MULTIPLICATIVE NEWS IN FOREIGN EXCHANGE MARKETS

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Abstract. Slope coefficients from time series regressions of exchange rate returns on forward premiums are well below the efficient market value, but are also highly unstable. We focus on the econometric identification of this slope variability through panel estimation. No arbitrage arguments imply that the forward premium slopes are identical across exchange rates at each point in time. The slope variability reflects the heavy tail nature of the news distribution and the news dominance feature. Surprisingly, the cross sectional deviations from market efficiency are statistically insignificant for a majority of considered months. The slope estimates reveal that the deviations from unbiasedness are smaller when forward premiums are large and volatile as during currency crises.

Key words and phrases. forward premium; panel regression; foreign exchange news; time variation; triangular arbitrage; numeraire invariance; Peso effect
JEL classification: F3.

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1. Introduction

The efficient market hypothesis applied to foreign exchange implies that the expected excess forex returns should be zero. In other words, spot speculation net of the spread between spot and forward prices in the foreign exchange market should not be profitable if forex markets process information efficiently. Let s and f denote the log spot foreign exchange rate and 1-period log forward exchange rate, respectively. The efficient market hypothesis for foreign exchange can then be stated as:

$$\mathbf{E}_{t-1}[s_t - s_{t-1}] = \beta\left(f_{t-1} - s_{t-1}\right), \tag{1.1}$$

with $\beta = 1$. The left hand side and right hand side of the equation stand for the expected spot depreciation rate and the lagged forward premium, respectively. The forward premium equals the interest differential by covered interest parity. This part is the reward for currency speculation known at the time of investment. Stated otherwise, in an efficient market environment the forward premium should be an unbiased predictor of the future spot rate change.

If (1.1) holds, running a linear regression of the spot return on a constant and the lagged forward premium should give an intercept estimate close to zero and a slope estimate $\widehat{\beta}$ close to one. This so-called "forward premium regression" typically renders a slope coefficient that is significantly below one, and often negative, see e.g. Lewis (1995) and Engel (1996) for comprehensive literature surveys on this anomaly. The carry trade phenomenon led to a revived interest in the topic, see e.g. Burnside et al. (2010) and Brunnermeier et al. (2008).

Instead of developing yet another theoretical explanation for the downward bias, this paper focuses on the observed temporal instability of the forward premium slope. In contrast to the slope bias anomaly, the slope variation has not received much attention in the literature. For an early reference, see Fama (1984) who documented temporal instability in $\widehat{\beta}$ for the early post-Bretton Woods years. Using rolling regressions, Barnhart and Szakmary (1991) find evidence for time variation of the slope, but, except for the latter half of the 1980's, the hypothesis $\widehat{\beta} = 1$ can often not be rejected. Bekaert and Hodrick (1993) conclude that "formal tests of the stability of the coefficients indicate that the parameters have changed over time".

A time varying forward premium slope does not necessarily violate the efficient market hypothesis (1.1). More specifically, assume that the exchange rate return can be decomposed into its conditionally expected value and into "additive" as well as "multiplicative" news:

$$s_t - s_{t-1} = (\beta + \varepsilon_t)\, \mathbf{E}_{t-1}\, (s_t - s_{t-1}) + v_t. \tag{1.2}$$

News terms ε and v are (conditionally and unconditionally) zero mean, serially and mutually uncorrelated innovations. Moreover, the expectations formation is rational if $\beta = 1$. By using eq. (1.2) to substitute for the expected exchange rate change in (1.1), the actual change in the exchange rate can be expressed as:

$$s_t - s_{t-1} = (\beta + \varepsilon_t)(f_{t-1} - s_{t-1}) + v_t, \tag{1.3}$$

with $\beta = 1$. Clearly, (1.1) easily follows from taking expectations with respect to time $t-1$ information on both sides of (1.3). Traditionally the additive innovation v is interpreted as forex news, see Frenkel (1981). In a similar vein, ε constitutes the multiplicative news channel. It gives the direction and magnitude by which the interest differential is propagated through the forex spot market. The latter factor expresses that there is no economic necessity at any point in time for the ex post realized spot returns across different countries to be aligned along lines with a slope of 45 degrees with respect to the forward premium. This should only hold in expectation.

The cross sectional identification of the slope variation in (1.3) constitutes the main contribution of this paper. Notice that a time varying slope that also varies across currencies is not identified. However, no arbitrage conditions in the foreign exchange market enable us to identify the cross sectional slope in a numeraire invariant way, i.e., $\varepsilon_t^{ij} = \varepsilon_t$ for all currency pairs ij. Anticipating on our results, we find very large swings in the slope parameter over certain time periods. Model (1.3) can also be augmented with a true cross sectional model of risk, allowing better identification of time variation in risk factors. As such, our panel data approach also stands into the risk premium tradition.[1] As concerns the economic rationale behind the slope variation, we show that ε is small when forward premia are large and volatile. This may be triggered by transaction costs and no arbitrage bands in foreign exchange markets. Moreover, large and volatile forward premia often coincide with currency crises which suggests that at least part of the slope variation may be related to "Peso" bubbles and bubble bursts.

The remainder of the paper is structured as follows. In the next section we present a panel data specification of the stochastic coefficient model (1.3). A Quasi Maximum Likelihood (QML) panel estimation

[1] Lustig et al. (2011) constitutes another recent example of modelling the cross sectional variation in currency returns.

procedure is proposed in section 3. The empirical results are contained in sections 4 and 5. We summarize and conclude in section 6.A panel with multiplicative effects

Rolling regressions and other nonparametric methods implicitly assume that the parameters vary slowly over time. Here we make a different identifying assumption. By assuming that the time varying slope coefficient in (1.3) is uncorrelated with the forward premium, the slopes have to be the same for each bilateral exchange rate because of the triangular arbitrage relations between exchange rates. A cross sectional regression of, say, multiple dollar spot returns against the respective forward premia will then provide an estimate of the slope for each time period.

Let s_t^{i0} be the log-price of the foreign currency i in units of the numeraire currency 0 at time t and f_t^{i0} be the log-forward price at t of the exchange rate with a 1-month maturity ($i = 1, \cdots, N$; $t = 1, \cdots, T$). For brevity we introduce the shorthand notation $y_t^{i0} = s_t^{i0} - s_{t-1}^{i0}$ and $x_t^{i0} = f_{t-1}^{i0} - s_{t-1}^{i0}$ for the spot return and the lagged forward premium, respectively. We propose the model

$$(1.4) \qquad y_t^{i0} = \gamma_t^{i0} + \beta_t x_t^{i0} + v_t^{i0}, \quad i = 1, \ldots, N; \quad t = 1, \cdots T.$$

where both the intercept γ_t^{i0} and the slope β_t are time varying. Moreover, the intercept will depend on the specific currency pair. We assume that the intercept can be factorized in a fixed time effect γ_t, common to all currencies, and fixed currency effects δ_i that are constant over time:

$$(1.5) \qquad \gamma_t^{i0} = (\delta_i - \delta_0)\gamma_t, \qquad i = 1, \ldots, N; \quad t = 1, \cdots T,$$

and where δ_0 stands for the fixed currency effect of the numeraire currency. Notice that the intercept satisfies the triangular arbitrage restrictions between exchange rates. Moreover, the slope β_t is a numeraire invariant fixed time effect.

The slope's numeraire invariance easily follows from the triangular arbitrage restrictions in the foreign exchange market. Consider (1.4) for currencies i and k. Subtracting y_t^{k0} from y_t^{i0} renders an equation for the cross rate return y_t^{ik}. The latter spot return only depends on x_t^{ik} provided that the time varying slope is numeraire invariant. As for the factor structure in (1.5), it ensures that the cross rate regression for y_t^{ik} has a time varying intercept $(\delta_i - \delta_k)\gamma_t$. Our factor assumption is that the same γ_t appears in all exchange rates.

The above multiplicative specification for the intercept term can be interpreted as a risk factor in an international CAPM (ICAPM) model. Let R^i and R^0 denote the nominal interest rates on internationally traded comparable one-period bonds denominated in currency i and the numeraire currency 0, respectively. The return on the foreign bond expressed in the numeraire currency 0 equals $R^i_{t-1} - \left(s^{i0}_t - s^{i0}_{t-1}\right)$. Finally let R^M_t be the excess return on a world market index in a one factor version of the International CAPM, see e.g. Lewis (1995). The ICAPM then implies the two relations

$$R^i_{t-1} - \mathbf{E}_{t-1}\left(s^{i0}_t - s^{i0}_{t-1}\right) = \delta_i \mathbf{E}_{t-1}[R^M_t], \tag{1.6}$$

$$R^0_{t-1} = \delta_0 \mathbf{E}_{t-1}[R^M_t], \tag{1.7}$$

where δ_ℓ $(\ell = i, 0)$ is the sensitivity of both required returns denominated in the numeraire currency with respect to the return on the world market portfolio. Upon subtracting (1.7) from (1.6) we obtain:

$$\mathbf{E}_{t-1}[s^{i0}_t - s^{i0}_{t-1}] = -(\delta_i - \delta_0)\mathbf{E}_{t-1}[R^M_t] + \beta(R^i_{t-1} - R^0_{t-1}), \quad \beta = 1 \tag{1.8}$$

One easily recognizes the efficient market hypothesis (1.1) augmented with a term that can be interpreted as a foreign exchange risk premium. The risk premium is the product of a fixed time effect, common to all currencies, which is the price of risk $(-\mathbf{E}_{t-1}[R^M_t])$, and the relative amount of risk or currency specific effect measured by the difference in the 'beta's'.

The expected spot return in (1.8) can be expressed in terms of the realized spot return and multiplicative and additive forex news by means of the expectations decomposition (1.2). Rearranging terms, one obtains:

$$s^{i0}_t - s^{i0}_{t-1} = (\delta_i - \delta_0)\gamma_t + \beta_t(f^{i0}_{t-1} - s^{i0}_{t-1}) + v^{i0}_t, \tag{1.9}$$

with $\gamma_t = -\beta_t \mathrm{E}_{t-1}[R^M_t]$, $\beta_t = \beta + \varepsilon_t$ and where the forward premium equals the interest differential due to the covered interest parity. Model (1.9) boils down to the stochastic coefficient model (1.3) augmented with a time varying intercept that is interpretable as a forex risk premium. Specification (1.9) is equivalent to the multiplicative panel model (1.4)-(1.5).

In general a fixed time effect γ_t only appears for those currency pairs that have different factor loadings than the numeraire currency ($\delta_i \neq \delta_0$). We refer to the system (1.4)-(1.5) as a panel with multiplicative fixed time effects γ_t (common to all currencies) and multiplicative fixed currency effects δ_i (constant over time). Notice that the δ_i's in (1.5) are only identified in a panel data framework under the identifying restriction that all δ_i are constant over time. Since the δ_i only appear as differentials, normalization rules are necessary to fix their absolute value and their scale.[2]

As a special case, set $\beta_t = \beta$ and take a first order Taylor expansion of the multiplicative term (1.5) about its mean in order to get

$$y_t^{i0} = \phi + \eta_t + \mu_i + \beta x_t^{i0} + v_t^{i0}, \tag{1.10}$$

where $\phi = \overline{\delta}\overline{\gamma}$ is a constant term, $\eta_t = \overline{\delta}\gamma_t$ is an additive fixed time effect common across currencies, $\mu_i = \overline{\gamma}\delta_i$ is an additive fixed currency effect constant over time, and a bar over a variable denotes mean values across time or across currencies. This panel specification has been employed by e.g. Mayfield and Murphy (1992). Flood and Rose (1996) also consider this type of panel regressions without the fixed time effect η_t. The problem with this model is that the country and time effect appear in an additive way that is not numeraire invariant, i.e. the fixed time effect η_t drops out of the panel regression for cross rate y_t^{ij} because one has to subtract the corresponding panel equations for y_t^{i0} and y_t^{j0} in order to obtain the cross rate panel. For this reason we favor the numeraire invariant specification (1.5).

2. Estimation

We will estimate γ_t and β_t in eqs. (1.4)-(1.5) period by period using cross sectional regressions. The methodology is similar to the Fama-MacBeth (1973) procedure. This is standard fare in models where stock returns are regressed on the characteristics of a stock like its size and various accounting ratios. In our setup the currency characteristics are the forward premium x_t^{i0} and the sensitivity of individual currencies to the world market.

We also need to impose some assumptions on the additive disturbance terms v^{i0}. First, we make the standard assumptions that v_t^{i0} in (1.4) has mean zero and is orthogonal to all information variables dated t or earlier. Also, we know from the univariate time series regressions that the additive disturbances exhibit conditional heteroskedasticity. We therefore allow for time varying variances without making any assumptions on the time series properties of the error variance. This can be done in a panel data framework at the price of making additional cross sectional assumptions. We assume that the error term can be decomposed as in Koedijk and Schotman (1990),

$$v_t^{i0} = v_t^i - v_t^0. \tag{2.1}$$

[2]For that purpose we set $\delta_0 = 0$ and $\delta_N = 1$ in the empirical application.

This decomposition expresses the additive news in exchange rate s^{i0} as the difference between news about currencies i and 0 separately. It is consistent with the triangular identity, and assures that the whole system is completely numeraire invariant. We assume that the components in (2.1) are mutually uncorrelated and have a common variance $\frac{1}{2}\sigma_t^2$. The vector of N error terms against the common numeraire currency 0 will than have a covariance matrix $\sigma_t^2\mathbf{S}$ with

$$\mathbf{S} = \frac{1}{2}\left(I + \iota\iota'\right). \tag{2.2}$$

The decomposition (2.2) explicitly recognizes that the error terms contain the 'numeraire news' v_t^0 as a common factor, introducing strong cross sectional correlations. In other words, it states that all exchange rates against the dollar will be correlated, because they share the dollar component. Implicitly the decomposition introduces a random time effect, since the numeraire error v_t^0 is present in all equations of the model. The assumption that all v_t have the same variance $\frac{1}{2}\sigma_t^2$ accommodates the time series heteroskedasticity.

Estimation of model (1.4)-(1.5) with covariance structure (2.2) proceeds through Quasi Maximum Likelihood (QML). Assume that currency 0 is the numeraire, and write the model in vector notation as

$$y_t = X_t\theta_t + v_t \qquad t = 1, \ldots, T \tag{2.3}$$

where y_t is an N-vector containing the exchange rate changes y_t^{i0} against the numeraire currency 0. The numeraire invariant parameters are gathered in the parameter vector $\theta_t = (\gamma_t\ \beta_t)'$. The matrix of explanatory variables $X_t = (\delta\ x_t)$ is the $(N \times 2)$ matrix in which δ contains elements $(\delta_i - \delta_0)$ and x_t are forward premia x_t^{i0}, $(i = 1, \cdots, N)$. The log-likelihood function now takes the form

$$\ln L = -\frac{1}{2}\sum_{t=1}^{T}\left(N\ln\sigma_t^2 + \frac{v_t'^{-1}\mathbf{S}^{-1}v_t}{\sigma_t^2}\right), \tag{2.4}$$

From the first order conditions with respect to θ_t and σ_t^2 we find the conditional estimators

$$\hat{\theta}_t = \left(X_t'\mathbf{S}^{-1}X_t\right)^{-1}X_t'\mathbf{S}^{-1}y_t \tag{2.5}$$

$$\hat{\sigma}_t^2 = \frac{1}{N}y_t'\left(\mathbf{S}^{-1} - \mathbf{S}^{-1}X_t(X_t'\mathbf{S}^{-1}X_t)^{-1}X_t'\mathbf{S}^{-1}\right)y_t \tag{2.6}$$

Substituting (2.6) in (2.4) renders the concentrated likelihood function

$$\ln L^* = -\frac{N}{2}\sum_{t=1}^{T}\ln\hat{\sigma}_t^2. \tag{2.7}$$

This is still a function of δ, and must be maximized numerically. Robust standard errors for δ are obtained from the Hessian and Information matrix of the concentrated likelihood function. The information matrix can be computed from the outer product of the scores of the concentrated likelihood function (2.7).

The covariance matrix for the time effects β_t and γ_t can be computed conditional on the estimated value of δ using the standard least squares formula

$$\mathbf{V}\left(\hat{\theta}_t\right) = \hat{\sigma}_t^2 \left(X_t' \mathbf{S}^{-1} X_t\right)^{-1}. \tag{2.8}$$

3. Empirical results

This section reports the results of the QML estimator for the panel specification (1.4)-(1.5). We further distinguish between individual currency effects δ_i (4.1.) and estimates of the fixed effects θ_t (4.2). The model is fitted for two distinct currency panels. The "fixed" panel spans European currencies that previously belonged to the European Monetary System (EMS) but that were abolished with the introduction of the European single currency. As for the "floating" panel it solely contains non-Eurozone currencies and runs until February 2008. The fixed or EMS panel (January 1976 until December 1998) spans Austria, Belgium, United Kingdom, France, Germany, Ireland, Italy, Portugal, Spain and the Netherlands. The floating panel (January 1976 until February 2008) contains Canada, United Kingdom, Denmark, Japan, Norway, Sweden and Switzerland. All currencies are expressed against the US$ (we also use Dmark as benchmark currency in the fixed panel).[3] Further details on the data are provided in Appendix A.

3.1. **Individual currency effects.**

Table 1 reports estimates and accompanying standard errors for the country specific intercept terms $\delta_{i0} = (\delta_i - \delta_0)$.

[Insert Table 1]

As the country effect δ_{i0} is defined as a differential relative to a numeraire currency it is not numeraire invariant. We explained earlier that the fixed time effects (γ_t, β_t) are numeraire invariant by the triangular arbitrage condition.

[3] The inclusion of the British Pound in both panels is somewhat of a compromise: although Britain is not part of the eurozone it made part of the EMS during several important episodes of its history. Moreover, given the 1992 EMS crisis (strong involvement of £sterling), we want to keep it in the fixed panel in order to study the relation between currency crises and slope variation.

Given the numeraire dependence of the country effects, we distinguish between fixed currency panel outcomes and floating currency panel outcomes in Table 1. Thus it makes sense to calculate the country effects against different 'economically meaningful' numeraire currencies. The individual currency effects for the floating panel are expressed against the US$ numeraire. As for the fixed panel, we distinguish between fixed country effects both against the US$ and Deutsche Mark numeraires. The latter numeraire choice enables one to check whether there is an influence of German monetary policy on the other EMS currencies. The US is also taken as numeraire for both panels because it assumed to be the numeraire currency for most of the literature on the subject.

The country specific effects relative to the US$ are all significant in both panels except for the Canadian Dollar. The Canadian case can be understood from the high degree of monetary integration between the two countries. If monetary integration is indeed reflected by the insignificance of $\widehat{\delta}_{i0}$, it is also of interest to consider the significance of the country specific effect within the former EMS. When Germany is used as the numeraire country the price of country specific risk decreases for all EMS cross rate returns and even becomes statistically insignificant for Belgium, the Netherlands and Austria. These outcomes confirm earlier work on the so-called 'German Dominance' hypothesis, i.e., national Central Banks tried to mimic the Bundesbank policies in order to keep their exchange rates against the Deutsche Mark within the bands.[4]

3.2. **Time variation.** The estimated time series for $\theta_t = (\gamma_t\ \beta_t)'$ are reported in Figure 1. The Figure further distinguishes between fixed and floating panel outcomes. The top and bottom row of pictures in the Figure correspond with β_t and γ_t, respectively.

[Figure 1]

The figure shows huge parameter variation but it's severity differs over time. More specifically, the variation in the forward premium slope seems to increase towards the end of the sample period for both currency panels. Also, the time variation seems lower for the fixed panel. The smaller time variation may be due to the stabilizing effect of EMS target zone bands on exchange rate volatility.[5]

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[4]Earlier work on the German Dominance Hypothesis include e.g. von Hagen and Frattianni (1990) and Kirchgassner and Wolters (1992).

[5]Upon comparing cross sectional standard deviations of spot returns for the two currency panels, we indeed find that freely floating standard deviations exceed

The observed parameter variability is remarkable in view of the presumed smoothness that is inherent to the traditional univariate time series approach. One might question whether the observed time variation cum time series properties is not a mere statistical "artefact". However, it can be easily shown that although some of the volatility in $\widehat{\theta}_t$ can be attributed to estimation error, the cross sectional covariance matrix $V(\widehat{\theta}_t)$ is small enough to conclude that θ is far from constant. In order to compute the covariance matrix of θ_t , as distinct from the covariance matrix of the estimated $\widehat{\theta}_t$, we use the decomposition

$$\widehat{\theta}_t = \theta_t + (\widehat{\theta}_t - \theta_t). \tag{3.1}$$

Since $\widehat{\theta}_t$ is the best linear estimator of θ_t, the estimation error $(\widehat{\theta}_t - \theta_t)$ is uncorrelated with θ_t. Let us denote the time series covariance matrix of $\widehat{\theta}_t$ by $var(\widehat{\theta}_t)$ and the cross sectional covariance matrix of θ_t by $\mathbf{V}(\theta_t)$. It then immediately follows from (3.1) that

$$var(\widehat{\theta}_t) = \mathbf{V}(\theta_t) + \mathbf{E}_t[\mathbf{V}(\widehat{\theta}_t)], \tag{3.2}$$

from which we can estimate the covariance matrix of θ_t as

$$\mathbf{V}(\theta_t) = \frac{1}{T}\sum_{t=1}^{T}(\widehat{\theta}_t - \overline{\theta}) \cdot (\widehat{\theta}_t - \overline{\theta})' - \frac{1}{T}\sum_{t=1}^{T}\mathbf{V}\left(\widehat{\theta}_t\right). \tag{3.3}$$

The time series average of the cross sectional covariance matrix $\mathbf{V}\left(\widehat{\theta}_t\right)$ filters out the noise from the time series covariance matrix of $\widehat{\theta}_t$. For our dataset this results in

$$\mathbf{V}(\theta_t; \text{float ing}) = \begin{bmatrix} 9.40 & 0.57 \\ 0.57 & 19.15 \end{bmatrix} - \begin{bmatrix} 1.80 & 0.32 \\ 0.32 & 5.98 \end{bmatrix} = \begin{bmatrix} 7.6 & 0.25 \\ 0.25 & 13.17 \end{bmatrix},$$

and

$$\mathbf{V}(\theta_t; \text{fixed}) = \begin{bmatrix} 10.39 & -2.59 \\ -2.59 & 6.75 \end{bmatrix} - \begin{bmatrix} 1.50 & -0.08 \\ -0.08 & 1.40 \end{bmatrix} = \begin{bmatrix} 8.89 & -2.51 \\ -2.51 & 5.35 \end{bmatrix},$$

for the floating and the fixed panel, respectively. From the above covariance matrices we see that the "genuine" slope variability dominates its estimation error variability (e.g. 13.17 > 5.98 for the floating case and 5.35 > 1.40 for the fixed case, respectively). Thus we can safely conclude that the observed time variation is real and not merely reflecting estimation error variability.[6]

their EMS counterparts most of the time (currency crisis periods constitute an exception).

[6]We also sorted the fixed time effects and the slope estimates on ascending values of the (absolute value) of the cross sectional t-statistics using the variances on the

Next, Figure 2 contains the empirical distributions (histograms) together with some descriptive statistics for both panels.

[Insert Figure 2]

The time series average of the forward premium slope equals 0.535 and 0.010 for the fixed panel and the floating panel, respectively. One rejects that the average slope is equal to 1 using the normal based Z-test.[7] Thus, although the average slope estimate comes closer to its efficient market value of 1 in comparison to univariate studies, the cross sectional approach does not resolve the downward bias puzzle. Moreover, this rejection is not caused by a few outliers. Both average slope estimates as well as the accompanying test statistic values decline (stronger rejection) when truncating the most extreme observations from the tails of $\widehat{\beta}_t$. We also applied some standard nonparametric median tests (binomial sign test, Wilcoxon signed-ranks test) to evaluate whether the slope's median equals one. This makes sense given the observed skewness in the empirical slope distributions of Figure 2. However, the median is also found to deviate from one.[8]

The availability of the cross sectional standard errors in $V\left(\widehat{\theta}_t\right)$ also enables one to test the unbiasedness hypothesis $H_0 : \beta_t = 1$ for each month in the sample period. Surprisingly, this cross sectional unbiasedness hypothesis is rejected in only 40% of the months in the fixed panel and for 28% of the months in the floating panel. At the 1% significance level, these rejection rates further reduce to 23% and 10% for the fixed panel and the floating panel, respectively. Otherwise stated, although slope variation is real, the slope is insignificantly different from 1 in a majority of time periods.

Serial correlation tests (Ljung-Box test statistics) do not reveal any significant linear dependence in the estimated time series. However, there is some evidence for nonlinear dependence in $\widehat{\gamma}_t$ (from the ARCH LM test), which is in accordance with a risk premium interpretation of the time varying intercept.[9]

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diagonal of $V\left(\widehat{\theta}_t\right)$. The most sizeable intercept and slope estimates also exhibit the highest statistical significance.

[7] The "normal" standard deviations are $2.598/\sqrt{275} = 0.156$ for the fixed sample and $4.376/\sqrt{385} = 0.223$ for the floating panel.

[8] In contrast with the slope results, the null hypothesis of a zero mean and median for the fixed time effect γ_t cannot be rejected.

[9] The test results for serial correlation and heteroskedasticity are available upon request, but were omitted for sake of space considerations.

4. Economic origins of the slope variation

Previous research on the origins of the forward premium puzzle suggests that the severity of the forward premium bias is inversely related to the size/volatility of the forward premium, see e.g. Flood and Rose (1996) or Huisman et al. (1998). It is interesting to investigate whether this "size" hypothesis also holds for the time varying deviations from unbiasedness identified in this paper, i.e. is the slope less vehemently fluctuating when the forward premium is large and volatile?

A simple way to proceed is to sort the slope estimate $\widehat{\beta}_t$ for ascending values of the cross-sectional average absolute forward premium and the cross sectional variance of the forward premium, respectively.[10] The big advantage of this approach as compared to the earlier cited approaches is that it does not require dummying out data points.[11] The sorted vectors are shown in Figure 3 for the EMS and the floating currency panel, respectively. The figures show that the slope variation and the (absolute) size and volatility of the forward premium appear to be inversely related; the inverse relation is most clearly present for the EMS slopes.

[Insert Figure 3]

Two economic explanations come to mind for the observed inverse relationship. First, transaction costs in the foreign exchange market may induce 'no-arbitrage' or 'inactivity' bands, see e.g. Baldwin (1990) or, more recently, Sercu and Vandebroek (2005). In the absence of forward speculation, spot and forward rates (or, alternatively, spot returns and lagged forward premiums) might wander away from each other. As a result, financial frictions can cause such arbitrage inactivity bands making β indeterminate. Figure 4 shows the time variation in the (annual percentage) cross sectional forward premium means and standard deviations for the fixed panel and the floating panel, respectively.

[insert Figure 4]

Average (absolute) cross sectional interest differentials and accompanying volatilities have clearly diminished over the 2nd half of the sample. For the fixed panel, this reflects the convergence in cross border interest spreads anticipating the introduction of the single currency.

[10] The numeraire invariant estimation approach in section 3 enables us to calculate numeraire invariant versions of the cross sectional means and variances for the forward premium.

[11] Calculating the forward premium slope on a subset of the total forex sample may induce a truncation bias, even if the true underlying slope is constant, see e.g. Boyer et al. (1999) on truncation biases when the normal distribution applies. The cross sectional slope estimates do not suffer from this potential inconvenience.

More surprisingly, the same phenomenon seems to have occurred in the floating sample indicating that money market integration is not just a European, but a world wide trend due to benign world economic conditions in the low inflation era. Given that Baldwin's (1990) model calibrations suggests that it may not be worthwhile to arbitrage for interest differentials up to 4% on an annual basis, it is not unlikely that a relative lack of arbitrage may have occurred during several subperiods of our sample, and more in particular towards the end. So that the model (1.4)-(1.5) does not necessarily perform well in times of quiescence.

The Peso effect constitutes a second possible explanation for the slope variation, see e.g. Flood and Rose (1996) or Sercu and Vinaimont (2006). Currency speculators or carrytraders who foresee a future currency realignment require large interest differentials which may temporarily induce a downward 'peso bias' in the cross sectional slopes ($\widehat{\beta}_t < 1$) as long as the regime shift does not materialize in the left hand side variable $s_{t+1} - s_t$. A regime change (or peso 'bubble burst') aligns the actual spot returns with the lagged interest differential (i.e., it forces the slope back to one). More recently, it has been argued that carry trade pay-offs may also be related to the existence of a peso problem. More specifically, the pay-off may be interpreted as compensation for the currency realignment risk or 'Peso event' risk, see e.g. Burnside et al. (2010).

In order to give some empirical content to this "Peso hypothesis", consider the slope variation around ERM realigments. We limit ourselves to investigating the Peso hypothesis for the EMS panel given that currency realignments in this framework are well documented. Moreover, potential Peso effects can be expected to be stronger for regulated currencies because target zone bands can potentially be tested by speculators if they perceive the bands as nonsustainable or noncredible. Table 2 reports slopes before, during and after ERM realignments together with some test statistics.

[Table 2]

All 18 documented realignment episodes from the EMS history are considered, see e.g. Baldwin and Wyplosz (2006, ch. 11) for their political and monetary backgrounds. The number of currencies (n) that were simultaneously realigned reflects the systemic importance of the realignment shock. The realignment slope $\widehat{\beta}_r$ is defined as the end-of-the-month slope for the month containing the realignment; whereas the pre-realignment slope $\widehat{\beta}_{r-1}$ and the post-realignment slope $\widehat{\beta}_{r+1}$ correspond with the month preceding (following) the realignment month.

We also report the cross sectional t-test for forward premium unbiasedness for each of these slopes estimates.

Table 2 shows that the realignment slopes are nearly always positive and exceeding the pre-realignment slope values ($\widehat{\beta}_{r-1} < \widehat{\beta}_r$), which is suggestive of a peso effect. In some cases, however, the realignment slopes, are far larger than one, which may reflect overshooting behavior in exchange rates during periods of financial market turmoil. The often observed inverted U-shape around realignments ($\widehat{\beta}_{r-1} < \widehat{\beta}_{r+1} < \widehat{\beta}_r$) is in line with this observation. The nonrandomness of the slope variation around realignments can be easily illustrated by counting the number of times $\widehat{\beta}_r$ either exceeds $\widehat{\beta}_{r-1}$ or $\widehat{\beta}_{r+1}$. In case of random time variation, the realignment slope is equally likely to lie above or below its preceding (or consecutive) value. However, we find that the realignment slope exceeds its pre-realignment and post-realignment value in a majority of the cases ($\widehat{\beta}_{r-1} < \widehat{\beta}_r$ in 17 cases and $\widehat{\beta}_{r+1} < \widehat{\beta}_r$ in 15 cases). A simple nonparametric runs test easily rejects the null hypothesis that the probabilities $\widehat{P}(\widehat{\beta}_r > \widehat{\beta}_{r-1}) = 17/18$ and $\widehat{P}(\widehat{\beta}_r > \widehat{\beta}_{r+1}) = 15/18$ do not differ from $1/2$.[12]Thus we find that during realignments a vast majority of forward premium slopes exceeds their pre-crisis and post-crisis values and come closer to 1 which is consistent with the peso hypothesis in foreign exchange markets.

The table also reports a cross sectional unbiasedness t-test for each reported slope using the cross sectional standard deviation generated by the panel estimator. Although the slope estimates lie closer to 1 during realignment months, the unbiasedness hypothesis ($\beta_r = 1$) is still rejected in 7 cases out of the 18 (as compared to 6 cases in the pre-realignment month). Overall, we can conclude that the nonparametric tests and cross sectional t-test provides mixed evidence that part of the slope variation can be rationalized by Peso bubbles and consequent bubble bursts.

5. Conclusions

Time series regressions of monthly forex spot returns on the lagged 1-month interest differential (so-called forward premium regressions) typically yield slope estimates well below the efficient market value of

[12]The probability that $\widehat{\beta}_r$ exceeds $\widehat{\beta}_{r-1}$ can be estimated by $\widehat{p} = R^{-1}\sum_{r=1}^{R} 1\left(\widehat{\beta}_r > \widehat{\beta}_{r-1}\right)$ with standard deviation s.e.$(\widehat{p}) = \sqrt{\widehat{p}(1-\widehat{p})/R}$. The exceedance probability that $\widehat{\beta}_r$ exceeds $\widehat{\beta}_{r+1}$ is defined in an analogous way. The statistic $(\widehat{p}-1)/s.e.(\widehat{p})$ is approximately normal for $R = 18$. Testing the null hypothesis $p = 1/2$ leads to strong rejections in both cases.

one. The slope estimates are also found to be highly unstable across subsamples. As to date, however, this latter stylized fact remained a largely underexposed anomaly. This paper focussed on the econometric identification of the slope variability and its potential causes, rather than the downward bias which already has been extensively documented. To that aim, we adopted a stochastic coefficient specification that encompasses the rational expectations hypothesis whilst allowing for time varying parameters in the forward premium regression.

A simple no arbitrage argument implies that the forward premium slope parameter (dubbed "multiplicative" news) must be identical for all exchange rates at each point in time and against each possible benchmark currency. Next, we estimated this slope parameter at each point in time by means of a cross sectional estimator. The estimator also controlled for the possibility of fixed time and country-specific effects in the currency panel. We performed the estimation for a panel of European Monetary System (EMS) currencies that seized to exist in 1999 as well as a panel of floating currencies still in existence. We found that time varying slopes vary considerably and that the slope variation is increasing towards the end of the sample for both panels. Interestingly, the monthly deviations from unbiasedness identified in our panel regression framework are still compatible with market efficiency for a majority of the considered months. This contradicts with most preceding studies that typically kept forward premium slopes constant over time for sake of econometric identification.

The high slope variability is related to the heavy tail nature of the news distribution and the news dominance feature (i.e. forward premiums are relatively small as compared to spot returns). We argued that part of the slope variation may be due to transaction costs and resulting nontrading or inactivity bands (especially in periods characterized by small interest differentials). We also observed that forward premium slopes during currency realignment episodes tend to rise and exceed their pre-realignment counterparts, i.e. existing market inefficiencies tend to diminish. This implies that 'Peso' bubbles and bubble bursts like currency crises resulting in forex regime collapses may also be responsible for part of the slope variation during these periods of financial turmoil.

Appendix A. Data description

The paper uses end-of-the-month spot and forward exchange rates from January 1976 onwards. Both data frequency and forward contract maturity are monthly in order to circumvent overlapping data problems. All series are London mid prices (Datastream) quoted at 12:00 GMT against the pound sterling for 16 countries: Austria, Belgium, Canada, Denmark, France, Germany, Ireland, Italy, Japan, the Netherlands, Norway, Spain, Sweden, Switzerland and the United States. We download exchange rates against pound sterling because this provides us with the largest possible cross section of currencies. Cross rates used in the calculations (either US$ or DMARK numeraire) are determined by assuming triangular arbitrage. We solely consider industrial currencies because of their liquidity and data availability. The Japanese forward rate is missing in Datastream prior to July 1978 and we therefore extend this time series with Harris bank data.[13]

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Table 1. Country-specific currency effects

	EMS				floating	
	$\delta_i - \delta_0$ (s.e.)					
Currencies	vs. US$		vs. DM		vs. US$	
	$\delta_i - \delta_\$$	s.e.	$\delta_i - \delta_{DM}$	s.e.	$\delta_i - \delta_\$$	s.e.
United States	0	-	-1.007	0.014	0	-
Germany	1.007	0.014	0	-	-	-
France	0.978	0.016	-0.029	0.016	-	-
United Kingdom	0.814	0.021	-0.192	0.023	0.707	0.044
Spain	0.871	0.028	-0.135	0.030	-	-
Danmark	-	-	-	-	0.919	0.035
Norway	-	-	-	-	0.778	0.040
Sweden	-	-	-	-	0.749	0.041
Italy	0.932	0.025	-0.075	0.025	-	-
Japan	-	-	-	-	0.635	0.036
Canada	-	-	-	-	-0.056	0.028
Belgium	1.003	0.015	-0.004	0.015	-	-
Netherlands	1.001	0.014	-0.006	0.014	-	-
Portugal	0.915	0.037	-0.091	0.038	-	-
Switzerland	-	-	-	-	1	-
Ireland	0.972	0.017	-0.034	0.019	-	-
Austria	1.000	-	-0.007	0.014	-	-

Note: The table reports country specific effects for two panels consisting of EMS currencies and floating currencies, respectively. EMS panel results are reported against USdollar numeraire and Dmark numeraire. Floating panel results are solely reported against USdollar numeraire. Standard errors are reported within brackets. The currency specific effect for the numeraire currency is set to zero for identification purposes. Furthermore, identifiability of the fixed time effect from the fixed currency effect requires restricting the currency specific effect of yet another country equal to 1. We have chosen Austria (EMS panel) and Switzerland (floating panel) for this purpose but other choices are valid as well.

TABLE 2. Cross sectional forward premium slopes around ERM realignments

Realignments (dd/mm/yy)	n	$\widehat{\beta}_{r+1}$	$t\left(\beta_{r+1}=1\right)$	$\widehat{\beta}_r$	$t\left(\beta_r=1\right)$	$\widehat{\beta}_{r-1}$	$t\left(\beta_{r-1}=1\right)$
24/09/79	2	-0.338	-1.254	2.886	1.967	0.207	-4.235***
30/11/79	1	0.185	-1.633	2.047	0.799	-0.338	-1.254
22/03/81	1	-1.066	-8.520***	5.371	1.971	3.158	0.976
05/10/81	2	0.932	-0.098	1.646	1.218	-0.460	-1.325
22/02/82	2	1.102	0.176	1.490	0.282	0.169	-2.230
14/06/82	4	0.644	-1.951	4.487	4.561***	0.207	-5.877***
21/03/83	7	-0.041	-2.970**	0.621	-1.338	0.101	-7.520***
18/05/83	7	2.429	5.946***	0.003	-5.211***	-0.041	-2.970*
22/07/85	7	-0.128	-4.831***	0.958	-0.059	0.937	-0.169
07/04/86	5	0.598	-1.103	0.319	-1.055	-0.042	-1.371
04/08/86	1	1.425	0.525	2.239	0.650	1.407	0.284
12/01/87	3	-0.398	-3.192**	2.576	2.278	0.116	-1.406
08/01/90	1	-0.814	-4.576***	-1.410	-2.307*	1.421	0.588
14/09/92	3	-0.070	-0.816	15.76	2.988**	1.938	2.005
23/11/92	2	0.873	-0.142	0.958	-0.084	-0.070	-0.816
01/02/93	1	0.967	-0.026	2.042	2.481*	-0.061	-2.346*
14/05/93	2	-1.869	-3.109**	5.792	3.202**	1.169	0.099
06/03/95	2	-2.325	-1.635	12.708	3.295**	6.476	1.005

Realignment (time t) slopes and 'pre'-realignment slopes (time $t-1$) are reported together with their t-test statistics for $H_0 : \beta = 1$. Realignment slopes are end-of-the-month slope estimates for the month containing the realignment whereas the pre-realignment (post-realignment) slopes are end-of-the month slope estimates for the last month prior (next) to the realignment. Rejections at the 5, 2.5 and 1 percent significance level are denoted by *, * and ***, respectively (testong unbiasedness against a two-sided alternative hypothesis. The small cross sectional dimension implies that the cross sectional t-statistics of the slope estimates follow a Student-t distribution with 10-2=8 degrees of freedom. The two-sided 5, 2.5 and 1 percent critical values for the student-(8) distribution are equal to 2.306, 2.751 and 3.355, respectively.

FIGURE 1. QML estimates of fixed time effect ($\widehat{\gamma}_t$) and forward premium slope ($\widehat{\beta}_t$) for EMS and floating currency panel

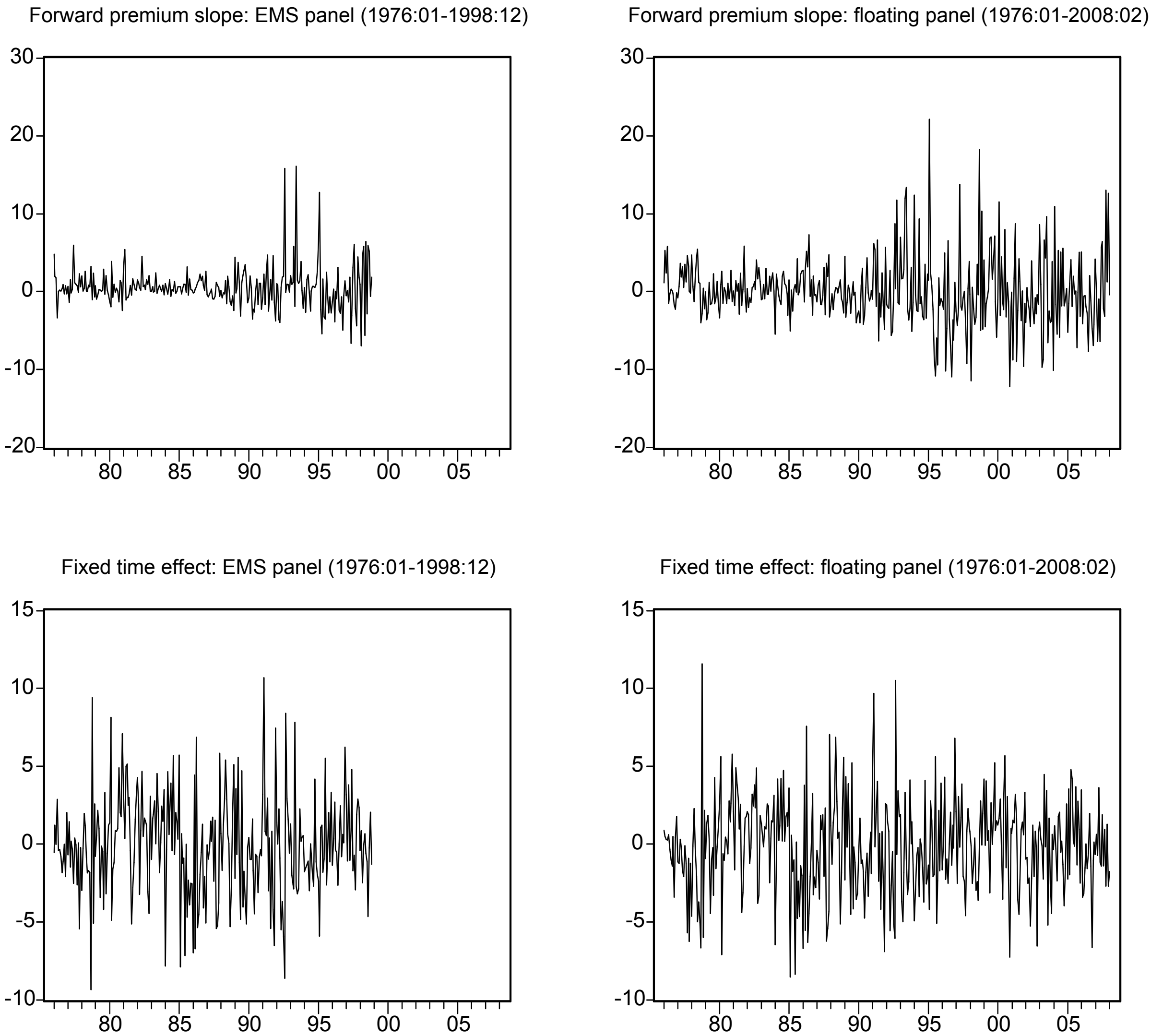

FIGURE 2. Empirical news distributions and descriptive statistics (EMS panel and floating panel)

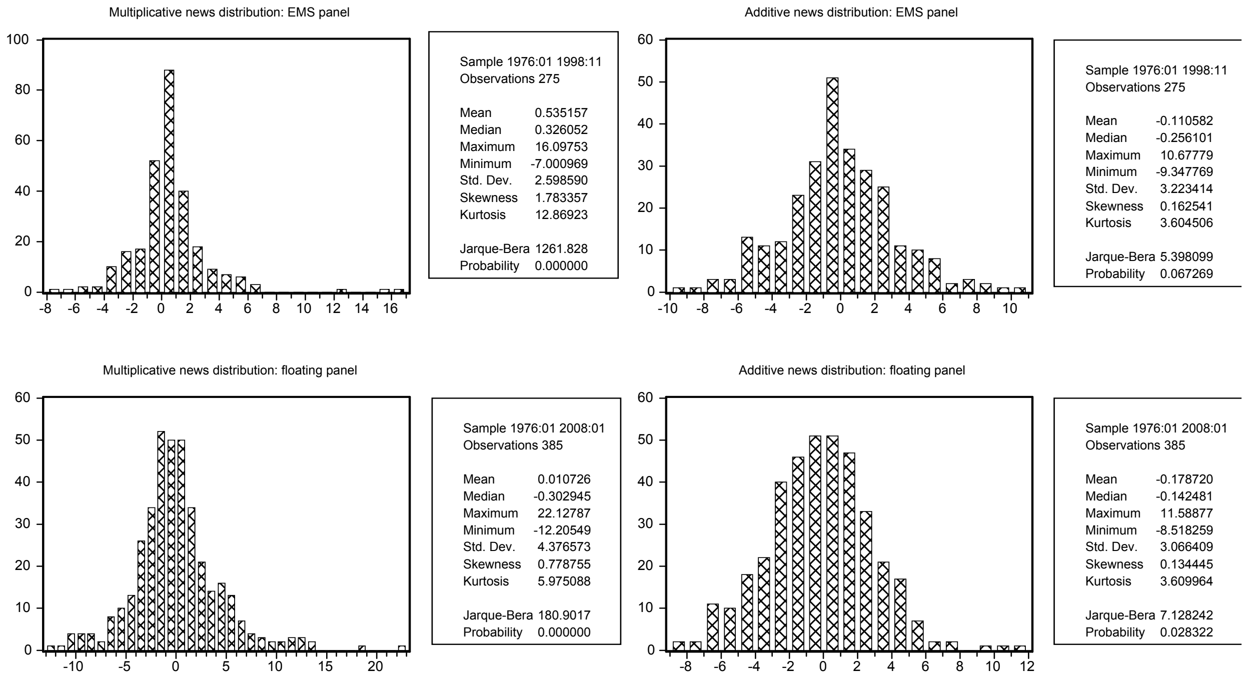

Figure 3. Forward premium slopes sorted on ascending average absolute forward premiums and ascending forward premium variances

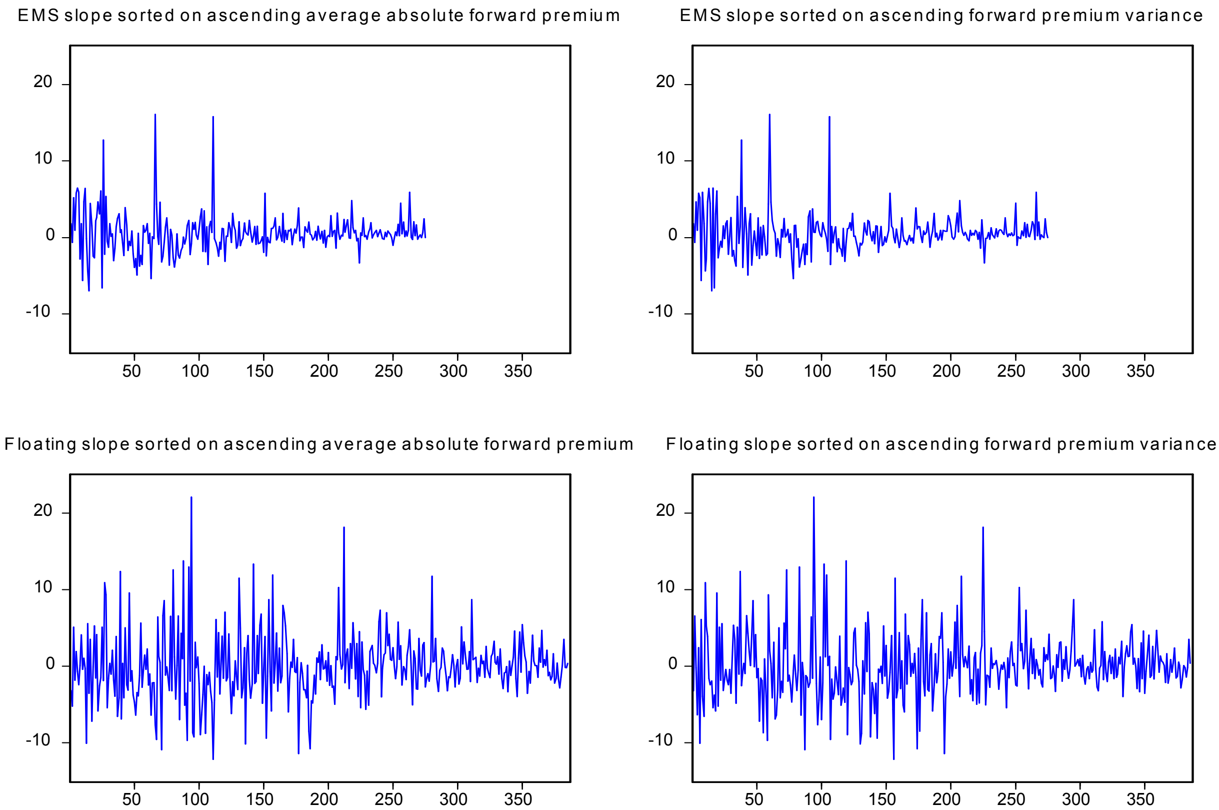

FIGURE 4. Cross sectional average and standard deviation of EMS and floating forward premiums