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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.

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

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

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 alterna-[illegible] Tj 3.9701 0 0 7.97 [illegible] (Fn[(ma)]-726.98 r)F12.5(i)-23(a)-22.9(n)-23(c)-22.4(e34)Tj 87F1 tive variance risk premium measures;

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

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 *XVPs*, 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

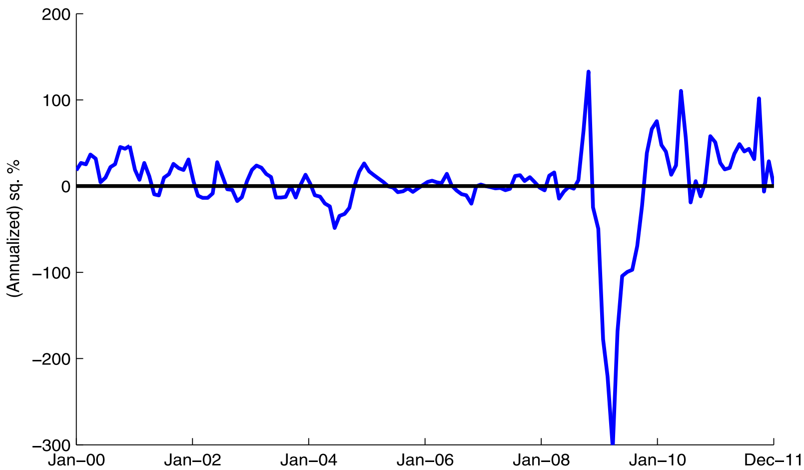

Fig. 1. World currency variance risk premium (XVP). worThe figure shows the six-month world XVP, which is calculated as the equally weighted average of the variance risk premiums of 17 currencies with respect to the U.S. dollar (see Table 2). Each currency's variance risk premium is measured as the difference 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 a

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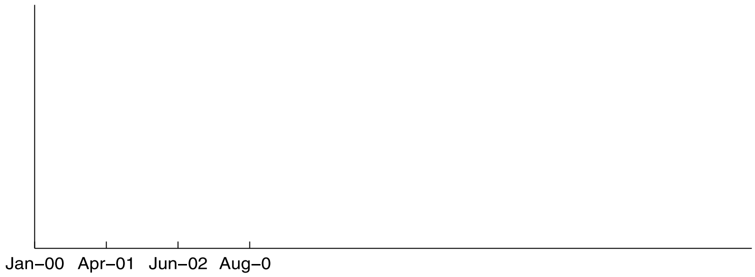
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Table 6
The predictive power of the world XVP and the U.S. VP for exchange rate returns with respect to the U.S. dollar.

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 sample period runs from January 2000 to December 2011. The currency-specific estimated constants are left unreported, to save space. ^{2} of R regression and ^{2}s with respect to a univariate regression for interest rate differential (Panel A in Table 4), $R^2 - R^2_y$, and with respect to a multivariate regression for interest rate differential and U.S. VP (Panel A in Table 5), $R^2 - R^2_{y,VP}$.

$y_{US}(h) - y_i(h)$	−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
$R^2 - R$	1.20	3.46	5.48	5.98	6.20	2.90	1.56
$R^2 - R$	6.51	5.18	9.30	10.42	7.12	2.92	1.56

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 1

3 0.3

Table 7 (*continued*)

h	1	2	3	4	6	9

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 R 2 R 2	VP_{US}	4.27***	2.70***	3.18***	3.09***			

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" account-to-GDP

t heterogeneous world
find that a portfolio formed currencies

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

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, $XVP2$ 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 ($E_t^P(\sigma_{c,t+1}^2)$ and $E_t^P(\sigma_{r,t+1}^2)$) is approximated using an AR(1) forecast of the realized variance. In Panel B, $XVP3$ 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 $XVP4$, 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 in

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, $R^2 - R_y^2$.

	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***	0.74ﬃﬃ	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***0					

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.


0
8
9
12
nths)

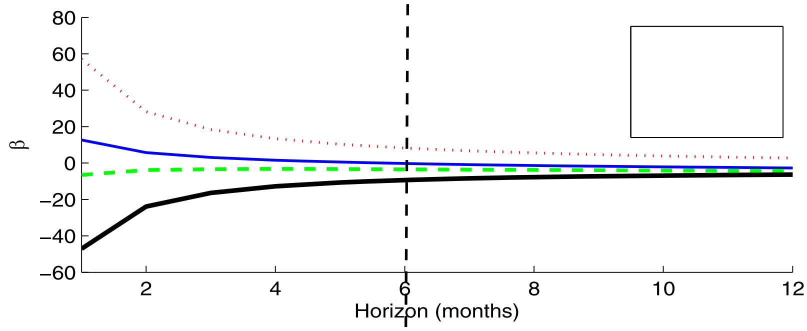
β
Horizon (months)

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

$$A_q^{\pm} = \frac{(1-\kappa_1\rho_q) \pm \sqrt{(1-\kappa_1 \ \rho q \ \Delta}}{}$$

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

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:

$$\frac{1}{h}(s_{t+h}-s_t) \simeq \frac{1}{h}\sum_{j=1}^{h}(s_{t+h}-s_t)$$

$$= \frac{1}{h}\Bigg[c_{x,h} + b_{x,\sigma_l}\left(\frac{1-\rho_l^h}{1-\rho_l}\right)\sigma_{l,t}^2 + b_{x,\sigma_l^*}\left(\frac{1-\rho_l^{*h}}{1-\rho_l^*}\right)\sigma_{l,t}^{*2}$$

$$+ b_{x,q}\left(\frac{1-\rho_q^h}{1-\rho_q}\right)q_t + b_{x,q^*}\left(\frac{1-\rho_q^{*h}}{1-\rho_q^*}\right)q_t^*$$

$$-\rho_\pi\frac{1-\rho_\pi^h}{1-\rho_\pi}\pi_t + \rho_\pi^*\frac{1-\rho_\pi^{*h}}{1-\rho_\pi^*}\pi_t^*$$

$$+ b_{x,\pi_w}\pi_{w,t} + f_c(z_{y,t+1},..z_{y,t+h})\Bigg], \tag{B.1}$$

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

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

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,

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

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

$$cov(\frac{1}{h}\sum_{j=1}^{h}(s_{t+j}-s_{t+j-1}), VP_t)$$

$$= \frac{1}{h}b_{vp,q}b_{x,q}\left(\frac{1-\rho_q^h}{1-\rho_q}\right)var(q_t)$$

and

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

The T-month ahead XVP is given by

$$XVP_t(T) \approx \frac{1}{T}\sum_{j=1}^{T}XVP_{t+j} = \Bigg[b_{xvp,q}q_t\left(\frac{1-\rho_q^T}{1-\rho_q}\right)$$

$$+ b_{xvp,\sigma_w}\sigma_{w,t}^2\left(\frac{1-\rho_w^T}{1-\rho_q}\right) + f_{xvp}(z_{t+1,..}z_{t+T})\Bigg], \tag{B.2}$$

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:

$$cov\left(\frac{1}{h}\sum_{j=1}^{h}(s_{t+j}-s_{t+j-1}), XVP_t(T)\right)$$

$$= TCov(c_{t+1}, XVP_{t+1}) + \sum_{j=1}^{T-1}(T-j)cov(c_{t+1}, XVP_{t+j+1})$$

$$+ \sum_{j=1}^{h-1}(h-j)cov(c_{t+1+j}, XVP_{t+1}),$$

where

$$Cov\left(\frac{1}{h}\sum_{j=1}^{h}(s_{t+j}-s_{t+j-1}), XVP_{t+1}\right) = b_x$$

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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