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Variance risk premiums and the forward premium puzzle[☆]

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ABSTRACT

We provide new empirical evidence that world currency and U.S. stock variance risk premiums have nonredundant and significant predictive power for the appreciation rates of 22 with respect to the U.S. dollar, especially at the four-month and one-month horizons, respectively. The heterogeneous exposures of currencies to the currency variance risk premium are systematically rising along the line of inflation risk. We rationalize these findings in a consumption-based asset pricing model, with local consumption uncertainty and global inflation uncertainty characterized, respectively, by the stock and currency variance risk premiums.

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1. Introduction

This paper provides new empirical evidence that the time variation in expected currency returns is strongly related to the world currency variance risk premium (XVP) and to the stock variance risk premium (VP). The world XVP is measured as an average of the variance risk premiums of 17 available currencies with respect to the U.S. dollar. Each currency-pair's variance risk premium is measured as the option-implied variance minus the realized variance of currency returns. The VP is measured alternatively as the U.S. stock variance risk premium or as a global average of major countries' stock variance risk premiums. We find that an increase in XVP predicts a depreciation of foreign currencies with respect to the U.S. dollar, while an increase in VP predicts an appreciation of these currencies. Thus, XVP and VP seem to have different informational contents for future exchange rate returns.

We set our empirical exercise against the background of pervasive violations in the uncovered interest parity (UIP). For a large panel of 22 available currency rates against the U.S. dollar from 2000 to 2011, interest rate differentials are insignificant predictors for exchange rate returns, often with wrong negative signs and low R^2s of less than 1% for one- to four-month horizons. However, including the world

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XVP increases the R^2 by 2.2% at the one-month horizon, 8% at the four-month horizon, and 1.5% at the 12-month horizon. The slope coefficients associated with the world XVP are uniformly negative and significant—a higher world XVP indicates greater global uncertainty, hence higher U.S. dollar safety value for currency investors. Including the U.S. VP increases the R^2 by 5.3% at the one-month horizon and almost zero at the 12-month horizon. The slope coefficients are, in this case, uniformly positive and significant for one- to six-month horizons—a higher VP indicates greater U.S. uncertainty, hence higher return premium compensation for currency investors.

To better understand the underlying economic mechanism behind the predictive power of variance risk premiums, we perform several in-depth diagnostics. First we run the empirical tests for each of the 22 currencies individually, and the findings for the predictive power of variance risk premiums remain intact except for the Japanese yen, a traditional funding currency, and for a few outliers, like the Hong Kong dollar and the Philippines peso, which have pegged or managed floating exchange regimes. Then, we sort currencies into portfolios and find that currencies of countries with higher average inflation tend to have more negative loading coefficients on XVP and higher forex return prediction R^2s. The heterogeneous forex return predictability by XVP and the exposure of XVP predictability to inflation risk shed light on how to provide a structural interpretation of our new empirical findings.

The joint predictability of XVP and VP for exchange rate returns remains robust if we consider a pre-global financial crisis sample. For XVP, the results also remain the same if the realized variance is replaced by the expected variance from an AR(1) model (Drechsler and Yaron, 2011), if the Black-Scholes implied variance is replaced by a model-free implied variance (Britten-Jones and Neuberger, 2000), or if the realized variance is calculated from high-frequency intraday forex returns (Andersen et al., 2001). For the VP, it makes no material difference if we use the U.S. VP or an equally weighted or value-weighted average of major countries' VPs. The predictability patterns of variance risk premiums for forex returns also hold after controlling for the countercyclical risk premium component of forex returns (Lustig et al., 2014).

To rationalize our empirical findings, we introduce a two-country consumption-based asset pricing model. Our model follows Bollerslev et al. (2009) to model the real economy and introduces a process for inflation, in line with Bansal and Shaliastovich (2013) and Zhou (2011), for the model to have realistic implications for nominal appreciation rates. In our model, both countries' real consumption growth processes are orthogonal to each other, while their inflation processes are exposed to global inflation. Moreover, global inflation level and volatility shocks are correlated. The orthogonality of the real-economy components of our model and the heterogeneous exposures to common inflation (level and volatility) risk yields the key implications that support our forex return predictability evidence. On the one hand, the XVP implied by the model reveals information about the global inflation uncertainty that cannot otherwise be obtained from domestic VPs. Thus, the XVP contains useful information to explain the time variation of appreciation rates that is additional to the VP. On the other hand, the predictive power of the XVP for the appreciation rate between two currencies depends crucially on the heterogeneity in the exposure of each country's inflation process to the global inflation risk.

We calibrate the parameters in the model to match the observed real growth and inflation dynamics for the United States and the United Kingdom and the dollar-pound XVP. For the benchmark calibration scenario, our model is able to qualitatively replicate the patterns for the predictive power of the currency and stock variance premiums for the exchange rate return. We also find that predictability patterns are highly sensitive to the degree of heterogeneity in the exposure to global inflation across countries. In particular, the predictability pattern of XVP for appreciation rates becomes more negative—an increase in XVP is followed by a depreciation of the foreign currency with respect to the U.S. dollar—as the United States is assumed to be less exposed to global inflation than the foreign economy, which explains the empirical evidence for the inflation-sorted currency portfolios.

1.1. Literature

Recent literature focuses on the role of the volatility risk premium in explaining the time variation in currency returns. Della Corte et al. (2011) provide empirical evidence that the volatility term premium is positive, time-varying, and predictable. In a related paper, Menkhoff et al. (2012) document the finding that global forex volatility risk is priced in currency markets (see also Bakshi and Panayotov, 2013). Chernov et al. (2015) find evidence that jump risk in currency variance may be priced in forex markets but is unrelated to interest rates or macroeconomic news. Using different methodologies, Farhi et al. (2015), Jurek (2014) and Brunnermeier et al. (2009) relate the high observed prices of currency options to the desire of agents to hedge rare and severe changes in exchange rate movements.[1] Finally, Mueller et al. (2015) find that the forex correlation risk premium is also priced in currency markets. To the best of our knowledge, our paper is the first one to show that both currency and stock variance premiums provide useful information to explain exchange rate returns at short horizons.

Our work is also intimately related to the early evidence that exchange rate volatility is time varying (Engle, 1982; Baillie and Bollerslev, 1989; Engel and Hamilton, 1990; Engle et al., 1990; Gagnon, 1993). However, we focus on the unique information content from the forex derivatives market not only to pin down the dynamics of forex volatility but also to show that this volatility risk is actually priced in forex markets. Graveline (2006) shows that the information from exchange rate options is valuable for the estimation of the exchange rate volatility that is much harder to identify using only time-series data for exchange rates. Bakshi et al. (2008) show that jumps are crucial to capture the currency return dynamics and to

[1] The rare disaster model in Farhi and Gabaix (2016) aims to rationalize this empirical finding. Burnside et al. (2011) provide a related interpretation based on the peso problem.

generate realistic currency option pricing behaviors. In fact, Bates (1996) and Guo (1998) provide evidence that the dollar/German mark variance risk is priced in the forex options market within a Heston (1993)-type model.

There is certainly a large literature documenting the forward premium puzzle or the deviation from the uncovered interest parity (UIP). Early works by Hansen and Hodrick (1980), Fama (1984), Bansal (1997), and Backus et al. (2001), among others, find evidence that, as a consequence of this deviation, carry trade excess returns are large, on average positive, and predictable. Recent works by Lustig and Verdelhan (2007), Lustig et al. (2014), Verdelhan (2015), and Colacito et al. (2015) relate the cross-sectional evidence of carry trade strategies to fundamental risk factors (consumption, dollar, carry-trade, long-run growth). Motivated by the recent finding that the stock variance premium can predict international stock market returns (Bollerslev et al., 2014; Londono, 2015), we investigate the different informational content of currency and stock variance risk premiums for explaining the predictable time variation in the forward premium.

The rest of the paper is organized as follows. Section 2 introduces our XVP and VP measures and the data used to calculate them. In Section 3, we summarize the main empirical findings for the predictive power of XVP and VP for forex appreciation rates, the heterogeneous nature of this predictability, and the linkage to global inflation risk. In Section 4, we introduce a two-country general equilibrium model to understand our main empirical findings. Finally, Section 5 concludes.

2. The currency and stock variance risk premiums

In this section, we introduce a measure for the world currency variance risk premium calculated as the equally weighted average of the variance risk premiums of a total of 17 currencies with respect to the U.S. dollar. We also describe the stock variance risk premium (VP), which is measured as the U.S. VP or as an average of the VPs of major countries with stock options data available.

2.1. The world XVP

Following the convention for the stock VP (Bollerslev et al., 2009; Drechsler and Yaron, 2011), we define the forex or currency variance risk premium (XVP) of the returns in U.S. dollars per one unit of foreign currency as

$$XVP_t(h) \equiv E_t^Q\left(\sigma^2_{c,t,t+h}\right) - E_t^P\left(\sigma^2_{c,t,t+h}\right). \quad (1)$$

That is, the h-month ahead XVP equals the difference between the risk-neutral (Q) and the physical (P) expectations of the currency return variance between months t and $t+h$, $\sigma^2_{c,t,t+h}$. For the benchmark XVP measure in our empirical exercise in Section 3, we substitute the risk-neutral expectation with the h-month ahead currency option-implied variance, using Black-Scholes at-the-money (ATM) options; and we substitute the physical expectation with the realized variance calculated as the sum of squared log daily currency returns between $t-h$ and t. We also assess the robustness of our results to three alternative variance risk premium measures; one in which we use the model-free approach to measure the risk-neutral expected variance (Britten-Jones and Neuberger, 2000), one in which we replace the physical expectation by a forecast obtained from an AR(1) model (Drechsler and Yaron, 2011), and one in which we use high-frequency data to calculate the currency realized variance for some available currencies (Andersen et al., 2001).

The world XVP is calculated as an equally weighted average of all countries' XVPs.[2] In the remainder of the paper, we focus on the six-month horizon, as we find that, although the predictability patterns are very similar for world XVPs across different horizons, the gains in predictive power for future appreciation rates with respect to the U.S. dollar are much higher for the six-month XVP than for XVPs at any other horizon.[3]

2.2. The VP

Similar to the XVP, we define the one-month stock variance risk premium (VP) as

$$VP_t \equiv E_t^Q\left(\sigma^2_{r,t,t+1}\right) - E_t^P\left(\sigma^2_{r,t,t+1}\right), \quad (2)$$

where $\sigma^2_{r,t,t+1}$ is the stock return variance between months t and $t+1$. We calculate the VP as the difference between the (model-free) option-implied and the expected realized stock variance. As we did for the XVP, we assume that the expected stock realized variance is given by $E_t(RV^2_{t,t+1}) = RV^2_{t-1,t}$, where $RV^2_{t-1,t}$ is the realized variance of the stock index calculated using one-month nonoverlapping rolling windows of daily (log) stock returns. We consider the following three alternative measures of the world stock variance risk premium: the U.S. stock variance risk premium (VP_{US}), the equally weighted average stock variance risk premium (VP_{EW}), and the value-weighted average stock variance premium (VP_{VW}). The average VPs are calculated using the variance risk premium for the headline stock indexes for the following countries: United States, Germany, Japan, and the United Kingdom.[4] The monthly value-weighted average VP is calculated, following Bollerslev et al. (2014), using lagged total market capitalization for the four countries considered.

2.3. Data

Our sample runs from January 2000 to December 2011 and covers the exchange rates (with respect to the U.S. dollar) of the following countries and their respective currencies (in parentheses): the Euro Area (EUR), Japan (JPY), Great Britain (GBP), Switzerland (CHF), Australia

[2] In Section 2.4, we summarize the results from a principal components analysis which supports the election of the equally weighted average to characterize the world XVP.

[3] The finding that the six-month world XVP is a more useful predictor of appreciation rates with respect to the U.S. dollar holds for the alternative variance risk premium measures considered. The results for the predictive power of one-, three-, and 12-month world XVPs are omitted to save space and are available, upon request, from the authors.

[4] Considering alternative average VP measures, including one in which we use VPs for all other countries with model-free option-implied volatility data available, has virtually no impact on the main empirical results in our paper.

Table 1
One-month currency appreciation rates with respect to the U.S. dollar, summary statistics.
This table reports the summary statistics for the time series of one-month fluctuations of the logarithm of foreign exchange rates with respect to the U.S. dollar. The appreciation rates are expressed in percent. The exchange rates are quoted in units of U.S. dollar per one unit of foreign currency—a positive sign corresponds to an appreciation of the foreign currency with respect to the U.S. dollar. We also report the average pairwise correlation between each currency and all other currencies considered (Avg. corr.).

	EUR	JPY	GBP	CHF	AUD	CAD	SEK	NZD	KRW	SGD	NOK
Mean	0.20	0.23	−0.03	0.40	0.33	0.24	0.31	0.31	−0.17	0.19	0.23
Median	0.26	−0.02	−0.02	0.14	0.53	0.27	0.88	0.88	−0.06	0.23	0.29
St. dev.	3.22	2.81	2.64	4.04	4.04	2.80	3.61	4.07	1.71	1.71	3.44
Skew.	−0.21	−0.30	−0.32	0.07	−0.76	−0.61	−0.10	−0.52	−0.59	−0.84	−0.55
Kurt.	3.89	3.41	4.83	4.51	5.14	6.30	4.50	4.50	3.47	7.22	4.51
AR(1)	0.02	−0.04	0.10	−0.08	0.06	−0.06	0.06	0.06	−0.09	−0.09	0.07
Avg. corr.	0.60	0.19	0.45	0.54	0.45	0.45	0.60	0.54	0.39	0.56	0.56
	PLN	**ZAR**	**CZK**	**DKK**	**THB**	**TWD**	**HKD**	**HUF**	**INR**	**MYR**	**PHP**
Mean	0.14	−0.02	0.44	0.20	0.12	0.01	0.06	0.06	−0.14	0.13	−0.05
Median	0.47	0.75	0.75	0.20	0.28	−0.02	0.00	0.61	0.00	0.00	0.02
St. dev.	4.32	3.54	3.87	1.75	1.75	1.46	0.14	4.47	1.42	1.42	2.03
Skew.	−0.89	−0.27	−0.40	−0.20	−0.29	−0.01	0.99	−1.21	−0.62	−0.85	−1.09
Kurt.	4.85	7.25	3.51	3.89	3.73	3.94	6.57	6.57	5.79	8.56	7.64
AR(1)	0.13	−0.05	0.04	0.03	0.13	0.21**	0.00	0.07	0.18*	−0.09	0.06
Avg. corr.	0.55	0.49	0.57	0.60	0.47	0.47	0.16	0.56	0.43	0.34	0.34

(AUD), Canada (CAD), Sweden (SEK), New Zealand (NZD), South Korea (KRW), Singapore (SGD), Norway (NOK), Poland (PLN), South Africa (ZAR), the Czech Republic (CZK), Denmark (DKK), Thailand (THB), Taiwan (TWD), Hong Kong (HKD), Hungary (HUF), India (INR), Malaysia (MYR), and the Philippines (PHP). For 17 of these 22 currencies (excluding the HKD, the HUF, the INR, the MYR, and the PHP), we can calculate the XVP as the difference between the option-implied and the realized currency return variance. The ATM implied volatility for these 17 currency pairs is obtained from J.P. Morgan's over the counter (OTC) currency options database while the spot rates are obtained from Bloomberg.

The stock option-implied volatility and the daily spot price for the headline stock indexes of the United States, Germany, Japan, and the United Kingdom are obtained from Bloomberg. Monthly total market capitalizations for the four countries, which are used to calculate the value-weighted average VP, are obtained from Compustat.

We also calculate the interest rate differential between each country and the United States from h-month zero-coupon rates calculated by the Board of Governors of the Federal Reserve system using data from each country's central bank.

Finally, to assess the fundamental determinants of the heterogeneous exposure of each country's currency appreciation rate to the world XVP, for all countries, we collect data on real gross domestic product (GDP) deflator from the Federal Reserve Board and Haver Analytics.

2.4. Summary statistics and stylized features

Table 1 reports summary statistics and average pairwise correlations for one-month currency appreciation rates with respect to the U.S. dollar. The mean appreciation against the U.S. dollar ranges between −0.17% (KRW) and 0.44% (CZK). Appreciation rates display a relatively high volatility (2.95% on average). The appreciation rate volatility is unusually low for the HKD (0.14%), most likely because this currency has been pegged to the U.S. dollar since 1983.[5] In contrast, the volatility is the highest for the KRW (5.12%). Some currencies, other than the HKD, deviate from the normal distribution. In particular, kurtosis is relatively high for the SGD (7.22), the ZAR (7.25), the MYR (8.56), and the PHP (7.64). Also, skewness is negative for all of the currencies in our sample except for the CHF and the HKD. Skewness is particularly negative for the HUF (−1.21), the SGD (−0.84), the PLN (−0.89), and the MYR (−0.85). It is particularly interesting to note that currency rates with respect to the U.S. dollar have a common component. In particular, the average pairwise correlation for all currencies' depreciation rates with respect to the U.S. dollar is 0.48. Some currencies display relatively high average pairwise correlations, such as the EUR (0.60), the SEK (0.60), the SGD (0.58), and the DKK (0.60). In contrast, the JPY (0.19) and the HKD (0.16) have the lowest average pairwise correlations with all of the other currencies considered.

Table 2 reports summary statistics for six-month XVPs for the 17 countries in our sample with currency options data available. In the first column of the table, we also report summary statistics for the world XVP calculated as the equally weighted average of all currencies' variance risk premiums. The XVP is, on average, positive and significant for six currencies (EUR, GBP, SGD, DKK, THB, and TWD) at confidence levels above 5%. The world XVP, plotted in Fig. 1, is also, on average, positive and displays relatively large volatility (49.46%^{2}). The XVP is particularly volatile for the AUD (160.97%^{2}), the KRW (125.04%^{2}), and the ZAR (182.89%^{2}). The world XVP deviates from the normal distribution with large spikes, especially around the Lehman Brothers episode, a relatively large kurtosis (16.33), and negative skeweness.[6] The world XVP is quite

[5] The HKD has been pegged to the dollar at 7.8 since 1983. In 2005, the Hong Kong monetary authority also committed to keep the exchange rate with respect to the U.S. dollar between HKD 7.74 and HKD 7.85.

[6] Fan et al. (2016) find that negative variance risk premiums, also prominent in the U.S. stock market around the Lehman Brothers episode,

Table 2
Currency variance risk premiums (XVPs), summary statistics.
This table reports the summary statistics for the six-month currency variance risk premiums (XVPs) of all available currencies with respect to the U.S. dollar. The XVPs are expressed in annualized squared percent. We also report the summary statistics for the world XVP, which is calculated as the equally-weighted average of all currencies' variance risk premiums. Our sample runs from January 2000 to December 2011. Each currency's variance risk premium is measured as the difference between the square of the six-month at-the-money forex option-implied volatility and the realized variance of the exchange rate appreciation with respect to the U.S. dollar. The forex return realized variance is calculated using six-month lagged rolling windows of daily (log) appreciation rates between each currency and the U.S. dollar. *, ** and *** represent the usual 10%, 5%, and 1% significance levels. To assess the significance of the mean *XVP*s, the standard errors are corrected by Newey-West with six lags. We also report the average correlation between each currency's and all other currencies' variance risk premiums (Avg. corr.).

	World *XVP*	EUR	JPY	GBP	CHF	AUD	CAD	SEK	NZD
Mean	1.44	16.45***	9.65	13.24**	−7.39	−42.57*	0.06	−5.63	−12.34
Median	4.64	10.32	6.45	9.53	−8.36	−8.36	2.16	2.62	0.80
St. dev.	49.46	42.02	43.57	39.76	62.29	160.97	35.56	78.99	75.84
Skew.	−2.71	1.47	−1.10	0.61	−1.97	−4.43	−2.06	−2.36	−2.51
Kurt.	16.33	9.35	7.20	11.68	9.19	23.88	14.86	12.70	13.63
AR(1)	0.79***	0.67***	0.74***	0.77***	0.87***	0.87***	0.67***	0.83***	0.79***
Avg. corr.	0.37	0.49	0.35	0.51	0.28	0.28	0.51	0.51	0.44

	KRW	SGD	NOK	PLN	ZAR	CZK	DKK	THB	TWD
Mean	2.78	12.15***	0.21	2.76	−42.45	−8.62	16.71***	30.15***	18.17***
Median	8.25	4.36	4.36	10.72	8.39	1.38	18.36	18.36	11.97
St. dev.	125.04	18.56	65.07	92.14	182.89	66.46	42.66	36.21	23.10
Skew.	−4.59	−2.34	−2.34	−1.41	−3.29	−2.19	1.53	1.17	2.08
Kurt.	29.63	12.23	12.23	8.93	15.58	11.47	9.52	3.84	8.41
AR(1)	0.71***	0.77***	0.81***	0.70***	0.78***	0.67***	0.67***	0.79***	0.84***
Avg. corr.	0.38	0.24	0.54	0.30	0.30	0.46	0.48	−0.05	0.03

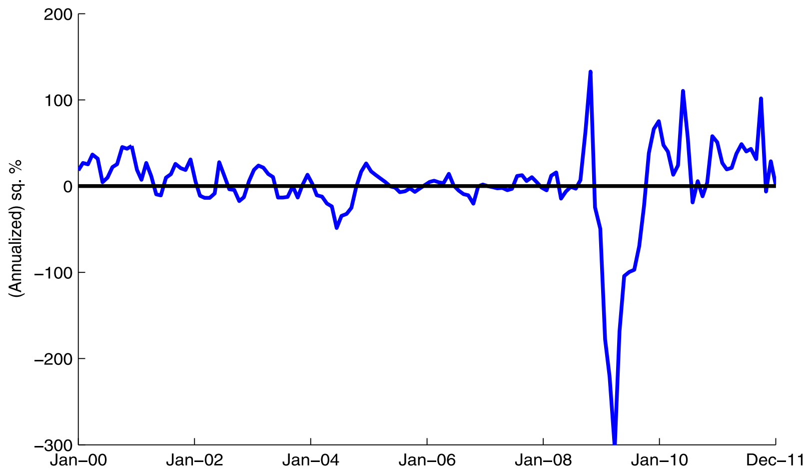


Fig. 1. World currency variance risk premium (XVP). worThe figure shows the six-month world XVP, which is calculated as the equally weighted average of the variance risk premiums of 17 currencies with respect to the U.S. dollar (see Table 2). Each currency's variance risk premium is measured as the difference between the square of the six-month at-the-money forex option-implied volatility and the realized variance of the exchange rate appreciation with respect to the U.S. dollar. The forex return realized variance is calculated using six-month lagged rolling windows of daily (log) appreciation rates between each currency and the U.S. dollar.

persistent (its AR(1) coefficient is 0.79), which is not surprising, as the six-month horizon requires a large number of overlapped windows to calculate the realized currency variance. Another interesting feature of XVPs is their large average pairwise correlation (0.37). In fact, the first principal component of XVPs explains 50% of the total variation. The evidence from the principal component analysis supports the use of the equally weighted average of XVPs to proxy the world XVP, as the weights associated with all countries' XVPs in the first principal component are positive for almost all currencies and of a similar magnitude.[7]

can be explained by the gains and losses on market makers delta-hedged positions. An alternative hypothesis to explain negative variance risk premiums is related to the predictive power of implied variance for realized variance (Jiang and Tian, 2005; Ait-Sahalia et al., 2015). To be sure, as we show in Section 3.4, our main empirical findings are robust to considering a subsample before the Lehman Brothers episode and to alternative variance premium measures that are less prone to experience large negative spikes.

[7] In unreported results, we show that the main empirical results in Section 3 are virtually unchanged when we approximate the world XVP as the first principal component of all countries' XVPs.

Table 3
Stock variance risk premiums (VPs), summary statistics.
This table reports the summary statistics for the stock variance risk premium (VP), which is calculated as the difference between the (model-free) option-implied and the realized stock return variance. The VPs are expressed in annualized squared percent. The VP is alternatively measured as the U.S. stock variance premium (VP_{US}), the equally weighted average stock variance premium (VP_{EW}), and the value-weighted average stock variance premium (VP_{VW}). The average stock variance risk premiums are calculated using the VPs for the following countries: United States, Germany, Japan, and the United Kingdom. For these four countries, the weights in the value-weighted measure are calculated using lagged total market capitalizations. We also report the correlation between the VP measures and the six-month world XVP, *corr*(*VP, XVP*), as well as the cross-correlations among the three VPs.

	VP_{US}	VP_{EW}	VP_{VW}
Mean	88.69*	104.18***	100.36**
Median	109.31	115.77	105.33
St. dev.	418.86	351.10	379.72
Skew.	−4.32	−5.54	−5.16
Kurt.	32.97	53.04	46.26
AR(1)	0.32***	0.18**	0.28***
corr(*VP, XVP*)	−0.17	−0.30	−0.24
Correlations	VP_{US}	VP_{EW}	VP_{VW}
VP_{US}	1		
VP_{EW}	0.89	1	
VP_{VW}	0.98	0.96	1

Table 3 reports summary statistics for the one-month VPs, while Fig. 2 shows their time series. Irrespective of the proxy measure used, VP is, on average, positive and significant at confidence levels above 10%. VPs are also relatively volatile (418.9%^{2}, 351.10%^{2}, and 379.7%^{2} for VP_{US}, VP_{EW}, and VP_{VW}, respectively). Interestingly, although all VP measures are highly correlated with each other, as is also evident from the figure, their correlations with the world XVP are moderate and negative (−0.17, −0.30, and −0.24 for VP_{US}, VP_{EW}, and VP_{VW}, respectively).

In the following section, we show that stock and currency variance risk premiums contain differential information to explain the time variation in forex returns. In Section 4, we also propose a model that provides the intuition for the differential informational content of VP and XVP and for the low correlation between the two premiums. Specifically, we show that, while the VP characterizes domestic real uncertainty, the XVP reveals information about global inflation uncertainty.

3. The predictive power for forex appreciation rates

In this section, we conduct a comprehensive analysis of return predictability for the 22 currencies in our sample from the world currency and stock variance risk premiums (XVP and VP). We first present our benchmark results for the panel-data regression setting. We then explore the heterogeneity in the predictability patterns across currencies using individual-currency regressions, and provide a fundamental explanation to these patterns based on inflation risk. In the final part of the section, we provide a set of robustness tests for our benchmark panel-data setting.

3.1. Panel data regressions

Our benchmark empirical regression setting for the predictive power of the world XVP is

$$s_{i,t+h} - s_{i,t} = b_{i,0}(h) + b_{IR}(h)[y_{US,t}(h) - y_{i,t}(h)] + b_{XVP}(h)XVP_t + u_{i,t+h}, \tag{3}$$

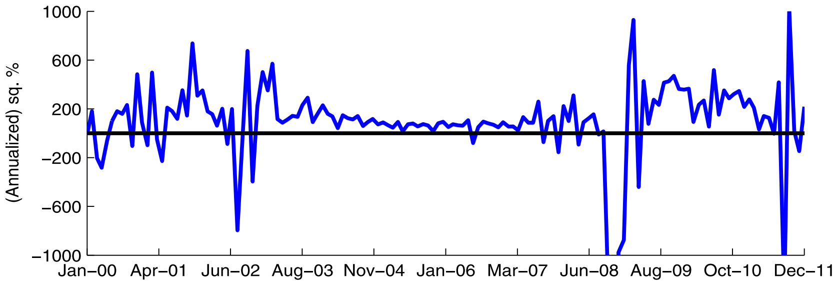
where $s_{i,t}$ is the log of the exchange rate (in dollars per one unit of each one of the foreign currencies considered), $y_{US,t}(h) - y_{i,t}(h)$ is the interest rate differential for $h-$period zero-coupon bonds between the United States and the foreign country, and XVP_t is the world XVP calculated as described in Section 2.1. The coefficients in Eq. (3) are estimated using pooled ordinary least squares (OLS) in which the coefficients associated with the interest rate differential and XVP are restricted to be homogeneous across currencies.[8]

Table 4 reports the predictive power of the interest rate differential and the additional predictive power of the world XVP for h-month ahead appreciation rates in Panels A and B, respectively. The uncovered interest rate parity (UIP) predicts that the expected *appreciation* of the foreign currency must equal the difference between domestic and foreign interest rates, such that an investor is indifferent between holding a domestic or a foreign bond. However, vast empirical evidence since Fama (1984) shows exactly the opposite—an increase in the domestic interest rate corresponds rather to a *depreciation* of the foreign currency. The UIP violation is especially challenging at short horizons (Hodrick, 1987). Our results are in line with deviations from the UIP reported in the literature. That is, the coefficient associated with the interest rate differential is significantly different from one for all horizons considered, and the estimated coefficient is even negative for the one-, nine-, and 12-month ahead appreciation rates (Panel A). The R^2s in the univariate regressions for the predictive power of interest rate differentials are as low as 0.26% for the one-month horizon and reach a maximum of 3.3% at the 12-month horizon.

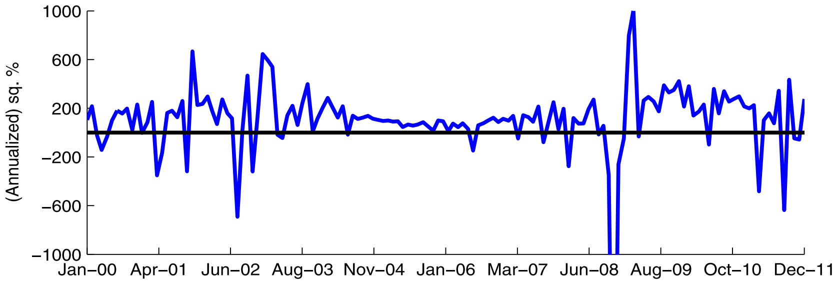
The results in Panel B suggest that adding the XVP does not have much impact on the deviations from the UIP. Specifically, in the multivariate regression setting with the interest rate differential and the world XVP, the estimated coefficients associated with the interest rate differential are statistically different from one for all horizons considered and follow a pattern similar to the coefficients for the univariate regression setting in Panel A. Interestingly, however, the estimated coefficients associated with the interest differential all become slightly less deviated from one, when the world XVP is included. Nevertheless, the improvement is not large enough to resolve the UIP puzzle.

The results in Panel B also suggest that the XVP plays a key role in predicting future appreciation rates. The statistical significance of XVP is above the 1% level for all horizons considered. Moreover, the estimated coefficient for the predictive power of XVP is economically meaningful—a one-hundred units increase in the monthly XVP, which

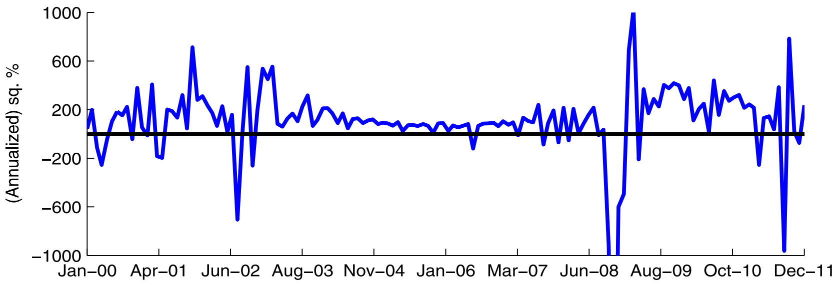
[8] As pointed out by Bansal and Dahlquist (2000), a panel-data setting reduces imprecision in the estimation of currency-specific parameters (also see Baillie and Bollerslev, 2000).



Panel A. VP_{US}



Panel B. VP_{EW}



Panel C. VP_{VW}

Fig. 2. Stock variance risk premiums (VPs). The figure shows the VPs measured as the difference between the square of the (model-free) stock-option-implied volatility and the realized stock return variance. We report the U.S. variance risk premium and the equal- and value-weighted average stock variance risk premiums (for the United States, Germany, Japan, and the United Kingdom) in Panels A, B, and C, respectively (see Table 3).

is equivalent to an increase of 1,200 units in the observed XVP (see Fig. 1), corresponds to a four-month ahead annual *depreciation* of 11.95% of the foreign currencies with respect to the U.S. dollar. The predictive power of the world XVP is maximized at a medium four- to six-month horizon. Moreover, the gains in predictive power with respect to the individual predictive power of the interest rate differential, $R^2 - R^2_y$, are considerable and maximized at the four-month horizon (8%). The evidence for the predictive power of the world XVP for appreciation rates relates to the finding in Della Corte et al. (2016) that country-level currency volatility risk premium has predictive power for the cross section of currency returns.

Table 5 reports the results for the predictive power of the VP for forex returns. Our regression setting for the predictive power of the VP is similar to that in Eq. (3),

$$s_{i,t+h} - s_{i,t} = b_{i,0}(h) + b_{IR}(h)[y_{US,t}(h) - y_{i,t}(h)] + b_{VP}(h)VP_t + u_{i,t+h}, \tag{4}$$

where VP_t is alternately the U.S. stock variance risk premium (in Panel A), the equally weighted average stock variance risk premium (in Panel B), or the value-weighted average stock variance risk premium (in Panel C). The results in Panel A suggest that the U.S. VP plays a key role in explaining the future appreciation rate for all currencies considered, especially at the short one- to four-month horizon. In particular, following an increase in the U.S. VP, the U.S. dollar tends to depreciate with respect to all currencies—a one-hundred units increase in the monthly U.S. VP corresponds to a one-month ahead annual *appreciation* of 2.10% of the foreign currencies with respect to the U.S. dollar. The statistical significance is above the 1% level for all currencies for horizons between one and six months. This result extends the evidence in Zhou (2009), who shows that the U.S. VP has predictive power for one-month dollar/EUR and dollar/GBP returns.[9] The predictive

[9] Interestingly, Londono (2015) finds that the U.S. VP is also a useful predictor for international stock returns, and Rapach et al. (2013) find that

Table 4
The predictive power of the world XVP for exchange rate returns with respect to the U.S. dollar.
This table reports the estimated coefficients for the following panel-data regressions:

$$s_{i,t+h} - s_{i,t} = b_{i,0}(h) + b_{IR}(h)[y_{US,t}(h) - y_{i,t}(h)] + b_{XVP}(h)XVP_t + u_{i,t+h},$$

where $s_{i,t}$ is the dollar exchange rate of currency i (in units of U.S. dollar per one unit of foreign currency), $y_{US,t}(h) - y_{i,t}(h)$ is the interest rate differential for $h-$month zero-coupon bond rates between the United States and country i, and XVP is the six-month world currency variance risk premium (XVP) calculated as the equally weighted average of all available currencies' variance risk premiums (see Table 2). To facilitate the interpretation of the estimated coefficients, we divide XVP by 12 (equivalent to monthly XVP). The standard errors are corrected by panel-data Newey-West with h lags (standard deviations are reported in parentheses). *, **, and *** represent the usual 10%, 5%, and 1% significance levels. For the interest rate differential, $y_{j,t}(h) - y_{i,t}(h)$, the null hypothesis corresponds to $b_{IR} = 1$ (that is, whether the UIP holds). The sample period runs from January 2000 to December 2011. The estimated regression currency-specific constants are left unreported, to save space. We report the R^2 of the regression and, for the multivariate regression in Panel B, the gains in R^2s with respect to a univariate regression for the interest rate differential, $R^2 - R^2_y$.

Panel A: Interest rate differential (deviations from UIP)							
h	1	2	3	4	6	9	12
$y_{US}(h) - y_i(h)$	−0.13***	0.02**	0.08**	0.16**	0.07**	−0.07***	−0.11***
	(0.41)	(0.39)	(0.41)	(0.41)	(0.41)	(0.40)	(0.39)
R^2_y	0.26	0.50	0.75	0.95	1.35	2.17	3.28
Panel B: Adding the world XVP							
	1	2	3	4	6	9	12
$y_{US}(h) - y_i(h)$	−0.08***	0.07**	0.14**	0.23*	0.13**	−0.02**	−0.06***
	(0.40)	(0.38)	(0.39)	(0.40)	(0.40)	(0.40)	(0.38)
XVP	−11.33***	−11.55***	−12.21***	−11.95***	−9.34***	−4.78***	−2.82***
	(1.70)	(1.35)	(1.19)	(1.19)	(0.93)	(0.66)	(0.57)
R^2	2.43	4.72	7.70	8.95	8.28	5.03	4.79
$R^2 - R^2_y$	2.16	4.22	6.95	8.00	6.94	2.86	1.51

Table 5
The predictive power of VP for exchange rate returns with respect to the U.S. dollar.
This table reports the estimated coefficients for the following panel-data regressions:

$$s_{i,t+h} - s_{i,t} = b_{i,0}(h) + b_{IR}(h)[y_{US,t}(h) - y_{i,t}(h)] + b_{VP}(h)VP_t + u_{i,t+h},$$

where $s_{i,t}$ is the dollar exchange rate of currency i (in units of U.S. dollar per one unit of foreign currency), $y_{US,t}(h) - y_{i,t}(h)$ is the interest rate differential for $h-$month zero-coupon bond rates between the U.S. and country i, and VP is the (one-month) stock variance premium. We consider three alternative measures for the VP: the U.S. stock variance premium (VP_{US}), the equally weighted average stock variance premium (VP_{EW}), and the value-weighted average stock variance premium (VP_{VW}) (see Table 3). To facilitate the interpretation of the estimated coefficients, we divide the VPs by 12. The standard errors are corrected by panel-data Newey-West with h lags (the standard deviations are reported in parentheses). *, **, and *** represent the usual 10%, 5%, and 1% significance levels. For the interest rate differential, $y_{US,t}(h) - y_{i,t}(h)$, the null hypothesis corresponds to $b_{IR} = 1$ (that is, whether the UIP holds). The sample period runs from January 2000 to December 2011. The currency-specific estimated constants are left unreported, to save space. We report the R^2 of the regression and the gains in R^2s with respect to a univariate regression for the interest rate differential, $R^2 - R^2_y$.

Panel A: VP_{US}							
	1	2	3	4	6	9	12
$y_{US}(h) - y_i(h)$	−0.39***	−0.10***	−0.07***	0.01**	0.05**	−0.07***	−0.11***
	(0.39)	(0.38)	(0.38)	(0.39)	(0.41)	(0.40)	(0.39)
VP	2.10***	0.87***	1.07***	1.06***	0.42***	0.04	0.01
	(0.26)	(0.17)	(0.16)	(0.14)	(0.10)	(0.06)	(0.05)
R^2	5.57	2.21	4.58	5.39	2.26	2.19	3.28
$R^2 - R^2_y$	5.31	1.72	3.83	4.44	0.91	0.02	0.00
Panel B: VP_{EW}							
	1	2	3	4	6	9	12
$y_{US}(h) - y_i(h)$	−0.20***	0.00**	0.02**	0.10**	0.07**	−0.07***	−0.11***
	(0.40)	(0.39)	(0.39)	(0.39)	(0.41)	(0.40)	(0.39)
VP	1.48***	0.39**	1.02***	1.03***	0.45***	0.16**	0.09*
	(0.23)	(0.17)	(0.17)	(0.15)	(0.10)	(0.06)	(0.05)
R^2	2.11	0.74	3.19	4.00	2.16	2.34	3.37
$R^2 - R^2_y$	1.85	0.24	2.44	3.05	0.81	0.17	0.08
Panel C: VP_{VW}							
	1	2	3	4	6	9	12
$y_{US}(h) - y_i(h)$	−0.28***	−0.04***	−0.02***	0.06**	0.06**	−0.07***	−0.11***
	(0.39)	(0.38)	(0.39)	(0.39)	(0.40)	(0.40)	(0.39)
VP	1.96***	0.74***	1.11***	1.11***	0.47***	0.11*	0.05
	(0.26)	(0.18)	(0.17)	(0.15)	(0.10)	(0.06)	(0.05)
R^2	4.07	1.51	4.15	5.03	2.35	2.26	3.31
$R^2 - R^2_y$	3.81	1.01	3.40	4.08	1.00	0.09	0.03

Table 6
The predictive power of the world XVP and the U.S. VP for exchange rate returns with respect to the U.S. dollar.
This table reports the estimated coefficients for the following panel-data regressions:

$$s_{i,t+h} - s_{i,t} = b_{i,0}(h) + b_{IR}(h)[y_{US,t}(h) - y_{i,t}(h)] + b_{XVP}(h)XVP_t + b_{VP}(h)VP_{US,t} + u_{i,t+h},$$

where $s_{i,t}$ is the dollar exchange rate of currency i (in units of U.S. dollar per one unit of foreign currency), $y_{US,t}(h) - y_{i,t}(h)$ is the interest rate differential for $h-$month zero-coupon bond rates between the U.S. and country i, XVP is the six-month world currency variance risk premium, and VP_{US} is the U.S. stock variance premium. To facilitate the interpretation of the estimated coefficients, we divide XVP and VP_{US} by 12. The standard errors are corrected by panel-data Newey-West with h lags (the standard deviations are reported in parentheses). *, **, and *** represent the usual 10%, 5%, and 1% significance levels. For the interest rate differential, $y_{US,t}(h) - y_{i,t}(h)$, the null hypothesis corresponds to $b_{IR} = 1$ (that is, whether the UIP holds). The sample period runs from January 2000 to December 2011. The currency-specific estimated constants are left unreported, to save space. We report the R^2 of the regression and the gains in R^2s with respect to a univariate regression for the interest rate differential (Panel A in Table 4), $R^2 - R^2_y$, and with respect to a multivariate regression for the interest rate differential and the U.S. VP (Panel A in Table 5), $R^2 - R^2_{y,VP}$.

	1	2	3	4	6	9	12
$y_{US}(h) - y_i(h)$	−0.33***	−0.02***	0.01***	0.11**	0.12**	−0.02**	−0.05***
	(0.39)	(0.37)	(0.37)	(0.38)	(0.40)	(0.40)	(0.39)
XVP	−8.56***	−10.61***	−10.99***	−10.56***	−9.02***	−4.91***	−2.94***
	(1.73)	(1.39)	(1.18)	(1.15)	(0.95)	(0.72)	(0.63)
VP_{US}	1.93***	0.66***	0.85***	0.80***	0.19*	−0.08	−0.07
	(0.26)	(0.17)	(0.16)	(0.14)	(0.10)	(0.07)	(0.07)
R^2	6.77	5.68	10.05	11.37	8.46	5.09	4.84
$R^2 - R^2_{y,VP}$	1.20	3.46	5.48	5.98	6.20	2.90	1.56
$R^2 - R^2_y$	6.51	5.18	9.30	10.42	7.12	2.92	1.56

power of the VP for h-month ahead appreciation rates remains statistically significant, mostly at the same levels, when equally or value-weighted average VPs are considered instead of the U.S. VP. In independent concurrent work, Aloosh (2014) finds positive evidence of one-month ahead dollar/EUR, dollar/JPY, and dollar/GBP return predictability from the value-weighted average VP.

Adding the VP also yields gains in predictive power with respect to the individual predictive power of the interest rate differential (maximized at 5.31% for the U.S. VP at the one-month horizon). However, as for the XVP, adding the VP does not seem to affect much the deviations from the UIP reported in Panel A of Table 4.

Table 6 shows the (additional) predictive power of the world currency and stock variance risk premiums for h-month ahead appreciation rates. Our results reveal that the world XVP has additional predictive power (after controlling for the U.S. VP) for forex returns for all horizons considered. The estimated XVP coefficient displays an inverted hump-shaped predictability pattern that peaks at the three-month horizon (−10.99). The gains in predictive power after adding XVP with respect to the predictive power of the interest rate differential and the U.S. VP, $R^2 - R^2_{y,VP}$, are maximized at the six-month horizon (6.20%). The U.S. VP also has additional predictive power for horizons between one and six months, and its estimated coefficient is positive and displays a decreasing pattern.[10] Adding VP and XVP simultaneously yields large gains in predictive power with respect to the univariate predictive power of the interest rate differential, $R^2 - R^2_y$, that are maximized at 10.42% at the four-month horizon.

3.2. Heterogeneous exposures to variance risk premiums

We find so far, in a panel-data setting, that an increase in the world XVP predicts an appreciation of the U.S. dollar with respect to foreign currencies, while an increase in VP predicts a depreciation of the U.S. dollar. A natural question at this point is whether the predictive power of variance risk premiums depends on the type of foreign currency considered. We now investigate the differential or heterogeneous predictive power of variance risk premiums for forex returns using an individual-currency regressions setting.

Table 7 reports the predictive power of the world XVP for individual currency appreciation rates.[11] Our results reveal that the coefficient associated with XVP is significant at confidence levels above 10% for 18 of the 22 currencies considered at the four-month horizon. However, the coefficient associated with XVP varies substantially across countries. At the four-month horizon, the coefficient is large (negative and significant) for the ZAR (−24.75), the NZD (−23.90), the HUF (-23.49), and the PLN (−20.64). In contrast, at this horizon, the world XVP is not a useful predictor for appreciation rates of the following currencies: the JPY, the HKD, the MYR, and the PHP. Interestingly, the Malaysian Ringgit, MYR, was pegged to the U.S. dollar between September 1998 and July 2005. On July 21, 2005, the Malaysian monetary authority announced the adoption of a managed float system. The HKD is also pegged at 7.8 to the U.S. dollar, as of May 2005, but can trade between 7.75 and 7.85. The number of currencies for which XVP is significant at any standard confidence level falls to 11 at the 12-month horizon.

The individual-currency regression setting reveals that, except for the Japanese yen, a traditional *funding* currency,

the U.S. lagged stock return is a main predictor for international stock returns.

[10] The decreasing (additional) predictability pattern of the VP is robust to considering the equally weighted or the value-weighted world VP. The results for the alternative VP measures are omitted to save space and are available, upon request, from the authors.

[11] Considering a currency-specific regression setting might be subject to inference problems due to imprecise parameter estimates (Bansal and Dahlquist, 2000).

Table 7
The predictive power of XVP for exchange rate returns with respect to the U.S. dollar, individual-currency regressions.
This table reports the estimated coefficients for the following individual-currency regressions:

$$s_{i,t+h} - s_{i,t} = b_{i,0}(h) + b_{i,IR}(h)[y_{US,t}(h) - y_{i,t}(h)] + b_{i,XVP}(h)XVP_t + u_{i,t+h},$$

where $s_{i,t}$ is the dollar exchange rate of currency i, $y_{US,t}(h) - y_{i,t}(h)$ is the interest rate differential for h−month zero-coupon bond rates between the United States and country i, and XVP is the six-month world currency variance premium (see Table 2). To facilitate the interpretation of the estimated coefficients, we divide XVP by 12. The standard errors are corrected by Newey-West with h lags (the standard deviations are left unreported, to save space). *, **, and *** represent the usual 10%, 5%, and 1% significance levels. The sample period runs from January 2000 to December 2011. The estimated regression constants and coefficients associated with the interest rate differential are also left unreported, to save space. We report the R^2 of the regression and the gains in R^2s with respect to a univariate regression for the interest rate differential, $R^2 - R^2_y$.

	h	1	2	3	4	6	9	12
EUR	XVP	−10.67**	−10.48***	−12.10***	−11.47***	−9.42***	−2.59**	0.86
	R^2	1.89	3.54	7.29	8.21	8.71	1.11	0.20
	$R^2 - R^2_y$	1.88	3.53	7.27	8.17	8.69	1.11	0.19
JPY	XVP	5.23	1.71	−0.30	−2.00	−2.18	0.31	0.88
	R^2	3.05	4.70	7.19	9.66	17.40	34.08	37.69
	$R^2 - R^2_y$	0.58	0.13	0.01	0.34	0.66	0.02	0.24
GBP	XVP	−11.88*	−15.15***	−15.76***	−14.59***	−9.52***	−3.17*	−1.11
	R^2	3.64	11.00	16.78	17.06	12.32	4.20	2.33
	$R^2 - R^2_y$	3.48	10.14	14.92	14.35	8.05	1.34	0.24
CHF	XVP	−5.97	−6.59	−8.02**	−8.12**	−5.66**	−0.61	1.17
	R^2	0.86	2.16	4.14	5.22	5.76	3.16	4.28
	$R^2 - R^2_y$	0.54	1.44	3.46	4.53	3.99	0.09	0.47
AUD	XVP	−21.25***	−20.47***	−20.51***	−19.70***	−15.20***	−9.14***	−6.13**
	R^2	4.76	7.99	11.66	13.13	11.28	6.23	4.51
	$R^2 - R^2_y$	4.72	7.99	11.66	12.87	10.61	6.13	4.51
CAD	XVP	−13.19*	−13.23***	−12.03***	−12.44***	−9.53***	−6.15***	−5.21***
	R^2	4.04	8.51	10.80	15.05	11.77	7.61	8.09
	$R^2 - R^2_y$	3.79	8.06	10.20	14.09	11.08	7.43	8.07
HKD	XVP	−0.12	−0.13	−0.11	−0.06	−0.02	0.04	0.06
	R^2	1.65	1.31	1.02	0.88	0.13	0.77	2.12
	$R^2 - R^2_y$	0.12	0.28	0.33	0.16	0.03	0.24	0.79
SEK	XVP	−14.14*	−15.40***	−15.78***	−16.53***	−13.49***	−5.78**	−3.09
	R^2	2.62	5.79	8.93	11.67	10.24	2.87	1.87
	$R^2 - R^2_y$	2.62	5.78	8.91	11.64	10.22	2.84	1.26
NZD	XVP	−23.28***	−24.20***	−24.85***	−23.90***	−17.30***	−10.02***	−6.28***
	R^2	5.61	11.30	17.80	18.67	13.96	7.48	4.56
	$R^2 - R^2_y$	5.52	11.12	17.44	18.17	13.10	6.54	3.84
KRW	XVP	−19.39	−15.56***	−14.79***	−15.17***	−11.79***	−9.32***	−6.86***
	R^2	5.57	8.11	11.78	14.42	11.17	8.75	5.96
	$R^2 - R^2_y$	5.06	6.63	9.55	11.89	9.57	8.14	5.61
SGD	XVP	−4.83	−4.59*	−5.49***	−5.47***	−3.96***	−1.89*	−1.07
	R^2	2.00	4.05	8.63	9.92	9.37	8.32	10.25
	$R^2 - R^2_y$	1.35	2.66	6.24	7.58	6.23	2.25	1.10
NOK	XVP	−15.88***	−15.44***	−14.60***	−13.56***	−11.54***	−5.88***	−2.62**
	R^2	3.64	6.40	8.18	8.71	9.35	4.25	1.24
	$R^2 - R^2_y$	3.63	6.27	7.94	8.35	8.81	3.72	1.23
INR	XVP	−1.67	−4.58	−4.47	−4.18*	−4.11***	−3.88***	−3.61***
	R^2	8.60	16.36	21.51	24.47	24.24	22.22	18.82
	$R^2 - R^2_y$	0.12	1.52	2.09	2.22	3.09	3.85	4.37
PLN	XVP	−15.54	−15.89***	−19.33***	−20.64***	−17.26***	−8.16***	−4.05**
	R^2	2.71	4.82	9.26	11.73	10.76	4.77	3.33
	$R^2 - R^2_y$	2.17	3.89	8.05	10.60	9.68	3.49	1.58
ZAR	XVP	−31.93***	−29.09***	−27.57***	−24.75***	−16.39***	−11.36***	−9.92***
	R^2	8.20	13.29	17.95	18.71	17.03	22.04	24.00
	$R^2 - R^2_y$	6.66	10.26	13.25	13.11	8.18	6.34	6.16
CZK	XVP	−16.59**	−16.33***	−18.58***	−17.46***	−13.23***	−3.64	−1.39
	R^2	3.25	6.05	11.82	12.98	11.75	4.04	6.88
	$R^2 - R^2_y$	3.12	5.79	11.41	12.20	10.63	1.37	0.36

(*continued on next page*)

Table 7 (*continued*)

	h	1	2	3	4	6	9	12
DKK	XVP	−11.14**	−10.83***	−12.45***	−11.86***	−9.67***	−2.66**	0.84
	R^2	2.03	3.76	7.68	8.79	9.14	1.18	0.20
	$R^2 - R_y^2$	2.03	3.74	7.64	8.68	9.08	1.17	0.18
THB	XVP	−4.13	−4.82**	−4.94***	−4.88***	−4.10**	−2.39**	−2.10**
	R^2	2.42	5.23	7.88	9.51	13.69	14.35	16.34
	$R^2 - R_y^2$	0.95	2.29	3.43	3.99	4.03	2.12	2.41
TWD	XVP	−3.77	−5.08**	−5.20***	−5.11***	−3.95***	−3.27**	−2.13
	R^2	1.15	3.42	5.11	6.15	5.54	6.38	4.09
	$R^2 - R_y^2$	1.14	3.42	5.11	6.14	5.49	6.00	4.00
HUF	XVP	−20.86**	−18.86***	−24.38***	−23.49***	−18.56***	−6.82***	−3.27
	R^2	4.40	6.42	13.37	14.59	13.37	3.00	1.73
	$R^2 - R_y^2$	3.44	5.08	12.38	13.75	12.60	2.88	1.20
MYR	XVP	−2.21	−2.60	−2.61	−2.81	−2.27	−2.07	−2.29*
	R^2	0.91	2.58	3.80	5.16	4.21	4.04	5.32
	$R^2 - R_y^2$	0.41	1.24	2.03	2.84	2.60	3.10	5.26
PHP	XVP	0.11	−0.15	−0.11	−0.46	−1.90	−2.58	−2.39*
	R^2	0.00	0.28	0.63	1.29	3.90	5.04	7.30
	$R^2 - R_y^2$	0.00	0.00	0.00	0.03	0.64	1.59	1.90
Avg. R^2		**3.32**	**6.23**	**9.69**	**11.18**	**10.69**	**8.00**	**7.78**
Avg. $(R^2 - R_y^2)$		**2.42**	**4.60**	**7.42**	**8.44**	**7.14**	**3.26**	**2.50**

the gains in predictive power after adding XVP, $R^2 - R_y^2$, at the four-month horizon are surprisingly high for almost all major free-floating currencies—EUR (8.17%), GBP (14.35%), AUD (12.87%), CAD (14.09%). Gains in R^2s are also very high for the currencies for which we obtain a large estimated coefficient—ZAR (13.11%), NZD (18.17%), HUF (13.75%), and PLN (10.60%). On average, the gain in predictive power with respect to the univariate predictive power of the interest rate differential is maximized at the four-month horizon (8.44%).

Similarly, Table 8 reports the currency-specific predictive power of the U.S. VP. In line with the results for the panel-data setting in Table 5, the U.S. VP is a useful predictor for one-month-ahead appreciation rates for all currencies in our sample except for the EUR, the HKD, the INR, and the PHP. At the five-month horizon, the VP coefficient remains significant for 13 currencies for confidence levels above 10%, and, at the 12-month horizon, the coefficient is insignificant for all currencies at any standard confidence level. The coefficient associated with the U.S. VP also varies across countries, from −1.49 for the JPY to 4.29 for the NZD at the one-month horizon. The average gain in R^2 is maximized at the one-month horizon (6.8%), and is almost null for horizons longer than nine months.

These individual-currency regressions yield two main results. First, by and large, most currencies confirm our main panel regression findings that world XVP and VP have nonredundant short-term predictive power for forex returns. Second, there are large variations in the predictive power of variance risk premiums across individual currencies. In particular, while for the JPY, a traditional funding currency, and currencies pegged to the U.S. dollar, the predictability from XVP and VP are weak or nonexistent, for traditional investment currencies and emerging market currencies, such as the ZAR, the coefficient associated with XVP is largely negative and significant—an increase in XVP is followed by a larger depreciation of these currencies with respect to the U.S. dollar—and the gains in predictive power increase considerably.

3.3. Inflation-sorted currency portfolios

To understand the economic determinants of the heterogeneous predictability patterns across currencies, especially for the world XVP, documented in Section 3.2, in this section, we form currency portfolios based on each country's average inflation (between 2000 and 2011), and investigate the heterogeneous predictive power of XVP and VP across portfolios. The evidence in this section motivates our international general equilibrium model in Section 4, wherein the exposure of each country's inflation process to global inflation risk plays a key role in explaining the cross-currency heterogeneity in the model-implied predictability of XVP.

Table 9 reports the results for the predictive power of the world XVP and the U.S. VP for four-month ahead forex returns for currency portfolios sorted on average inflation. Our results suggest that the coefficient associated with XVP is negative for all portfolios considered, in line with the evidence from the panel-data and individual-currency regression settings. More importantly, the results suggest that this coefficient becomes more negative for the currencies of countries with higher average inflation. In fact, a formal test suggests that the regression coefficient for XVP is significantly larger for the high-inflation currency portfolio, which includes the ZAR, the HUF, the INR, the PHP, and the KRW, than for the low-inflation currency portfolio, which includes the JPY, the CHF, the HKD, and the TWD.[12]

[12] This formal test requires a panel-data setting for both extreme portfolios, wherein the XVP is allowed to interact with a dummy for the high-inflation portfolio. Sorting the portfolios on total 2000 to 2011 inflation

Table 8
The predictive power of U.S. VP for exchange rate returns with respect to the U.S. dollar, individual-currency regressions.
This table reports the estimated coefficients for the following individual-currency regressions:

$$s_{i,t+h} - s_{i,t} = b_{i,0}(h) + b_{i,IR}(h)[y_{US,t}(h) - y_{i,t}(h)] + b_{i,VP}(h)VP_{US,t} + u_{i,t+h},$$

where $s_{i,t}$ is the dollar exchange rate of currency i, $y_{US,t}(h) - y_{i,t}(h)$ is the interest rate differential for $h-$month zero-coupon bond rates between the United States and country i, and VP_{US} is the one-month U.S. stock variance risk premium (VP). To facilitate the interpretation of the estimated coefficients, we divide U.S. VP by 12. The standard errors are corrected by Newey-West with h lags (the standard deviations are left unreported, to save space). *, **, and *** represent the usual 10%, 5%, and 1% significance levels. The sample period runs from January 2000 to December 2011. The estimated regression currency-specific constants and coefficients associated with the interest rate differential are also left unreported to save space. We report the R^2 of the regression and the gains in R^2s with respect to a univariate regression for the interest rate differential, $R^2 - R^2_y$.

	h	1	2	3	4	6	9	12
EUR	VP_{US}	1.94	0.12	0.48	0.64**	0.05	−0.19	−0.15
	R^2	4.46	0.04	0.83	1.83	0.04	0.41	0.39
	$R^2 - R^2_y$	4.45	0.03	0.81	1.79	0.01	0.41	0.37
JPY	VP_{US}	−1.49***	−1.10**	−0.60	−0.22	0.01	0.09	0.04
	R^2	5.91	8.50	8.96	9.62	16.74	34.19	37.48
	$R^2 - R^2_y$	3.44	3.93	1.77	0.30	0.00	0.13	0.04
GBP	VP_{US}	2.43***	1.51***	1.45***	1.28***	0.57**	0.08	0.13
	R^2	10.56	8.04	10.78	10.55	6.16	2.92	2.30
	$R^2 - R^2_y$	10.40	7.17	8.92	7.84	1.89	0.05	0.20
CHF	VP_{US}	2.26**	0.08	0.51	0.65**	0.16	−0.01	0.06
	R^2	5.90	0.74	1.67	2.76	1.99	3.07	3.90
	$R^2 - R^2_y$	5.58	0.02	0.99	2.07	0.22	0.00	0.09
AUD	VP_{US}	3.54**	1.49*	1.62***	1.34***	0.45	−0.17	−0.20
	R^2	9.44	3.04	5.25	4.50	1.27	0.24	0.33
	$R^2 - R^2_y$	9.40	3.03	5.25	4.24	0.60	0.14	0.33
CAD	VP_{US}	2.20**	1.02**	1.14***	0.95***	0.44*	−0.03	−0.08
	R^2	7.82	3.86	7.11	6.81	2.27	0.20	0.14
	$R^2 - R^2_y$	7.57	3.41	6.51	5.86	1.58	0.02	0.13
HKD	VP_{US}	0.00	−0.01	0.00	−0.01	−0.01	0.00	0.00
	R^2	1.55	1.13	0.70	0.82	0.45	0.53	1.35
	$R^2 - R^2_y$	0.02	0.10	0.00	0.10	0.35	0.00	0.01
SEK	VP_{US}	3.14***	1.66***	1.77***	1.73***	0.79***	0.17	0.18
	R^2	9.21	4.81	8.04	9.11	2.32	0.19	0.89
	$R^2 - R^2_y$	9.21	4.80	8.02	9.08	2.31	0.17	0.28
NZD	VP_{US}	4.29***	2.15***	2.29***	2.06***	0.83*	0.03	−0.07
	R^2	13.35	6.37	10.78	9.97	2.79	0.94	0.75
	$R^2 - R^2_y$	13.26	6.19	10.42	9.47	1.93	0.00	0.03
KRW	VP_{US}	3.18***	1.20	1.17***	1.24***	0.32	0.09	0.01
	R^2	10.14	4.28	6.47	8.08	2.06	0.66	0.35
	$R^2 - R^2_y$	9.63	2.80	4.23	5.54	0.46	0.05	0.00
SGD	VP_{US}	1.36*	0.33	0.62***	0.62***	0.27***	0.11**	0.08
	R^2	8.43	2.39	8.11	9.40	5.00	6.62	9.58
	$R^2 - R^2_y$	7.78	1.00	5.72	7.05	1.87	0.55	0.43
NOK	VP_{US}	2.79***	1.22**	1.11***	0.89***	0.23	−0.02	−0.11
	R^2	8.01	2.93	3.51	2.94	0.77	0.53	0.14
	$R^2 - R^2_y$	8.00	2.80	3.27	2.57	0.23	0.00	0.13
INR	VP_{US}	0.82	0.03	0.28	0.40	0.15	−0.08	−0.14
	R^2	10.65	14.84	20.04	23.73	21.40	18.48	14.92
	$R^2 - R^2_y$	2.17	0.00	0.62	1.49	0.26	0.11	0.47
PLN	VP_{US}	4.27***	2.70***	3.18***	3.09***	1.69***	0.39	0.21
	R^2	12.50	9.14	17.09	18.31	7.23	1.82	2.05
	$R^2 - R^2_y$	11.97	8.21	15.87	17.18	6.15	0.55	0.29
ZAR	VP_{US}	3.08**	1.53*	1.60***	1.03**	−0.03	−0.36	−0.25
	R^2	5.93	5.05	7.86	7.19	8.84	16.12	18.10
	$R^2 - R^2_y$	4.39	2.02	3.16	1.60	0.00	0.42	0.26
CZK	VP_{US}	3.34***	1.75***	1.96***	2.03***	0.78***	0.10	0.10
	R^2	9.20	5.07	9.51	12.65	3.53	2.74	6.66
	$R^2 - R^2_y$	9.07	4.80	9.10	11.88	2.41	0.07	0.14

(*continued on next page*)

Table 8 (*continued*)

	h	1	2	3	4	6	9	12
DKK	VP_{US}	1.97*	0.10	0.47	0.63*	0.03	−0.20	−0.15
	R^2	4.51	0.04	0.82	1.82	0.07	0.45	0.40
	$R^2 - R_y^2$	4.51	0.02	0.78	1.71	0.01	0.45	0.38
THB	VP_{US}	1.06***	0.61***	0.57***	0.71***	0.43***	0.23**	0.18
	R^2	5.88	5.55	7.68	11.52	12.53	13.53	15.06
	$R^2 - R_y^2$	4.41	2.62	3.23	6.01	2.87	1.29	1.13
TWD	VP_{US}	1.11**	0.65**	0.70***	0.69***	0.34*	0.21	0.15
	R^2	7.13	4.04	6.53	8.02	2.75	2.08	1.45
	$R^2 - R_y^2$	7.12	4.04	6.53	8.01	2.70	1.70	1.36
HUF	VP_{US}	3.31**	1.66**	2.42***	2.55***	1.11***	0.06	0.06
	R^2	7.66	4.41	10.53	13.56	4.02	0.14	0.56
	$R^2 - R_y^2$	6.69	3.08	9.54	12.73	3.25	0.02	0.03
MYR	VP_{US}	1.23**	0.34	0.58***	0.57***	0.26*	0.16	0.05
	R^2	9.62	2.89	9.11	10.85	3.91	2.15	0.22
	$R^2 - R_y^2$	9.12	1.54	7.35	8.53	2.30	1.22	0.17
PHP	VP_{US}	0.40	0.00	0.22	0.28	0.30	0.18	0.11
	R^2	0.48	0.28	1.02	1.96	4.25	3.98	5.68
	$R^2 - R_y^2$	0.47	0.00	0.39	0.70	0.99	0.53	0.28
Avg. R^2		**7.65**	**4.43**	**7.38**	**8.45**	**5.02**	**5.09**	**5.58**
Avg. $(R^2 - R_y^2)$		**6.76**	**2.80**	**5.11**	**5.72**	**1.47**	**0.36**	**0.30**

Table 9

Heterogeneous predictability patterns of variance risk premiums across inflation-sorted currency portfolios.

This table reports the estimated coefficients for the panel-data regression setting including the interest rate differential, the six-month world XVP, and the U.S. VP (see Table 6) for currency portfolios sorted on country-specific average inflation for the sample running from January 2000 to December 2011. To save space, we only report the results for the four-month prediction horizon. The standard errors are corrected by Newey-West with four lags (the standard deviations are reported in parentheses). *, **, and *** represent the usual 10%, 5%, and 1% significance levels. For the interest rate differential, $y_{US,t}(h) - y_{i,t}(h)$, the null hypothesis corresponds to $b_{IR} = 1$ (that is, whether the UIP holds). The sample period for the regressions runs from January 2000 to December 2011. The estimated constants are left unreported, also to save space. We also report, in the last column, the difference in the estimated coefficients between portfolios 5 and 1 as well as the statistical significance of this difference, which is calculated in a panel-data setting for both extreme portfolios, wherein the right-hand-side variables are allowed to interact with a dummy for the high-inflation portfolio. We report the R^2s of the regression and the gains in R^2s from adding XVP, $R^2_{y,XVP} - R^2_y$, or VP, $R^2_{y,VP} - R^2_y$, to a univariate regression for the interest rate differential.

	Low inflation 1	2	3	4	High inflation 5	5–1
$y_{US}(h) - y_i(h)$	−1.51	0.15	1.78	4.04	2.25	2.25
	(1.10)	(1.88)	(1.99)	(2.58)	(1.30)	(2.41)
XVP	−3.88*	−10.69***	−11.01***	−14.66***	−13.18***	−9.30**
	(2.01)	(2.60)	(2.36)	(2.83)	(2.50)	(4.54)
VP	0.21	0.68**	0.87***	1.37***	0.86***	0.65
	(0.18)	(0.27)	(0.25)	(0.29)	(0.27)	(0.46)
R^2	8.68	12.85	21.26	27.17	20.36	11.68
$R^2_{y,XVP} - R^2_y$	4.54	9.82	18.32	12.92	16.14	11.60
$R^2_{y,VP} - R^2_y$	1.72	3.53	10.96	11.71	7.52	5.80

In other words, high-inflation currencies will depreciate more with respect to the U.S. dollar than low-inflation currencies following an increase in the world XVP. Finally, the results show that the gains in R^2 from adding XVP to the interest rate differential, $R^2_{y,XVP} - R^2_y$, are higher for high-inflation currencies than for low-inflation currencies.

We also find that the coefficient associated with the VP is positive for all currency portfolios, in line with the results for the panel-data and individual-currency regression settings. Although the VP coefficient for high-inflation currencies is higher than that for low-inflation currencies, the difference between these coefficients is not significant. Thus, our results suggest that average inflation does not explain the heterogeneous exposure of future forex returns to the U.S. VP. Nevertheless, as for XVP, the gains in predictive power from adding VP to the interest rate differential, $R^2_{y,VP} - R^2_y$, are higher for high-inflation currencies.

In unreported results, we explore a comprehensive set of variables that could explain the heterogeneous predictability patterns of world currency variance risk premium for appreciation rates against the U.S. dollar. We find that alternative variables characterizing inflation risk, including measures of inflation volatility and inflation exposure to global inflation level and volatility risks, play an insignificant role in explaining the heterogeneous pre-

instead of on average inflation leaves the results for the heterogeneous predictability patterns of XVP unchanged.

dictability patterns of XVP. Interestingly, we also find that variables characterizing each country's real economic growth, including real GDP growth and survey-based real GDP growth uncertainty, do not play a role in explaining the observed short-run heterogeneous predictability patterns.[13] Rather, we find that the current account-to-GDP deficit is the only other fundamental variable that explains the heterogeneous exposure to the world XVP. Specifically, we find that a portfolio formed by the currencies of countries with large relative current account deficits has a significantly larger (more negative) XVP coefficient than a portfolio formed by the currencies of countries with low current account deficits. Furthermore, we find that a set of currency-related variables explains the heterogeneous predictability patterns. In particular, we obtain that currencies with large currency uncertainty (proxied either by the currency-option implied volatility or the carry-to-risk ratio) have significantly larger coefficients associated with the predictive power of the world XVP.

These are all useful angles to explain the variance risk premiums' predictive power for currency returns. However, they may not be mutually exclusive with the inflation differential angle identified earlier, which we will focus on as the main economic direction for the theoretical modeling.

3.4. Additional robustness checks

In this section, we investigate whether our results for the predictive power of the world XVP and the U.S. VP hold for alternative regression specifications, subsamples, and alternative variance risk premium measures. We also investigate whether the predictive power of variance risk premiums holds after controlling for the countercyclical risk premium component of forex returns. To facilitate the comparison between the robustness tests and the benchmark results, we focus on the panel-data setting. Some results in this section are left unreported, to save space, and are available upon request from the authors.

We first consider an alternative regression setting wherein the coefficient associated with the interest rate differential is specific to each currency, while those of the world XVP and the U.S. VP are homogeneous across currencies. This specification does not seem to make any difference for the strong predictive power of currency or stock variance risk premiums. Specifically, in a setting with currency-specific b_{IR}, we find that the estimated coefficients associated with the world XVP are negative and of a similar magnitude as in the benchmark setup. More importantly, the world XVP remains a useful predictor of future appreciation rates for all within-one-year horizons considered. Similarly, the U.S. VP also remains a useful predictor for future appreciation rates with a positive sign but being significant only for short horizons.

To verify the sensitivity of our results to large fluctuations in the world XVP and the U.S. VP around the Lehman Brothers episode (see Figs. 1 and 2), in Table 10, we show the results for our benchmark panel-data setting for a pre-June 2008 sample (or pre-global financial crisis sample). For this subsample, our main results are almost unchanged, suggesting that the predictive power of world currency and stock variance risk premiums is not entirely driven by the global financial crisis. In particular, our main empirical findings do not seem to be affected considerably by the large variance premium spikes observed around this episode. For this subsample, the coefficient associated with the world XVP is negative and significant for all horizons considered and the coefficient of the U.S. VP is positive and significant up to the four-month horizon. If anything, the gains in predictive power when the world XVP and the U.S. VP are added to the interest rate differential are only slightly smaller than those for the full sample (see Table 6).

We also investigate the sensitivity of our results for three alternative variance risk premium measures. In the first alternative measure, the expectation of the forex and stock return variance under the physical distribution is approximated using an AR(1) estimation of their respective realized variance as in Drechsler and Yaron (2011). In the second alternative measure, the expected forex return variance under the risk-neutral measure is approximated using a model-free measure similar to the one used to calculate the VIX (Chicago Board Options Exchange Volatility Index).[14] In the third method, we calculate the forex realized variance using intraday (five-minute) exchange rates for the EUR, the JPY, the CHF, the CAD, the AUD, and the DKK and daily appreciation rates for all other currencies, and calculate the world XVP accordingly. Table 11 shows the predictability of the alternative currency and stock variance premiums for forex returns.

Our main result that both currency and stock variance premiums are useful predictors for future appreciation rates against the U.S. dollar holds for these alternative variance premium measures. For the first alternative measure (Panel A), however, the predictability patterns are slightly changed. In particular, the world XVP is a useful predictor only at horizons between one and four months, while the U.S. VP becomes a useful predictor for all horizons considered.[15] The results for the second alternative measure are almost indistinguishable from those in our

[13] Colacito et al. (2015) find that the currencies of countries with large exposure to long-run growth appreciate relative to those of countries with low exposure following a negative shock to the global economy. In unreported results, we show that, for the advanced-economy currencies in Colacito et al. (2015), the ranking based on the exposure to long-run growth and that based on our inflation risk measures are inversely correlated, although the cross-sectional variation of inflation risk measures decreases substantially for this subsample of currencies. Inflation risk seems to be a more prominent phenomenon for our larger sample of currencies, which includes currencies from both emerging market and advanced economies.

[14] We follow the method in Bakshi and Madan (2000) and Bakshi et al. (2003) to calculate the risk-neutral distribution of each currency's appreciation rate with respect to the U.S. dollar using currency options at different degrees of moneyness. We thank Wenxin Du and Jesse Schreger for kindly providing the code to calculate these risk-neutral distributions.

[15] The first alternative measure differs from the benchmark measure especially around the Lehman Brothers episode. Specifically, the alternative XVP measure is large and positive throughout most of the last quarter of 2008, while the benchmark measure displays a positive spike followed by a large negative spike in October 2008. The alternative U.S. VP also has a negative spike in October 2008, although less pronounced than the spike for the benchmark U.S. VP measure.

Table 10
The predictive power of XVP and VP for exchange rate returns with respect to the U.S. dollar, pre-global financial crisis sample.
This table reports the estimated coefficients for the panel-data regressions:

$$s_{i,t+h} - s_{i,t} = b_{i,0}(h) + b_{IR}(h)[y_{US,t}(h) - y_{i,t}(h)] + b_{XVP}(h)XVP_t^* + b_{VP}(h)VP_{US,t}^* + u_{i,t+h},$$

where $s_{i,t}$ is the dollar exchange rate of currency i, $y_{US,t}(h) - y_{i,t}(h)$ is the interest rate differential for h-month zero-coupon bond rates between the United States and country i, XVP_t is the six-month world XVP, and $VP_{US,\,t}$ is the U.S. VP. The sample period considered runs from January 2000 to June 2008—a few months before the collapse of Lehman Brother in October 2008. To facilitate the interpretation of the estimated coefficients, we divide the world XVP and the U.S. VP by 12. The standard errors are corrected by panel-data Newey-West with h lags (the standard deviations are reported in parentheses). *, **, and *** represent the usual 10%, 5%, and 1% significance levels. For the interest rate differential, $y_{US,t}(h) - y_{i,t}(h)$, the null hypothesis corresponds to $b_{IR} = 1$ (that is, whether the UIP holds). The currency-specific estimated constants are left unreported, to save space. We report the R^2 of each individual regression, and the gains in R^2s with respect to a univariate regression for the interest rate differential, $R^2 - R_y^2$.

	1	2	3	4	6	9	12
$y_{US}(h) - y_i(h)$	−0.13***	0.18**	0.22*	0.31	0.29	0.31	0.30
	(0.42)	(0.41)	(0.42)	(0.42)	(0.45)	(0.47)	(0.48)
XVP	−11.54***	−12.96***	−12.63***	−15.56***	−19.00***	−12.78***	−14.17***
	(2.10)	(1.73)	(1.47)	(1.80)	(1.96)	(1.94)	(1.75)
VP	1.18***	0.44**	0.61***	0.48***	−0.17	0.27**	0.64***
	(0.26)	(0.20)	(0.18)	(0.17)	(0.12)	(0.11)	(0.11)
R^2	5.65	6.79	9.58	11.25	6.80	5.92	9.53
$R^2 - R_y^2$	5.27	6.15	8.64	9.96	4.96	2.87	4.83

benchmark setting, which is not surprising, as the correlation between the second alternative and the benchmark XVPs is 0.90. This result also suggests that there is little gain in using currency options at different degrees of moneyness instead of more simple ATM currency options to calculate the implied volatility of forex returns. Similarly, the results obtained using high-frequency data to calculate forex realized volatilities confirm the evidence from our benchmark setup.

As a final robustness test, we explore the additional predictive power of variance risk premiums for future appreciation rates after controlling for the countercyclical risk premium component of forex returns. To do so, we calculate the U.S.-specific component of global industrial production following Lustig et al. (2014). The results in Table 12 suggest that the predictive power of currency and variance risk premiums is additional to that of the U.S.-specific component of global industrial production. Moreover, the predictability patterns of variance risk premiums are unchanged with respect to the benchmark specification in Table 6. The coefficient associated with the U.S. component of global industrial production is positive and significant for horizons of up to six months, in line with the evidence in Lustig et al. (2014).

To summarize, in this section, we find that the world currency and stock variance risk premiums have predictive power for the appreciation rates of currencies with respect to the U.S. dollar. This evidence is robust to alternative regression specifications and variance risk premium measures, to subsample analysis, and to controlling for the countercyclical component of forex returns. In an individual-currency regression setting, we also find that the predictability patterns of variance risk premiums for forex appreciation rates with respect to the U.S. dollar vary largely across currencies, especially for the world XVP. Using currency portfolios sorted according to country-specific inflation, we find that currencies of countries with high inflation have a higher (more negative) coefficient associated with the predictive power of the world XVP. That is, these currencies tend to depreciate more following an increase in XVP than low-inflation currencies. In addition, the gains in R^2 from adding XVP to the interest rate differential are also higher for high-inflation currencies.

4. A model with global inflation uncertainty

In Section 3, we find that world currency and stock variance risk premiums (XVP and VP) have predictive power for appreciation rates. We also show that currencies' exposures to the world XVP are systematic along the line of inflation risk. To rationalize these empirical findings, in this section, we introduce a two-country consumption-based asset pricing model that links the XVP to global inflation uncertainty.

In our model, both countries' real consumption growth processes are orthogonal while their inflation processes are exposed to global inflation, and this exposure is heterogeneous across countries. Moreover, we allow for shocks to global inflation level to be correlated with shocks to global inflation volatility. The independence of the real-economy components of our model, the heterogeneous exposures to common inflation, and the correlation between inflation level and inflation volatility risks yield the key implications that support our empirical evidence. On the one hand, the XVP implied by the model reveals information about the global inflation uncertainty that cannot otherwise be obtained from domestic stock and stock-options markets. Thus, the XVP contains useful information to explain the time variation of appreciation rates that is additional to the VP. On the other hand, the predictive power of the XVP for the appreciation rate between two currencies depends crucially on the heterogeneity in the exposure of each country's inflation process to global inflation.

In the first part of this section, we explain the model setup and its main implications for the predictive power of variance risk premiums for appreciation rates. In the second part, we compare the model-implied predictability patterns with those observed empirically, and discuss the sensitivity of these patterns to key parameters in the model, including the heterogeneous exposures to global inflation across countries, which explains the empirical evidence for the inflation-sorted currency portfolios.

Table 11
The predictive power of XVP and VP for exchange rate returns with respect to the U.S. dollar, alternative variance premium measures.

This table reports the estimated coefficients for the panel-data regressions:

$$s_{i,t+h} - s_{i,t} = b_{i,0}(h) + b_{IR}(h)[y_{US,t}(h) - y_{i,t}(h)] + b_{XVP}(h)XVP^*_t + b_{VP}(h)VP^*_{US,t} + u_{i,t+h},$$

where $s_{i,t}$ is the dollar exchange rate of currency i, $y_{US,t}(h) - y_{i,t}(h)$ is the interest rate differential for $h-$month zero-coupon bond rates between the United States and country i. We consider three alternative variance risk premium measures (*XVP** and *VP**). In Panel A, *XVP*2 and $VP2_{US}$ are alternative measures for the world currency and U.S. stock variance risk premium in which the expectation of the currency and stock return variance under the physical distribution ($E^P_t(\sigma^2_{c,t+1})$ and $E^P_t(\sigma^2_{r,t+1})$) is approximated using an AR(1) forecast of the realized variance. In Panel B, *XVP*3 is an alternative world XVP measure in which the expectation of the currency return variance under the risk-neutral measure is approximated by a model-free measure using at-the-money and out-of-the-money option prices. The method used to calculate this model-free measure is similar to that used to calculate the VIX, our proxy for the expectation of the stock return variance under the risk-neutral measure. In panel C, to calculate the alternative *XVP*4, we use intraday (five-minute) exchange rates for the EUR, the AUD, the CAD, the DKK, the JPY, and the CHF and daily appreciation rates for all other currencies, and calculate the world XVP accordingly. The intraday data are cleaned using standard techniques. In particular, besides identifying errors in the data, we also determine a threshold for the maximum number of runs of null appreciation rates to exclude quiet trading periods of each day and weekends. To facilitate the interpretation of the estimated coefficients, we divide XVP and the U.S. VP by 12. The standard errors are corrected by panel-data Newey-West with h lags (the standard deviations are reported in parentheses). *, **, and *** represent the usual 10%, 5%, and 1% significance levels. For the interest rate differential, $y_{US,t}(h) - y_{i,t}(h)$, the null hypothesis corresponds to $b_{IR} = 1$ (that is, whether the UIP holds). The sample period runs from January 2000 to December 2011. The currency-specific estimated constants are left unreported, to save space. We report the R^2 of each individual regression, and the gains in R^2s with respect to a univariate regression for the interest rate differential, $R^2 - R^2_y$.

Panel A: XVP2 and VP2 (AR(1) approximation of the physical variance)							
	1	2	3	4	6	9	12
$y_{US}(h) - y_i(h)$	−0.12***	−0.10***	0.01**	0.32	0.45	0.32	0.22*
	(0.42)	(0.41)	(0.42)	(0.43)	(0.46)	(0.46)	(0.44)
XVP2	−9.69***	−6.31***	−6.10***	−3.55***	0.08	2.10**	1.48
	(2.11)	(1.62)	(1.42)	(1.30)	(1.24)	(1.02)	(0.90)
VP2	3.88***	1.84***	2.06***	2.20***	1.58***	0.82***	0.57***
	(0.37)	(0.27)	(0.24)	(0.25)	(0.20)	(0.12)	(0.09)
R^2	7.78	4.14	6.99	7.90	5.49	4.49	4.98
$R^2 - R^2_y$	7.52	3.64	6.24	6.95	4.14	2.32	1.70
Panel B: XVP3 (model-free approximation of the risk-neutral variance)							
	1	2	3	4	6	9	12
$y_{US}(h) - y_i(h)$	−0.16**	0.27	0.39	0.61	0.85	0.85	0.67
	(0.48)	(0.47)	(0.48)	(0.50)	(0.52)	(0.51)	(0.48)
XVP3	−8.90***	−11.88***	−11.41***	−10.70***	−8.84***	−4.46***	−2.18***
	(1.68)	(1.40)	(1.26)	(1.31)	(1.09)	(0.79)	(0.62)
VP	1.93***	0.51***	0.73***	0.66***	−0.04	−0.23***	−0.16**
	(0.28)	(0.19)	(0.17)	(0.15)	(0.12)	(0.08)	(0.07)
R^2	8.33	7.25	11.85	12.63	8.16	4.90	4.18
$R^2 - R^2_y$	8.07	6.75	11.09	11.68	6.81	2.73	0.90
Panel C: XVP4 (intraday exchange rates)							
	1	**2**	**3**	**4**	**6**	**9**	**12**
$y_{US}(h) - y_i(h)$	−0.32	0.01	0.05	0.15	0.20	0.07	−0.03
	(0.39)	(0.37)	(0.38)	(0.38)	(0.39)	(0.41)	(0.40)
XVP4	−7.54***	−9.71***	−10.07***	−9.69***	−9.20***	−5.89***	−2.94***
	(1.69)	(1.36)	(1.16)	(1.14)	(1.04)	(0.79)	(0.65)
VP	1.93***	0.66***	0.85***	0.79***	0.37***	−0.08	−0.10
	(0.26)	(0.17)	(0.16)	(0.14)	(0.12)	(0.08)	(0.07)
R^2	6.56	5.29	9.45	10.68	7.82	4.79	4.66
$R^2 - R^2_y$	6.30	4.80	8.70	9.73	6.47	2.62	1.38

4.1. Model setup and implications

Our model extends the domestic framework of Bollerslev et al. (2009) to an international setting. Specifically, we assume that the real economic growth in the U.S. (the domestic economy in the model) follows the process

$$g_{t+1} = \mu + \phi_l \sigma_{l,t} z_{g_l,t+1}, \tag{5}$$

where the country's macroeconomic uncertainty, $\sigma^2_{l,t}$, is characterized by

$$\sigma^2_{l,t+1} = \mu_l + \rho_l \sigma^2_{l,t} + \phi_{\sigma_l}\sqrt{q_t} z_{\sigma_l,t+1},$$
$$q_{t+1} = \mu_q + \rho_q q_t + \phi_q \sqrt{q_t} z_{q,t+1}.$$

Any other economy (foreign economy hereafter) follows a similar process, with parameters marked with *.

We also assume that each country's representative agent is endowed with recursive preferences (Epstein and Zin, 1989). For simplicity, we assume that the parameters in the preference function are homogeneous across countries. Thus, for instance, the U.S. stochastic discount factor is given by

$$m_{t+1} = \theta \log \delta - \frac{\theta}{\psi} g_{t+1} + (\theta - 1) r_{t+1}, \tag{6}$$

where r_t is the return of an asset that pays the U.S. domestic consumption as dividends (stock return), $0 < \delta < 1$ is the time discount rate, $\gamma \geq 0$ is the risk aversion

Table 12
The predictive power of XVP and VP for exchange rate returns with respect to the U.S. dollar after accounting for the countercyclical risk premium component.

This table reports the estimated coefficients for the panel-data regressions:

$$s_{i,t+h} - s_{i,t} = b_{i,0}(h) + b_{IR}(h)[y_{US,t}(h) - y_{i,t}(h)] + b_{XVP}(h)XVP_t + b_{VP}(h)VP_{US,t} + b_{IP}(h)IP_{UScomp,t} + u_{i,t+h},$$

where $s_{i,t}$ is the dollar exchange rate of currency i, $y_{US,t}(h) - y_{i,t}(h)$ is the interest rate differential for $h-$month zero-coupon bond rates between the United States and country i, XVP_t is the six-month world XVP, $VP_{US,\,t}$ is the U.S. VP. $IP_{US_{comp}}$ is the U.S.-specific component of the world industrial production (IP) growth, which is calculated, as in Lustig et al. (2014), as the residual from the following regression:

$$\Delta IP_{US,t} = \alpha + \beta \frac{\sum_i \Delta IP_{i,t}}{n} + \epsilon_{US\_comp},$$

where the world IP growth, $\frac{\sum_i \Delta IP_{i,t}}{n}$, is calculated using industrial production for the following countries: Austria, Belgium, Canada, Chile, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Israel, Italy, Japan, Korea, Luxembourg, Netherlands, Norway, Poland, Portugal, Slovak Republic, Slovenia, Spain, Sweden, Turkey, the U.K., Brazil, Colombia, India, and Russia. The IP data are obtained from the Organisation for Economic Co-operation and Development (OECD). The sample period considered runs from January 2000 to December 2011. To facilitate the interpretation of the estimated coefficients, we divide the world XVP and the U.S. VP by 12. The standard errors are corrected by panel-data Newey-West with h lags (the standard deviations are reported in parentheses). *, **, and *** represent the usual 10%, 5%, and 1% significance levels. For the interest rate differential, $y_{US,t}(h) - y_{i,t}(h)$, the null hypothesis corresponds to $b_{IR} = 1$ (that is, whether the UIP holds). The currency-specific estimated constants are left unreported, to save space. We report the R^2 of each individual regression, and the gains in R^2s with respect to a univariate regression for the interest rate differential, $R^2 - R^2_y$.

	1	2	3	4	6	9	12
$y_{US}(h) - y_i(h)$	−0.46***	−0.16***	−0.10***	−0.01***	0.00***	−0.07***	−0.04***
	(0.38)	(0.36)	(0.36)	(0.37)	(0.39)	(0.40)	(0.39)
XVP	−8.46***	−10.45***	−10.89***	−10.42***	−8.87***	−4.83***	−2.96***
	(1.71)	(1.38)	(1.18)	(1.15)	(0.95)	(0.72)	(0.63)
VP_{US}	1.78***	0.51***	0.74***	0.69***	0.080	−0.120	−0.060
	(0.26)	(0.18)	(0.16)	(0.14)	(0.11)	(0.08)	(0.07)
$IP_{US_{comp}}$	13.81***	14.10***	10.35***	10.27***	8.97***	3.240	−0.770
	(3.21)	(3.06)	(2.70)	(2.77)	(2.71)	(2.15)	(1.60)
R^2	7.51	7.07	11.13	12.68	9.86	5.37	4.87
$R^2 - R^2_y$	7.25	6.57	10.38	11.73	8.52	3.20	1.59

parameter, and $\theta = \frac{1-\gamma}{1-\frac{1}{\psi}}$ for $\psi \geq 1$ is the intertemporal elasticity of substitution.

To solve the model, as is standard in the literature (Campbell and Shiller, 1988a; 1988b), we log linearize stock returns as

$$r_{t+1} = \kappa_0 + \kappa_1 z_{t+1} - z_t + g_{t+1}, \tag{7}$$

where z_t is the price-consumption ratio. We conjecture a solution for the price-consumption ratio as a function of all state variables as

$$z_{t+1} = A_0 + A_{\sigma_l}\sigma^2_{l,t+1} + A_q q_{t+1}, \tag{8}$$

where the price of risk of the state variables is explained in detail in Appendix A.

As shown by Bollerslev et al. (2009), the variance risk premium of the U.S. stock market implied by our model can be found as the conditional covariance between the variance of stock returns and the domestic stochastic discount factor; that is,

$$VP_t = cov_t(\sigma^2_{r,t+1}, m_{t+1}),$$

where $\sigma^2_{r,t}$ is the conditional variance of stock returns, $var_t(r_{t+1})$. Thus, it can be shown that

$$VP_t = b_{vp,q} q_t, \tag{9}$$

where $b_{vp,q} = (\theta - 1)\kappa_1(A_{\sigma_l}\phi_l^2\phi^2_{\sigma_l} + \kappa_1^2 A_q(A^2_{\sigma_l}\phi^2_{\sigma_l} + A_q^2\phi_q^2)\phi_q^2)$. The model-implied VP in each country is then a function of the country's domestic real growth uncertainty. In particular, the VP is a function of each country's volatility-of-volatility of real consumption growth, q_t. Therefore, the model-implied VP is positive as long as $\theta < 0$ and, following the intuition in Bollerslev et al. (2009), VP becomes a useful predictor for domestic stock returns for horizons for which q_t is the dominant source of variation in the equity premium.

Following Bansal and Shaliastovich (2013) and Zhou (2011), we impose a process for the dynamics of inflation in each country for the model to have realistic implications for nominal appreciation rates.[16] In particular, we assume that the U.S. inflation follows the process

$$\begin{aligned} \pi_{t+1} &= \mu_\pi + \rho_\pi \pi_t + \phi_{\pi\sigma_l}\sigma_{l,t} z_{g_l,t+1} + \phi_{\pi q}\sqrt{q_t} z_{\sigma_l,t+1} \\ &\quad + \phi_\pi \sigma_{\pi,t} z_{\pi,t+1} + \phi_{\pi w}\pi_{w,t+1}, \\ \sigma^2_{\pi,t+1} &= \mu_{\sigma_\pi} + \rho_{\sigma_\pi}\sigma^2_{\pi,t} + \phi_{\sigma_\pi}\sigma_{\pi,t} z_{\sigma_\pi,t+1}, \end{aligned} \tag{10}$$

while the foreign economy's inflation process is given by

$$\begin{aligned} \pi^*_{t+1} &= \mu^*_\pi + \rho^*_\pi \pi^*_t + \phi^*_{\pi\sigma_l}\sigma^*_{l,t} z^*_{g_l,t+1} + \phi^*_{\pi q}\sqrt{q^*_t} z^*_{\sigma_l,t+1} \\ &\quad + \phi^*_\pi \sigma^*_{\pi,t} z^*_{\pi,t+1} + \phi^*_{\pi w}\pi_{w,t+1}, \end{aligned}$$

$$\sigma^{*2}_{\pi,t+1} = \mu^*_{\sigma_\pi} + \rho^*_{\sigma_\pi}\sigma^{*2}_{\pi,t} + \phi^*_{\sigma_\pi}\sigma^*_{\pi,t} z^*_{\sigma_\pi,t+1}.$$

The common exposure of inflation processes to global inflation implies that inflation is correlated across countries.[17] The degree of cross-country inflation correlation

[16] Unlike Bansal and Shaliastovich (2013), in our model, inflation has a neutral effect on real economic growth.

[17] Parameters $\phi_{\pi\sigma_l}$ and $\phi_{\pi q}$ and their foreign counterparts, as shown by Zhou (2011), are crucial to fit each country's term structure of interest rates; however, they do not play a significant role for the predictive power

depends on the differential exposure to global inflation, which is driven by parameters $\phi_{\pi w}$ and $\phi^*_{\pi w}$. The degree of heterogeneity in the exposure to global inflation can be measured as

$$\omega = \frac{\phi^*_{\pi w}}{\phi_{\pi w}}.$$

Interestingly, ω can be expressed as a function of the unconditional mean of both countries' inflations, as follows:

$$\omega = \frac{E(\pi^*_t)(1-\rho^*_\pi) - \mu^*_\pi}{E(\pi_t)(1-\rho_\pi) - \mu_\pi}, \tag{11}$$

which intuitively links heterogeneity in the exposure to global inflation to heterogeneity in average inflation and, therefore, to the empirical evidence for the inflation-sorted currency portfolios in Section 3.3, as we discuss further in Section 4.2.

Global inflation follows the process

$$\pi_{w,t+1} = \mu_{\pi_w} + \rho_{\pi_w}\pi_{w,t} + \phi_{\pi\sigma_w}\sigma_{w,t} z_{\pi_w,t+1}, \tag{12}$$

where

$$\sigma^2_{w,t+1} = \mu_w + \rho_w\sigma^2_{w,t} + \phi_{\sigma_w}\sigma_{w,t} z_{\sigma_w,t+1}$$

is the global inflation uncertainty. We allow for the possibility of correlated shocks between the level and the volatility of global inflation. In particular, we assume $cov_t(z_{\pi_w,t+1}, z_{\sigma_w,t+1}) = \sigma_{\pi_w\sigma_w} = \rho_{\pi_w\sigma_w}$. This assumption implies, in turn, that inflation risk and inflation volatility risk are correlated.[18] To motivate this assumption, in Table 13, we report the correlation coefficient between the level of global inflation and various measures of global inflation uncertainty. Specifically, we calculate several measures of global inflation volatility, including simple measures, such as the absolute value or the square of the volatility, rolling-window estimates of volatility, and time-series based measures of volatility assuming different processes for inflation. We find that the unconditional correlation between global inflation level and volatility is always positive and ranges between 0.49, when global inflation uncertainty is measured as the realized variance of global inflation, and 0.79, when global inflation uncertainty is measured as the absolute value of global inflation.[19] At the country level, however, the correlation between inflation level and inflation volatility is much lower and, depending on the volatility measure, even negative for several countries in our sample.

The XVP, from the point of view of currency investors in the United States, implied by our model is given by:

$$XVP_t = cov_t(m^\$_{t+1}, \sigma^{\$2}_{c,t+1}),$$

where $m^\$_{t+1} = m_{t+1} - \pi_t$ is the nominal stochastic discount factor and $\sigma^{\$2}_{c,t+1}$ is the conditional variance of the nominal exchange rate. Thus,

$$XVP_t = b_{xvp,q}q_t + b_{xvp,\sigma_w}\sigma^2_{w,t}, \tag{13}$$

where

$$b_{xvp,q} = (A_{\sigma_l}\gamma^2\phi_l^2\phi^2_{\sigma_l} + A_q(\theta-1)^2\kappa_1^2(A^2_{\sigma_l}\phi^2_{\sigma_l} + A_q^2\phi_q^2) \times \phi_q^2)(\theta-1)\kappa_1$$

and

$$b_{xvp,\sigma_w} = -\phi_{\pi w}(\phi^*_{\pi w} - \phi_{\pi w})^2\phi_{\sigma_w}\phi^3_{\pi\sigma_w}\rho_{\pi_w\sigma_w}.$$

Comparing the expression for the XVP (Eq. (13)) with that for the VP (Eq. (9)) yields one of the key implications of our model: the XVP reveals information about the global inflation uncertainty, $\sigma^2_{w,t}$, that cannot otherwise be inferred from the purely domestic VPs if the following conditions are met: $\phi_{\pi_w}(\phi^*_{\pi_w}) \neq 0$, $\phi_{\pi_w} \neq \phi^*_{\pi_w}$, and $\rho_{\pi_w\sigma_w} \neq 0$.[20] That is, if (1) country-level inflation is exposed to global inflation, (2) exposures to global inflation are heterogeneous across countries, and (3) shocks to the level and the volatility of global inflation are correlated.

The expected variation in one-period-ahead nominal exchange rates of the foreign currency with respect to the U.S. dollar implied by our model is given by

$$E_t(s_{t+1}) - s_t = E_t(m^\$_{t+1}) - E_t(m^{*\$}_{t+1}) + \frac{1}{2}Var_t(m^\$_{t+1}) - \frac{1}{2}Var_t(m^{*\$}_{t+1}), \tag{14}$$

which is a function of the state variables,

$$E_t(s_{t+1}) - s_t = c_x + b_{x,\sigma_l}\sigma^2_{l,t} + b_{x,\sigma^*_l}\sigma^{*2}_{l,t} + b_{x,q}q_t + b_{x,q^*}q^*_t - \rho_\pi\pi_t + \rho^*_\pi\pi^*_t + (\phi^*_w - \phi_w)\rho_{\pi_w}\pi_{w,t} + \frac{1}{2}(\phi^2_\pi\sigma^2_{\pi,t} - \phi^{*2}_\pi\sigma^{*2}_{\pi,t} + b_{x,\sigma_w}\sigma^2_{w,t}),$$

where

$$b_{x,\sigma_l} = (\theta-1)A_{\sigma_l}\sigma^2_{l,t}(\kappa_1\rho_l - 1) + \frac{1}{2}(b_{mg} + b_{mr})^2\phi_l^2,$$

$$b_{x,q} = (\theta-1)A_q(\kappa_1\rho_q - 1) + \frac{1}{2}((\theta-1)\kappa_1 A_{\sigma_l}\phi_{\sigma_l} - \phi_{\pi q})^2 + \frac{1}{2}(\theta-1)^2\kappa_1^2A_q^2\phi_q^2,$$

and

$$b_{x,\sigma_w} = (\phi^2_{\pi w} - \phi^{*2}_{\pi w})\phi^2_{\pi\sigma_w}.$$

Comparing the expression for the expected nominal appreciation rates (Eq. (14)) with the expressions for the variance risk premiums (Eqs. (13) and (9) for XVP and VPs, respectively) yields the implication of our model for the predictive power of variance risk premiums for appreciation rates. On the one hand, XVP and VPs should contain useful

of XVP for appreciation rates. Therefore, for the calibration of the model in Section 4.2, we assume $\phi_{\pi\sigma_l} = \phi_{\pi q} = 0$.

[18] This assumption also implies that, because inflation risk premium and inflation volatility risk premium cannot be separately identified, there are no major gains in reducing the model's parsimony by adding time-varying inflation volatility-of-volatility. We do entertain the possibility of having time-varying volatility-of-volatility for real GDP growth mainly to be consistent with the existing literature. In particular, because time-varying volatility-of-volatility allows us to differentiate the volatility risk premium from the consumption risk premium (see, for instance, Tauchen, 2011; Bollerslev et al., 2009).

[19] In contrast, the related literature and the empirical evidence suggest that real growth and growth volatility are not tightly correlatedÂ or, if anything, slightly negatively correlated (see, for instance, Bansal et al., 2014; Bansal and Yaron, 2004).

[20] Moreover, $\phi_{\pi\sigma_w}$ and ϕ_{σ_w} should also be different from zero, which trivially means that global inflation level and volatility risks are nonzero.

information to predict exchange rate returns. On the other hand, the predictive power of our model's implied XVP is additional to that of the VP as long as $(\phi_{\pi w}^2 - \phi_{\pi w}^{*2}) \neq 0$; that is, as long as the exposure of both countries' inflation processes to the global inflation uncertainty is heterogeneous ($\omega \neq 1$, see Eq. (11)).[21] The additional predictive power of XVP should become more relevant for horizons at which the global inflation uncertainty dominates the domestic sources of uncertainty in explaining the expected appreciation rate.

4.2. Model-implied predictability patterns

In this section, we illustrate our model's ability to generate predictability patterns that are qualitatively comparable to those suggested by the empirical evidence in Section 3. In particular, we show that the model-implied slope coefficients for the predictive power of stock and currency variance risk premiums for appreciation rates and the (univariate-regression) coefficients of determination linked to these variance risk premiums qualitatively match the observed patterns. We also explore the sensitivity of these predictability patterns to two important economic parameters in our model: the heterogeneous exposure to global inflation and the correlation between global inflation level and volatility shocks. We show that the former parameter is key to understand the predictability patterns observed for the country-specific regressions and for the inflation-sorted currency portfolios, in Sections 3.2 and 3.3, respectively.

The model-implied slope coefficients for the predictive power of stock and currency variance risk premiums for h-month ahead appreciation rates are given by

$$\beta_{x,VP}(h) = \frac{cov(s_{t+h} - s_t, VP_t)}{var(VP_t)}, \tag{15}$$

and

$$\beta_{x,XVP}(h) = \frac{cov(s_{t+h} - s_t, XVP_t)}{var(XVP_t)}, \tag{16}$$

respectively. The coefficients of determination are given by

$$R^2_{x,VP}(h) = \frac{cov(s_{t+h} - s_t, VP_t)^2}{var(VP_t)var(s_{t+h} - s_t)}, \tag{17}$$

and

$$R^2_{x,XVP}(h) = \frac{cov(s_{t+h} - s_t, XVP_t)^2}{var(XVP_t)var(s_{t+h} - s_t)}, \tag{18}$$

for a regression wherein either the stock or the currency variance risk premium is considered, respectively. The components of Eqs. (15) to (18) are presented in Appendix B.

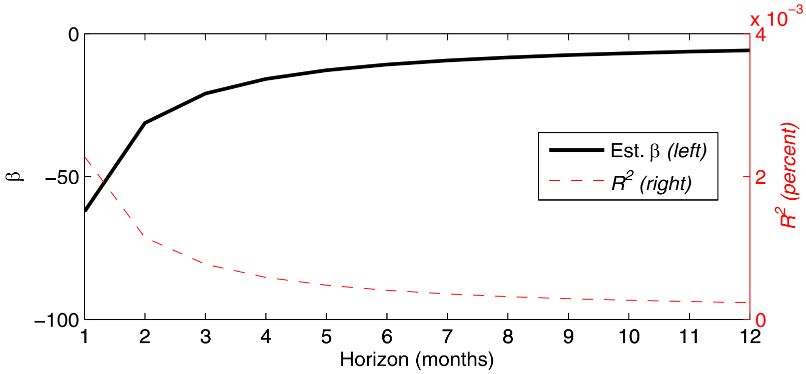
The numerical values for the components of the model-implied slope coefficients and coefficients of determination depend upon the values of the parameters that characterize the local and foreign real economic growth processes (Eq. (5) and its foreign counterpart), the parameters driving the inflation processes (Eq. (10) and its foreign counterpart), and the parameters of the preference function (Eq. (6)). In Appendix C, we explain in detail the method used to calibrate the parameters in the model with real growth, inflation, and XVP data for the United States and the United Kingdom.

In Fig. 3, we compare the observed and model-implied predictability patterns of variance risk premiums for the dollar-pound appreciation rate for the benchmark set of estimated parameters. The model-implied coefficient for the predictive power of the dollar-pound XVP for the dollar-pound appreciation rate, $\beta_{x,XVP}(h)$, is negative and decreases (approaches to zero) with the horizon (Panel A). That is, our model implies that an increase in the dollar-pound variance risk premium, which reveals information about the global inflation uncertainty, is followed by the appreciation of the U.S. dollar with respect to the pound for all horizons considered. The R^2 from a univariate regression with XVP decreases with the horizon and its magnitude is several orders of magnitude smaller than those observed empirically. The model-implied coefficient associated with the VP, $\beta_{x,VP}(h)$, is positive and decreases with the horizon (Panel B). Thus, in line with the empirically observed coefficient, an increase in U.S. VP, which reveals information about domestic real economic uncertainty, is followed by a depreciation of the U.S. dollar with respect to the U.K. pound. The R^2 for a univariate regression with VP follows a hump-shaped pattern that peaks at the five- to six-month horizon, although, as for XVP, the R^2s are several orders of magnitude smaller than those observed empirically.

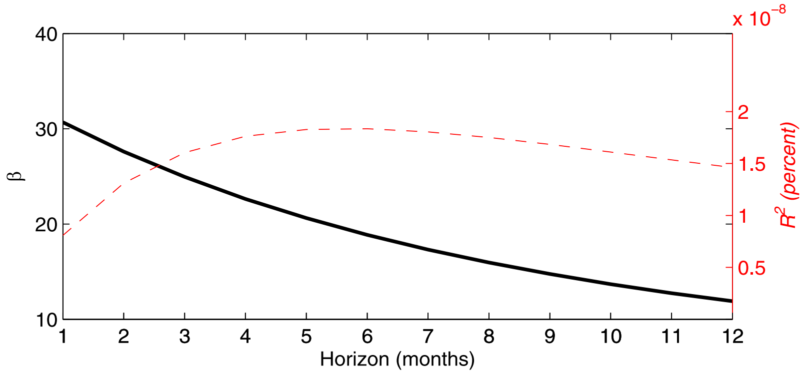
In Fig. 4, we focus on the sensitivity of the predictive power of XVP for appreciation rates to ω, the degree of heterogeneity in the exposure of inflation to global inflation across countries (see Eq. (11)). When the U.S. is assumed to be more exposed to global inflation than the foreign economy, that is, when $\phi_{\pi w} > \phi^*_{\pi w}$ $(w < 1)$, the model-implied coefficient associated with the XVP becomes positive. Thus, an increase in the dollar-pound variance risk premium predicts a depreciation of the U.S. dollar, in contrast to our empirical evidence in Table 7 for most currencies, except perhaps for the JPY and other hard-pegged currencies, such as the HKD. However, as long as $w > 1$ and, therefore, $\phi^*_{\pi w} > \phi_{\pi w}$, an increase in the dollar-pound variance risk premium predicts an appreciation of the U.S. dollar for all horizons considered, which is consistent with our benchmark panel regression results.[22] This finding suggests that, in line with the evidence in Section 3.3, the currencies of countries with higher aver-

[21] The relevance of having heterogeneous exposures to the common factor is acknowledged in Backus et al. (2001), Farhi et al. (2015), Lustig et al. (2011), Gourio et al. (2013), and, in a no-arbitrage setting, in Lustig et al. (2014). The global-uncertainty component in Bansal and Shaliastovich (2013) and Du (2013) cancels out in the expression for the expected appreciation rate precisely because of the homogeneous exposures of both countries to this factor.

[22] We obtain a range for ω using the ratio of average inflations in Eq. (11). For the countries in our sample, the minimum ω is −0.1 for Japan, and there are four countries with ω above 2.0: the Philippines (2.0), South Africa (2.4), Hungary (2.3), and India (2.8). For very high values of ω, however, the model-implied predictability patterns, although still negative, are not necessarily increasing in ω (that is, a higher exposure implies that currencies will depreciate more following an increase in XVP). This result is in line with our empirical evidence for portfolios sorted on inflation regarding the extreme portfolios 4 and 5 in Table 9.



Panel A. Dollar-pound XVP



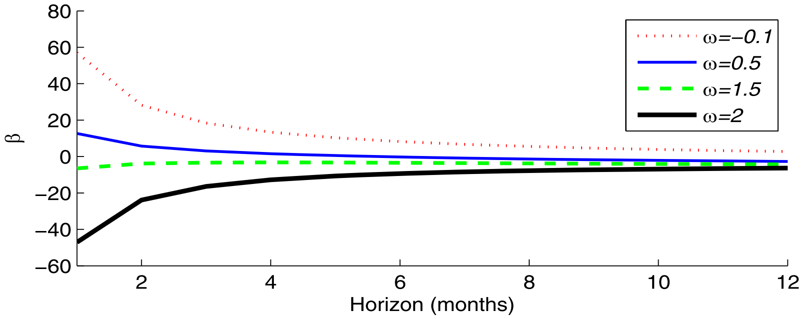
Panel B. U.S. VP

Fig. 3. Model-implied predictability of dollar-pound XVP and U.S. VP for the dollar-pound appreciation rate. The figure shows the model-implied regression coefficients for the predictive power of the dollar-pound variance risk premium (XVP) and the U.S. stock variance premium (VP) for the dollar-pound appreciation rate in Panels A and B, respectively. We also report the R^2s from a univariate regression with either XVP or VP. The model-implied predictability patterns and the parameters used in the calibration of the model are explained in Section 4.2 and Appendix B and Appendix C.

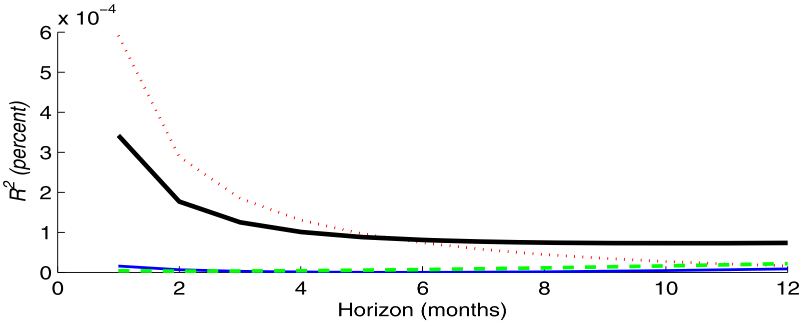
age inflation are more exposed to global inflation and will therefore depreciate more with respect to the currencies of lower-inflation countries following an increase in XVP. The economic intuition for this finding is as follows. As inflation in the United States is less exposed to global inflation risk, an increase in XVP should lead to a higher demand to hold safe assets. The increase in the demand for safe assets will make the U.S. dollar value relatively higher going forward. As can be seen in Panel B, although the model-implied R^2s are several orders of magnitude smaller than those observed empirically, these R^2s increase as the exposure to global inflation uncertainty is more heterogeneous, that is, as w deviates more from one in either direction.

Fig. 5 shows the sensitivity of the model-implied predictability patterns of XVP for future appreciation rates to $\rho_{\pi_w \sigma_w}$, the parameter driving the correlation between shocks to the level and the volatility of global inflation. If shocks to the level and the volatility of global inflation are orthogonal, that is, $\rho_{\pi_w \sigma_w} = 0$, the model-implied $\beta_{c,XVP}$ is relatively small and positive and the R^2s are virtually zero. The more correlated shocks to inflation and inflation volatility are, the larger (more negative given that $\omega > 1$ for the benchmark calibration) the model-implied $\beta_{c,XVP}$ and the higher the explanatory power of XVP for appreciation rates. This result echoes the motivating evidence that the systematic components of all countries' inflation level and volatility shocks are strongly and positively correlated, as shown in Table 13.

Our model can qualitatively match the predictive pattern of XVP for appreciation rates, but quantitatively does not match well many important asset pricing and macroeconomic moments (see Appendix C). To further improve on the quantitative dimension, there are several possible extensions, which we leave for further research. First, we could introduce long-run growth risk and the cross-country correlations in growth and growth volatility risk (Colacito and Croce, 2013; Colacito et al., 2015), which would complement our setup with cross-country correlation in inflation and inflation volatility and help to match important real quantities and equity premiums. Second, we could also allow for richer interactions between growth and inflation dynamics, therefore introducing money nonneutrality (see, for instance, Bansal and Shaliastovich, 2013), which would help us to match the moments of nominal interest rates and further enrich our model implications for joint dynamics of inflation, currency appreciation rates, and XVP. Finally, there are some differential effects between emerging and advanced countries, especially along the interest differential dimension, which mirrors the findings in Bansal and Dahlquist (2000). Therefore, it would be natural to further model the countries' heterogeneous exposures to global inflation risk as

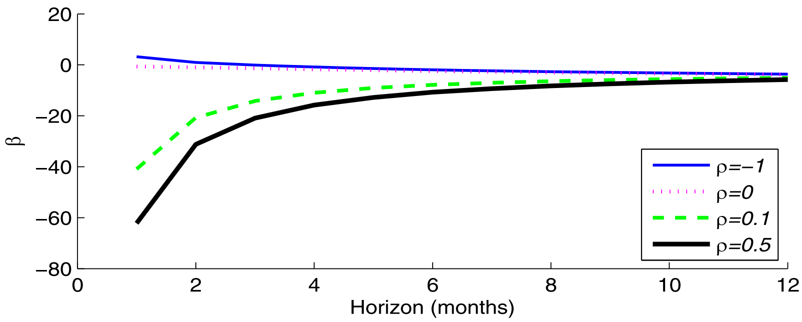


Panel A. $\beta_{x,XVP}$

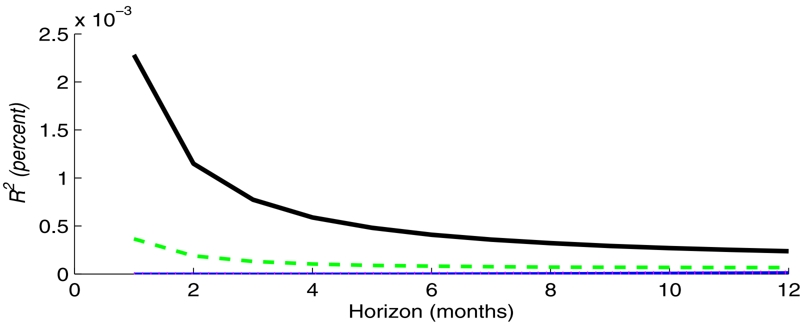


Panel B. $R^2_{x,XVP}$

Fig. 4. Sensitivity of the model-implied predictability of XVP to cross-country heterogeneity in the exposure to global inflation uncertainty. The figure shows the regression coefficients and R^2s for the predictive power of the dollar-pound variance risk premium for the dollar-pound appreciation rate implied by our model for alternative values of ω, the parameter driving the relative exposure of each country's inflation process to the global inflation uncertainty (see Eq. (11)). The predictability patterns are calculated using the parameters and expressions in Section 4.2 and Appendix B and Appendix C.



Panel A. $\beta_{x,XVP}$



Panel B. $R^2_{x,XVP}$

Fig. 5. Sensitivity of the model-implied predictability of XVP to the correlation between the global inflation level and volatility. The figure shows the regression coefficients and R^2s for the predictive power of the dollar-pound variance risk premium for the dollar-pound appreciation rate implied by our model for alternative values of $\rho_{\pi_w\sigma_w}$, the parameter driving the correlation between the level and the volatility of global inflation. The predictability patterns are calculated using the parameters and expressions in Section 4.2 and Appendix B and Appendix C.

Table 13
Correlation between inflation level and inflation uncertainty.

This table reports the unconditional correlation coefficient between the level of inflation and alternative measures of inflation uncertainty. The first measure of inflation uncertainty is the absolute value of global inflation. The second measure is the square of inflation. Rolling RV is the realized variance of monthly inflation calculated using nonoverlapping annual windows as the sum of the squared monthly inflation. Rolling RVol is the realized volatility calculated as the squared-root of the realized variance. The last measure is the time-series inflation uncertainty measure calculated as the volatility of the following inflation process proposed by Stock and Watson (2007): $\pi_t = \tau_t + \eta_t$, where $\eta_t \sim N(0, \sigma^2_{\eta,t})$, and $\tau_t = \tau_{t-1} + \epsilon_t$ is inflation's stochastic trend with $\epsilon_t \sim N(0, \sigma^2_{\epsilon,t})$. The volatilities of the permanent and noise components of inflation follow $\log(\sigma^2_{\eta,t}) = \log(\sigma^2_{\eta-1,t}) + \psi_{1,t}$ and $\log(\sigma^2_{\epsilon,t}) = \log(\sigma^2_{\epsilon-1,t}) + \psi_{2,t}$, respectively, where $\psi_t = (\psi_{1,t}, \psi_{2,t})'$ is iid (independent and identically distributed)$N(0, I_2)$. The total volatility of each country's inflation process is calculated as $\sigma_t = \sqrt{\sigma^2_{\eta,t} + \sigma^2_{\epsilon,t}}$. The column labeled "Global inflation" shows the unconditional correlation coefficient between global inflation, calculated as the equally weighted average of all countries' inflation, and each alternative inflation uncertainty measure. The column labeled "Cross-country average" shows the cross-country average of the correlation between country-level inflation and each measure of inflation uncertainty. All inflation measures are calculated for the sample running from January 2000 to December 2011.

	Global inflation	Cross-country average
Absolute value of inflation	0.79	0.02
Square of inflation	0.73	0.02
Rolling RV (12 months)	0.49	0.12
Rolling RVol (12 months)	0.59	0.12
Stock-Watson inflation uncertainty	0.56	0.18

"bipolar" for the advanced and emerging countries, which could potentially help to match some key dimensions in international finance moments.

5. Conclusion

The pervasive violations of the UIP, especially for short horizons, originally documented in Fama (1984), imply that there is a time-varying predictable component in the currency risk premium. In this paper, we provide new empirical evidence that the currency and stock variance risk premiums (XVP and VP) are useful predictors of future appreciation rates with respect to the U.S. dollar for 22 currencies.

We propose a measure for the world XVP as the average of 17 currencies' variance risk premiums. We show that the world XVP predicts currency *depreciation* against the U.S. dollar, especially at the short within-year horizon. The estimated world XVP coefficient displays an inverted hump-shaped predictability pattern, and the gains in predictive R^2s reach a maximum of 8% at the four-month horizon. We also document a finding that the U.S. VP can predict *appreciation* rates with respect to the U.S. dollar for the 22 currencies considered, especially at the one-month horizon, where the gains in predictive R^2s are maximized at 5.3%. Interestingly, XVP and VP have different informational content for future exchange rate returns and are not highly correlated with each other.

We also find evidence of heterogeneous forex predictability patterns across currencies and systematic exposure to inflation risk. In particular, we sort currencies into portfolios and find that the currencies of countries with high inflation depreciate more following an increase in XVP than low-inflation currencies. These findings motivate a two-country consumption-based asset pricing model, wherein both countries' real consumption growth dynamics are orthogonal to each other, while both countries' inflation processes are exposed to common global inflation. The currency variance risk premium implied by our model isolates the global inflation uncertainty as long as the exposures of the two countries to the global inflation uncertainty are not homogeneous and shocks to global inflation level and volatility are correlated. The model-implied stock variance risk premium for each country captures the domestic real consumption uncertainty, or volatility-of-volatility component. Therefore, XVP and VP have different informational content for the appreciation rates of currencies against the U.S. dollar, both in theory and empirically. The predictability pattern of XVP for appreciation rates depends crucially on the heterogeneity in the exposure to global inflation. In particular, the currencies of countries with higher exposure to global inflation will depreciate with respect to the currencies of low-exposure countries following an increase in XVP, which explains the empirical evidence for the inflation-sorted currency portfolios.

Appendix A. Solution to the price-consumption ratio

As is standard in the literature, we solve the model in Section 4 by log linearizing domestic stock returns following Campbell and Shiller (1988b) as

$$r_{t+1} = \kappa_0 + \kappa_1 z_{t+1} - z_t + g_{t+1}. \tag{A.1}$$

We then propose a process for the log of the wealth-consumption ratio of the asset that pays the consumption endowment in terms of the state variables (Eq. (8) written here again for completeness), that is,

$$z_{t+1} = A_0 + A_{\sigma_l}\sigma^2_{l,t+1} + A_q q_{t+1}. \tag{A.2}$$

Finally, we impose the general equilibrium condition $E_t(r_{t+1} + m_{t+1}) + \frac{1}{2}Var_t(r_{t+1} + m_{t+1}) = 0$. The solution yields

$$A_0 = \frac{(1-\gamma)\mu + \theta\log\delta + \theta\kappa_0 + \theta\kappa_1(A_{\sigma_l}\mu_l + A_q\mu_q)}{\theta(1-\kappa_1)}, \tag{A.3}$$

$$A_{\sigma_l} = \frac{(1-\gamma)^2\phi_l^2}{2\theta(1-\kappa_1\rho_l)}, \tag{A.4}$$

and

$$A_q^{\pm} = \frac{(1-\kappa_1\rho_q) \pm \sqrt{(1-\kappa_1\rho_q)^2 - \theta^2\kappa_1^4\phi_q^2\phi_{\sigma_l}^2 A_{\sigma_l}^2}}{\theta\kappa_1^2\phi_q^2}. \tag{A.5}$$

To avoid the load of time-varying domestic volatility-of-volatility, q_t, from growing without bounds, it only makes sense to keep A_q^-. The positive root is discarded as it is explosive in ϕ_q, that is, $\lim_{\phi_q \to 0} A_q^+ \phi_q \neq 0$. Also, A_q^- will be the solution to the model as long as $(1-\kappa_1\rho_q)^2 \geq \theta^2\kappa_1^4\phi_q^2\phi_{\sigma_l}^2 A_{\sigma_l}^2$. It is easy to show from these expressions that A_{σ_l} , $A_q \leq 0$ as long as $\theta < 1$.

Appendix B. Solution to prediction R^2s and slope coefficients

We now describe how to obtain the components of Eqs. (15) to (18). The model-implied h-period ahead exchange rate return can be approximated by the compound return based on monthly appreciation rates as follows:

$$\begin{aligned}\frac{1}{h}(s_{t+h}-s_t) &\simeq \frac{1}{h}\sum_{j=1}^{h}(s_{t+h}-s_t)\\ &= \frac{1}{h}\Bigg[c_{x,h}+b_{x,\sigma_l}\left(\frac{1-\rho_l^h}{1-\rho_l}\right)\sigma_{l,t}^2+b_{x,\sigma_l^*}\left(\frac{1-\rho_l^{*h}}{1-\rho_l^*}\right)\sigma_{l,t}^{*2}\\ &\quad+b_{x,q}\left(\frac{1-\rho_q^h}{1-\rho_q}\right)q_t+b_{x,q^*}\left(\frac{1-\rho_q^{*h}}{1-\rho_q^*}\right)q_t^*\\ &\quad-\rho_\pi\frac{1-\rho_\pi^h}{1-\rho_\pi}\pi_t+\rho_\pi^*\frac{1-\rho_\pi^{*h}}{1-\rho_\pi^*}\pi_t^*\\ &\quad+b_{x,\pi_w}\pi_{w,t}+f_c(z_{y,t+1},..z_{y,t+h})\Bigg],\end{aligned} \tag{B.1}$$

where $c_{x,h}$ is a constant term,

$$\begin{aligned}b_{x,\sigma_l} &= (\theta-1)b_{r,\sigma_l},\\ b_{x,\sigma_l^*} &= -(\theta-1)b_{r^*,\sigma_l^*},\\ b_{x,q} &= (\theta-1)b_{r,q},\\ b_{x,q^*} &= -(\theta-1)b_{r^*,q^*},\\ b_{x,\pi_w} &= \rho_{\pi_w}\Bigg(\frac{\phi_w^*\rho_\pi^*}{\rho_\pi^*-\rho_{\pi_w}}\left(\frac{1-\rho_\pi^{*h}}{1-\rho_\pi^*}-\frac{1-\rho_{\pi_w}^h}{1-\rho_{\pi_w}}\right)\\ &\quad-\frac{\phi_w\rho_\pi}{\rho_\pi-\rho_{\pi_w}}\left(\frac{1-\rho_\pi^h}{1-\rho_\pi}-\frac{1-\rho_{\pi_w}^h}{1-\rho_{\pi_w}}\right)\\ &\quad+(\phi_w^*-\phi_w)\frac{1-\rho_{\pi_w}^h}{1-\rho_{\pi_w}}\Bigg),\end{aligned}$$

and $b_{r,q}$ and b_{r,σ_l} are the stock return loads on the state variables $q_{l,t}$ and $\sigma_{l,t}$, respectively,

$$\begin{aligned}b_{r,q} &= (\kappa_1\rho_l-1)A_{\sigma_l},\\ b_{r,\sigma_l} &= (\kappa_1\rho_q-1)A_q.\end{aligned}$$

The model-implied one-month ahead VP is defined in Eq. (9). From this expression, the components of $\beta_{x,VP}$ and $R^2_{x,VP}$ are given by

$$\begin{aligned}&cov\Big(\frac{1}{h}\sum_{j=1}^{h}(s_{t+j}-s_{t+j-1}),VP_t\Big)\\ &=\frac{1}{h}b_{vp,q}b_{x,q}\left(\frac{1-\rho_q^h}{1-\rho_q}\right)var(q_t)\end{aligned}$$

and

$$var(VP_t)=b_{vp,q}^2var(q_t).$$

The T-month ahead XVP is given by

$$\begin{aligned}XVP_t(T) &\approx \frac{1}{T}\sum_{j=1}^{T}XVP_{t+j}=\Bigg[b_{xvp,q}q_t\left(\frac{1-\rho_q^T}{1-\rho_q}\right)\\ &\quad+b_{xvp,\sigma_w}\sigma_{w,t}^2\left(\frac{1-\rho_w^T}{1-\rho_q}\right)+f_{xvp}(z_{t+1},..z_{t+T})\Bigg],\end{aligned} \tag{B.2}$$

where $b_{xvp,\,q}$ and b_{xvp,σ_w} are defined in Eq. (13). Therefore, the components of $\beta_{x,XVP}$ and $R^2_{x,XVP}$ are given by the following expressions:

$$\begin{aligned}&cov\left(\frac{1}{h}\sum_{j=1}^{h}(s_{t+j}-s_{t+j-1}),XVP_t(T)\right)\\ &=TCov(c_{t+1},XVP_{t+1})+\sum_{j=1}^{T-1}(T-j)cov(c_{t+1},XVP_{t+j+1})\\ &\quad+\sum_{j=1}^{h-1}(h-j)cov(c_{t+1+j},XVP_{t+1}),\end{aligned}$$

where

$$\begin{aligned}&Cov\left(\frac{1}{h}\sum_{j=1}^{h}(s_{t+j}-s_{t+j-1}),XVP_{t+1}\right)=b_{xvp,q}b_{xq}var(q_t),\\ &Cov(s_{t+1}-s_t,XVP_{t+j+1})=b_{xvp,q}b_{cq}\rho_q^jVar(q_t)\\ &\quad+b_{xvp,q}(\theta-1)\kappa_1A_q\rho_q^{j-1}\phi_q^2E(q_t)\\ &\quad+b_{xvp,\sigma_w}\phi_{\sigma_w}\rho_w^{j-1}(\phi_w^*-\phi_w)\phi_{\pi\sigma_w}\rho_{\pi\sigma_w}E(\sigma_{w,t}),\end{aligned}$$

and

$$cov\left(\frac{1}{h}\sum_{j=1}^{h}(s_{t+j}-s_{t+j-1}),XVP_{t+1}\right)=b_{xvp,q}b_{xq}\rho_q^jvar(q_t).$$

Finally, the unconditional first- and second-order moments of the state variables and global inflation uncertainty can be found as follows:

$$\begin{aligned}E(q_t) &= \frac{\mu_q}{1-\rho_q};\ E(q_t^*)=\frac{\mu_q^*}{1-\rho_q^*};\\ E(\sigma_{l,t}^2) &= \frac{\mu_l}{1-\rho_l};\ E(\sigma_{l,t}^{*2})=\frac{\mu_l^*}{1-\rho_l^*};\ E(\sigma_{w,t}^2)=\frac{\mu_w}{1-\rho_w};\\ var(q_t) &= \frac{\phi_q^2E(q_t)}{1-\rho_q^2};\ var(q_t^*)=\frac{\phi_q^{*2}E(q_t^*)}{1-\rho_q^{*2}};\\ var(\sigma_{l,t}^2) &= \frac{\phi_{\sigma_l}^2E(q_t)}{1-\rho_l^2};\ var(\sigma_{l,t}^{*2})=\frac{\phi_{\sigma_l}^{*2}E(q_t^*)}{1-\rho_l^{*2}};\\ var(\sigma_{w,t}^2) &= \frac{\mu_w}{1-\rho_w}.\end{aligned}$$

Appendix C. Calibration of the model

In this section, we describe the method used to calibrate the parameters for the model in Section 4.

For the benchmark scenario, we calibrate the parameters for the real consumption growth processes (Eq. (5) and its foreign counterpart) to mimic the U.S. economy and the U.K. economy. In particular, we assume $\mu=0.18\%$ and $\mu^*=0.07\%$, equivalent to the average monthly industrial production growth for each country, respectively, for a sample period running from 1970 to 2011. For simplicity, we assume that all other parameters driving the real consumption growth volatility in each country are homogeneous. To calibrate the parameters driving the dynamics of local uncertainties, we follow Bollerslev et al. (2009) and set $\rho_l=\rho_l^*=0.979$. We also set $\phi_{\sigma_l}=\phi_{\sigma_l}^*=0.2<1$ to reduce the chance of finding nonreal solutions for the model (see Appendix A). To calibrate the parameters driving the dynamics of the volatility-of-volatility, we also follow Bollerslev et al. (2009) and set $\rho_q = \rho_q^* = 0.80$, $\mu_q = \mu_q^* = 1 \times 10^{-6}(1-\rho_q)$, and $\phi_q = \phi_q^* = 0.001$.

Campbell and Shiller's (1988b) constants, κ_0 and κ_1 (and their foreign counterparts), are estimated using an iterative procedure, as they depend on the parameters of the real component of the model. Specifically, we depart from initial values of κ_0 and κ_1 that match the unconditional mean of the industrial production growth of the U.S. and the U.K. between 1970 and 2011 (as in Londono, 2015). Given these initial values, we then find the parameters in the price-consumption ratio (see Appendix A in the paper), obtain new values for the constants given these parameters, and iterate until the sum of the absolute changes in the estimated Campbell and Shiller constants are below a tolerance level (1×10^{-6}).

To calibrate the preference-function parameters (Eq. (6)), we follow Bansal and Yaron (2004) and Bollerslev et al. (2009) and set $\delta = 0.997$, $\gamma = 10$, and $\psi = 1.5$.

To calibrate the parameters in each country's inflation processes (Eq. (10) and its foreign counterpart) and the global inflation (Eq. (12)), we use efficient GMM (generalized method of moments) to match a set of moments for the U.S. and the U.K. inflation and for the dollar-pound XVP. Specifically, we match the following set of moments:

$$m(\theta) = \begin{bmatrix} \sum_1^T(\pi_{t+1} - \mu_\pi - \rho_\pi\pi_t - \phi_{\pi_w}(\mu_{\pi_w} + \rho_{\pi_w}\pi_{w,t})) \\ \sum_1^T(\pi_{t+1} - \mu_\pi - \rho_\pi\pi_t - \phi_{\pi_w}(\mu_{\pi_w} + \rho_{\pi_w}\pi_{w,t}))\pi_t \\ \sum_1^T(\pi_{t+1} - \mu_\pi - \rho_\pi\pi_t - \phi_{\pi_w}(\mu_{\pi_w} + \rho_{\pi_w}\pi_{w,t}))\pi_{w,t} \\ \sum_1^T(\pi_{t+1} - \mu_\pi - \rho_\pi\pi_t - \phi_{\pi_w}(\mu_{\pi_w} + \rho_{\pi_w}\pi_{w,t}))\pi_t^2 \\ \sum_1^T(\pi_{t+1} - \mu_\pi - \rho_\pi\pi_t - \phi_{\pi_w}(\mu_{\pi_w} + \rho_{\pi_w}\pi_{w,t}))\pi_{w,t}^2 \\ \sum_1^T((\pi_{t+1} - \mu_\pi - \rho_\pi\pi_t - \phi_{\pi_w}(\mu_{\pi_w} + \rho_{\pi_w}\pi_{w,t}))^2 - \phi_\pi^2 E(\sigma_{\pi,t}^2) \\ \text{...(repeat for foreign economy)} \\ \sum_1^T(\pi_{w,t+1} - \mu_{\pi_w} - \rho_{\pi_w}\pi_{w,t}) \\ \sum_1^T(\pi_{w,t+1} - \mu_{\pi_w} - \rho_{\pi_w}\pi_{w,t})\pi_{w,t} \\ \sum_1^T(\pi_{w,t+1} - \mu_{\pi_w} - \rho_{\pi_w}\pi_{w,t})\pi_{w,t}^2 \\ \sum_1^T((\pi_{w,t+1} - \mu_{\pi_w} - \rho_{\pi_w}\pi_{w,t})^2 - \phi_{\pi\sigma_w}^2 E(\sigma_{w,t}^2) \\ \sum_1^T(XVP_t - b_{xvp,q}E(q_t) - b_{xvp,\sigma_w}E(\sigma_{w,t}^2)) \\ \sum_1^T(XVP_t - b_{xvp,q}E(q_t) - b_{xvp,\sigma_w}E(\sigma_{w,t}^2))^2 - b_{xvp,q}^2 Var(q_t) - b_{xvp,\sigma_w}^2 Var(\sigma_{w,t}^2) \\ \sum_1^T(XVP_{t+1} - XVP_t - b_{xvp,q}(\mu_q + (\rho_q - 1)E(q_t)) \\ -b_{xvp,\sigma_w}(\mu_{\sigma_\pi} + (\rho_{\sigma_\pi} - 1)E(\sigma_{w,t}^2)) \end{bmatrix},$$

where $\theta = \{\mu_\pi, \rho_\pi, \phi_\pi, \phi_w, \mu_\pi^*, \rho_\pi^*, \phi_\pi^*, \phi_w^*, \mu_{\pi_w}, \rho_{\pi_w}, \phi_{\pi\sigma_w}, \mu_{\sigma_\pi}, \rho_{\sigma_\pi}, \phi_{\sigma_\pi}, \mu_{\sigma_\pi}^*, \rho_{\sigma_\pi}^*, \phi_{\sigma_\pi}^*, \mu_w, \rho_w, \phi_{\sigma w}, \rho_{\pi w \sigma w}\}$ is the set of parameters to be estimated.

To reduce the dimension of the optimization problem and to implicitly focus our attention on matching the levels of inflation and XVP, we make a few simplifications. First, we estimate the parameters in two steps; a first step in which we estimate only the parameters affecting the first-order moments of inflation (GMM is robust to heteroskedasticity), and a second step in which we fix the parameters in the first step and estimate the parameters affecting second-order moments of inflation and XVP-related moments. In the second simplification, we assume that some of the non-key inflation volatility parameters are homogeneous; specifically, we assume $\mu_{\sigma_\pi} = \mu_{\sigma_\pi}^* = \mu_w$, $\rho_{\sigma_\pi} = \rho_{\sigma_\pi}^* = \rho_w$, and $\phi_{\sigma_\pi} = \phi_{\sigma_\pi}^* = \phi_{\sigma w}$. Third, we use grid search to identify $\rho_{\pi_w\sigma_w}$, the parameter driving the correlation between the level and the volatility of global inflation, as we find that, although the moments are mostly insensitive to it, this parameter is key to match the predictability patterns (see Fig. 5).[23]

We find the set $\widehat{\theta}$ that minimizes the functions $J1 = m_1(\theta_1)'W_1 m_1(\theta_1)$ and $J2 = m_2(\theta_2)'W_2 m_2(\theta_2)$, where m_1 (m_2) is a subset of $m(\theta)$ that includes only the moments related to the level of inflation (volatility of inflation and XVP), θ_1 (θ_2) is the subset of parameters in m_1 (m_2), and W_1 and W_2 are efficient GMM weighting matrices, which are obtained iteratively departing from the identity matrix (up to a maximum of 100 iterations).

Table C.1 shows the estimated parameters for the benchmark specification. To facilitate the interpretation of the parameters, Table C.2 compares a set of key model-implied moments for the U.S. and the U.K. economies with those observed for a sample between 2000 and 2011 for these two countries and for an average of all countries in

[23] The estimate of ϕ_{σ_π} is correlated with that of $\rho_{\pi_w\sigma_w}$ and tends to go to its boundaries: $\widehat{\phi}_{\sigma_\pi}$ is very low for low $\widehat{\rho}_{\pi_w\sigma_w}$ and very high for high $\widehat{\rho}_{\pi_w\sigma_w}$. As for $\rho_{\pi_w\sigma_w}$, however, the moments are largely insensitive to this parameter. The model-implied predictability patterns are qualitatively similar to the observed patterns only for large values of $\phi_{\pi\sigma_w}$. For small values of this parameter, the predictability patterns are almost flat, irrespective of the values of ω or $\rho_{\pi_w\sigma_w}$.

Table C1
GMM estimated parameters for the nominal component of the model.

This table shows the estimated parameters for each country's inflation processes (Eq. (10) and its foreign counterpart) and for the global inflation (Eq. (12)). To estimate these parameters, we use efficient GMM to match the set of moments for the U.S. and the U.K. inflation and for the dollar-pound XVP described in Appendix C.

Parameter	Estimated value
μ_π	$1.69x10^{-7}$
ρ_π	0.88
ϕ_w	0.11
μ_π^*	$9.44x10^{-15}$
ρ_π^*	0.88
ϕ_w^*	0.17
μ_{π_w}	$5.93E-05$
ρ_{π_w}	0.96
ϕ_π	0.07
$\mu_{\sigma_\pi} = \mu_{\sigma_\pi}^* = \mu_w$	$8.70E-05$
ρ_{σ_π}	0.64
$\phi_{\sigma_\pi} = \phi_{\sigma_\pi}^* = \phi_{\sigma_w}$	20.00
ϕ_π^*	0.04
$\phi_{\pi\sigma_w}$	0.02
$\rho_{\pi_w\sigma_w}$ (grid)	1.00
$J = J1 + J2$	27.86

Table C2
Targeted model-implied versus observed moments.

This table compares a set of model-implied moments for the U.S. and the U.K. economies with those observed for a sample between 2000 and 2011. The observed volatility of inflation is calculated as the absolute value of inflation. We also compare the moments for a global aggregate, which is calculated as the equally weighted average of all countries in our sample. The model-implied moments are calculated using the benchmark set of estimated parameters in Table C.1. All magnitudes are annualized, unless noted.

	Observed	Model-implied
U.S.		
Mean inflation	2.44%	1.68
Volatility inflation	1.36%	0.76
Correlation level and volatility inflation	0.14	0.00
U.K.		
Mean inflation	2.36%	2.64
Volatility inflation	1.09%	0.44
Correlation level and volatility inflation	0.51	0.00
Global		
Mean inflation	2.67%	1.89
Volatility inflation	0.82%	0.44
Correlation level and volatility inflation	0.79	1.00
Financial variables		
Mean app. rate (monthly)	−0.03%	2.69
Volatility app. rate (monthly)	2.64%	13.55
Mean GBP-dollar XVP	$13.33\%^2$	41.87
Volatility GBP-dollar XVP	$40.17\%^2$	0.27
Correlation (VP, XVP)	−0.40	0.75

our sample ("Global"). For the benchmark set of estimated parameters, our model underestimates the level of U.S. and the global inflation (1.68 compared to an observed 2.44% and 1.89 compared to an observed 2.67%, respectively) and overestimates the level of U.K. inflation (2.64 compared to an observed 2.36%). For both countries and for the global inflation, the model-implied volatility is lower than the observed values. Using a grid estimate, we find $\widehat{\rho}_{\pi_w\sigma_w} = 1$, which implies that global inflation level and inflation volatility are perfectly correlated. In contrast, at the country level, the model-implied correlation between the level and the volatility of inflation is virtually zero, although the observed values are 0.14 and 0.51 for the U.S. and the U.K., respectively.

While the deviations between model-implied and observed moments is relatively small for the nominal variables, these deviations are notably larger for the financial variables. In particular, the set of estimated parameters yields much higher values than the observed average appreciation rate, volatility of appreciation rate, and XVP, while it underestimates the volatility of XVP. Also, the correlation between XVP and VP is relatively high at 0.75, while, for our sample, the observed correlation is −0.40.

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