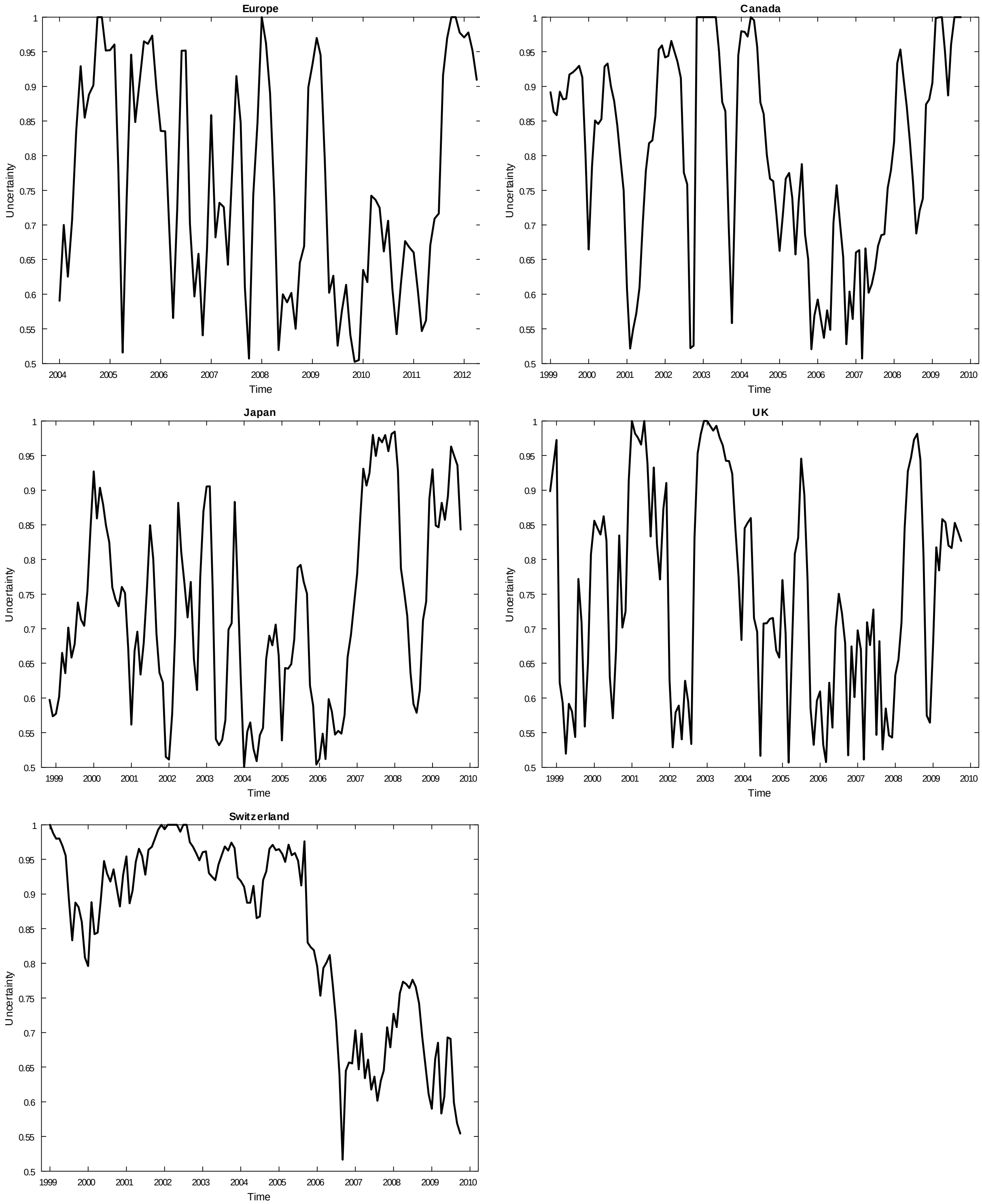
Table 10: UIRP Regressions for Additional Countries

	A. Full sample		B. Low Uncertainty		C. High Uncertainty	
Country	α_1	β_1	α_1	β_1	α_2	β_2
Australia	0.006	-0.302	-0.057	2.252	0.009	-0.634
	(-0.008;0.021)	(-1.007;0.403)	(-0.126;0.012)	(-0.322;4.825)	(-0.033;0.051)	(-1.909;0.641)
Sweden	0.001	-0.219	-0.004	0.451	0.013	-0.789
	(-0.009;0.011)	(-0.831;0.394)	(-0.019;0.011)	(-0.476;1.377)	(-0.004;0.03)	(-2.043;0.464)
South Africa	0.062	-0.706	0.046	-0.397	0.111	-1.685
	(0.031;0.094)	(-1.192;-0.219)	(-0.08;0.173)	(-2.57;1.776)	(-0.13;0.351)	(-5.417;2.047)
Norway	0.001	0.045	-0.049	3.33	-0.005	1.579
	(-0.009;0.01)	(-0.526;0.615)	(-0.093;-0.006)	(0.582;6.078)	(-0.042;0.032)	(-1.285;4.442)
New Zealand	-0.005	0.151	-0.123	4.147	-0.125	4.343
	(-0.028;0.017)	(-0.711;1.014)	(-0.19;-0.056)	(1.975;6.32)	(-0.299;0.049)	(-2.794;11.479)
Denmark	-0.001	-0.414	0.002	1.239	-0.001	0.397
	(-0.009;0.008)	(-1.073;0.245)	(-0.01;0.014)	(-0.223;2.701)	(-0.021;0.02)	(-0.755;1.548)

Notes to the table. The table reports parameter estimates of the traditional UIRP regression (Panel A) as well as parameter estimates in eq. (3), where the measures of uncertainty is the overall exchange rate uncertainty index constructed based on random walk forecast errors (Panel B).

Figures

Figure 1. Exchange Rate Uncertainty Indices



Notes to the figure. The figure plots the overall exchange rate uncertainty index for the benchmark countries in our sample.

Figure 2. Exchange Rate Uncertainty Indices and UIRP Coefficients

Figure 2(a)

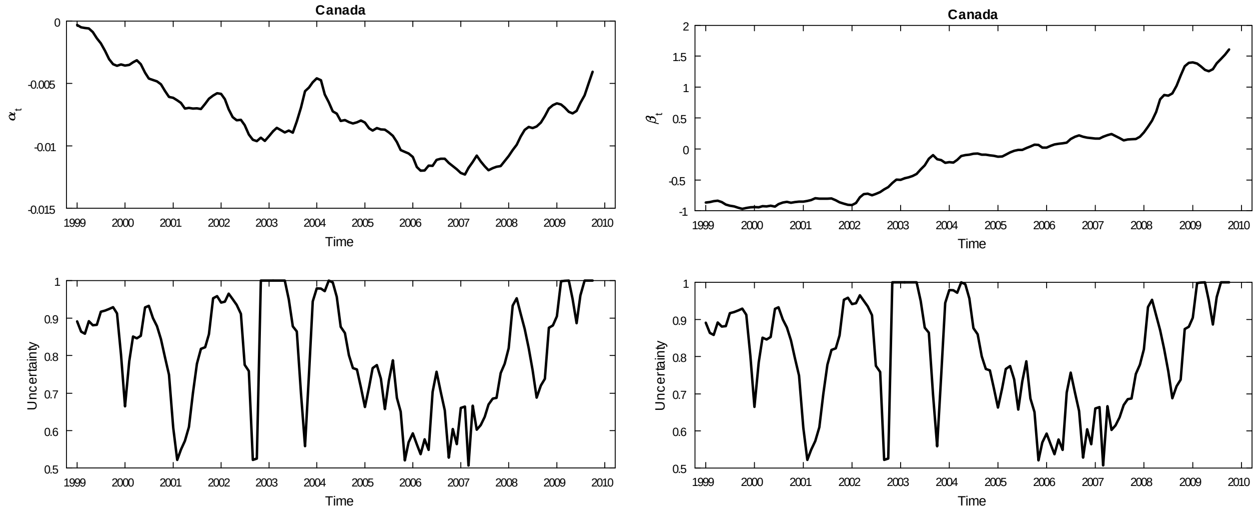


Figure 2(b)

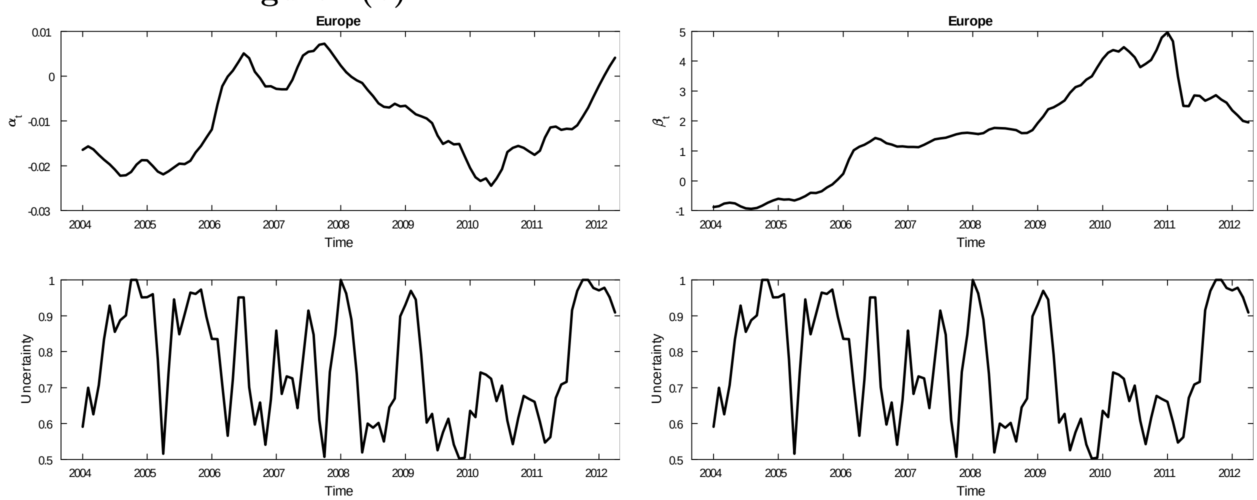


Figure 2(c)

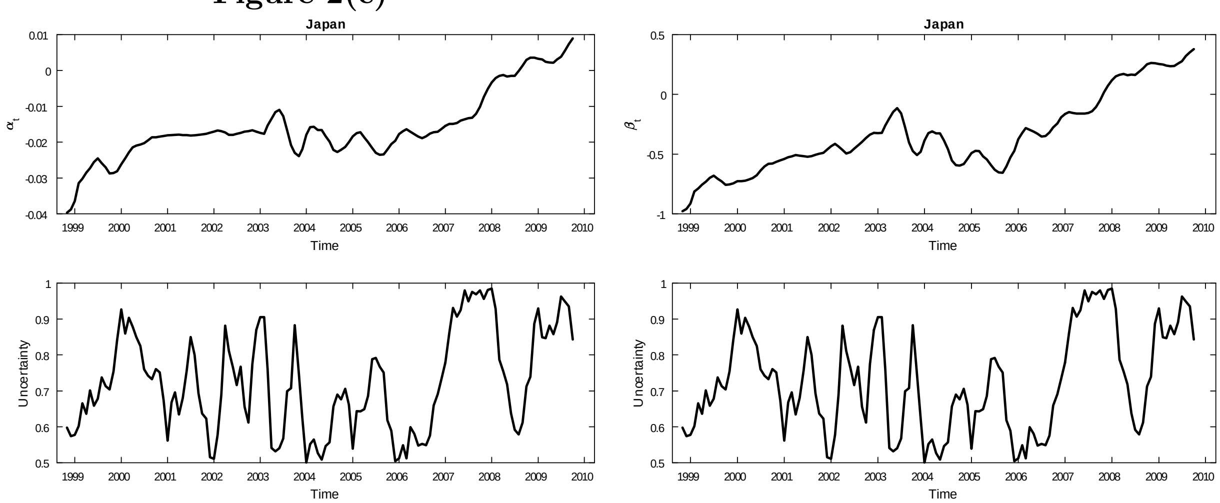


Figure 2(d)

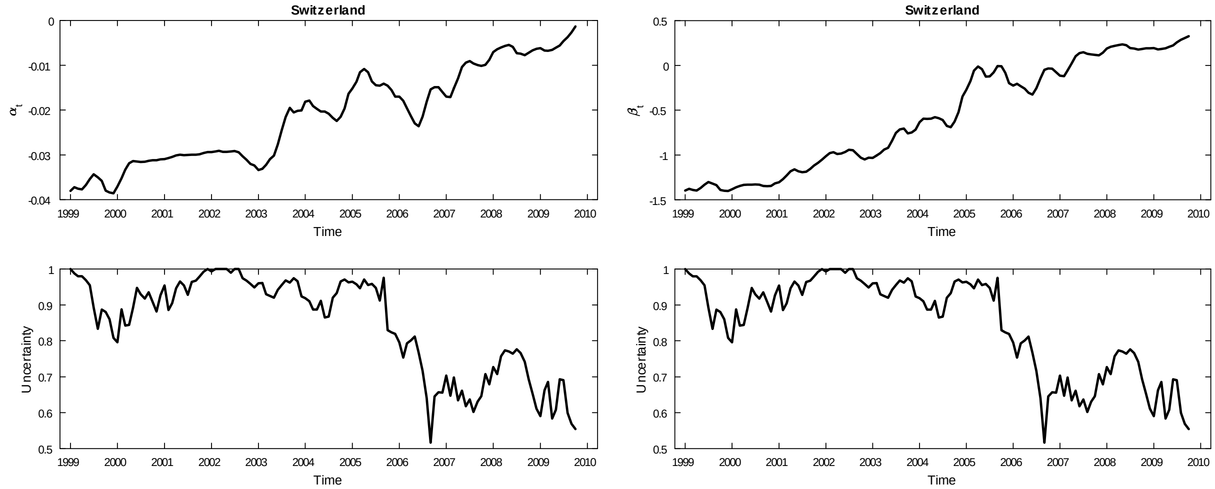
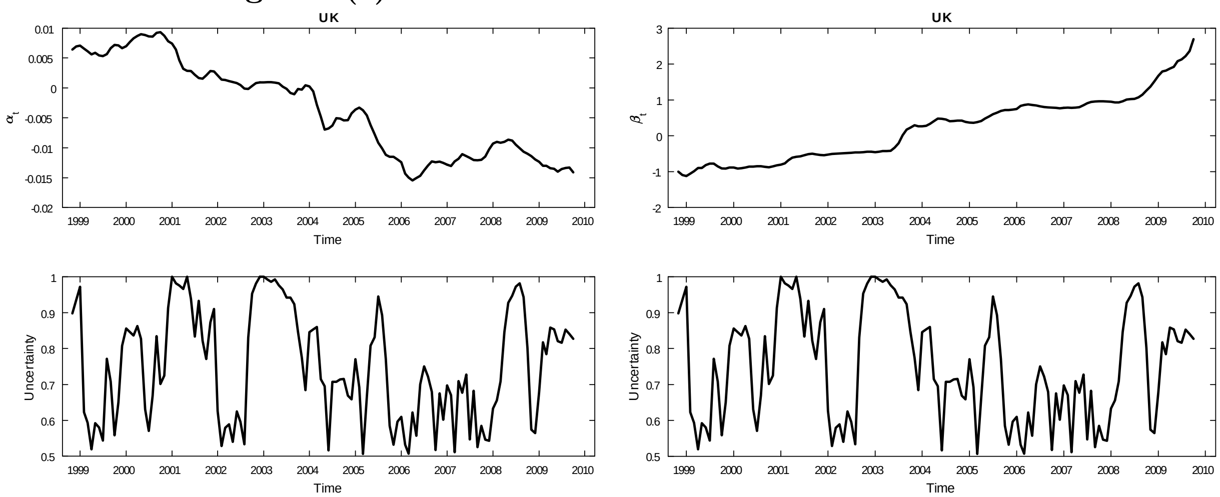
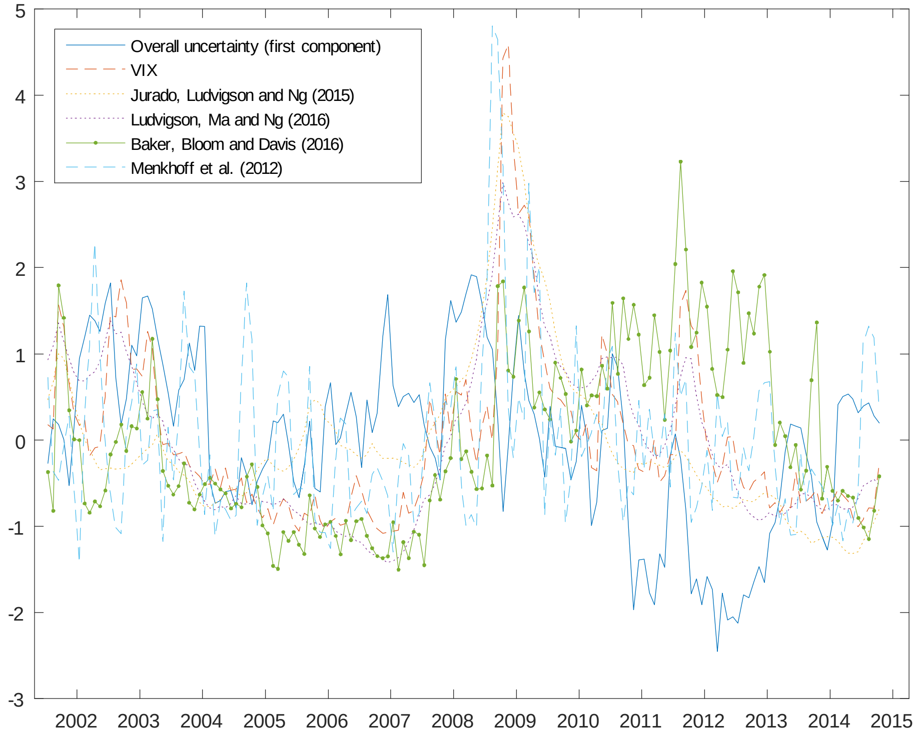


Figure 2(e)



Notes to Figure 2. The top panels in the figure plot the UIRP coefficients estimated in rolling windows (the constant is depicted on the left and the slope on the right). The bottom panels in the figures plot the overall exchange rate uncertainty index.

Figure 3. Global Uncertainty Indices



Note to Figure 3. The figure depicts time series of global uncertainty indices: the first principal component of the bilateral exchange rate uncertainty indices described in Section 3, labeled "Overall uncertainty (first component)"; the VIX; the Jurado, Ludvigson and Ng (2015) macroeconomic uncertainty index; the Ludvigson, Ma and Ng (2016) financial uncertainty index; the Baker, Bloom and Davis (2016) economic policy uncertainty index; and the Menkoff et al. (2012) global foreign exchange volatility risk.