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Common Volatility Trends in the Central and Eastern European Currencies and the Euro

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Common Volatility Trends in the Central and Eastern European Currencies and the Euro

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Abstract

How much convergence has been achieved between Central and Eastern European (CEE) economies and the eurozone? We explore this question by comparing long-run volatility trends in CEE currencies and the euro. We find that these trends are closely correlated, pointing to convergence in the economic and financial structures of these economies. Nonetheless, the degree of commonality remains weaker than what had been found for major European currencies before the introduction of the euro. Spillovers of volatility across regional markets appear to have diminished over time, with the exception of the Hungarian forint, which remains a source of volatility shocks to regional currencies.

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I. Introduction

The role of the exchange rate in adjustment to shocks lies at the heart of the optimum currency area theory (Mundell, 1961). Under a flexible exchange rate regime, the exchange rate can help buffer the economy from external shocks.[2] In a currency union, where the nominal exchange rate between two currencies disappears, adjustment would have to take place through relative prices in the two economies. A limited degree of volatility in the bilateral exchange rate would thus suggest that the two economies have achieved a sufficient degree of convergence in their economic and financial structures and face broadly similar shocks for a common monetary policy to be sustainable.[3]

Exchange rate volatility as such has received less attention in the literature on optimum currency areas than the analysis of business cycle correlations and equilibrium exchange rates. A few exceptions are studies by Harvey and others (1994), Klaassen (1999), Black and McMillan (2004), and Horváth (2005). Black and McMillan (2004), in particular, examined the degree of commonality in exchange rate volatility trends for industrial countries and have identified a strong long-run volatility trend in European currencies (the Deutsche mark, French franc, Italian lira, and the British pound sterling) for the period from 1974 to 1998. Using a larger sample of developed countries, Horváth (2005) found that countries fulfilling the optimum currency area criteria (including trade linkages and business cycle synchronization) tend to experience less bilateral exchange rate volatility. In particular, the author confirmed that this conclusion applies to Central and Eastern European (CEE) countries—the Czech Republic, Hungary, Poland, the Slovak Republic, and Slovenia—and the eurozone.

This study also focuses on the analysis of exchange rate volatility trends in CEE currencies and the euro. The literature on business cycle convergence concludes that the CEE economies have achieved a considerable degree of integration with the eurozone, although less than exists among the current core members of the eurozone.[4] We explore if a comparison of long-run volatility trends in CEE currencies and the euro would render a similar conclusion. A finding of a common long-run volatility trend in CEE currencies and the euro would suggest similarity in the underlying economic and financial structures, the shocks faced by these economies, and mechanisms for the transmission of these shocks through the currency markets. In line with the existing literature on the business cycle convergence of CEE economies, we would expect the degree of commonality in the volatility trends of CEE currencies and the euro to be weaker than what has been found for the original members of the European Union prior to the introduction of the euro.

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[2] See Obstfeld (2002).

[3] This is the main rationale for including the exchange rate convergence criterion among the Maastricht criteria for adopting the euro. As outlined in the Treaty on European Union, fulfilling the exchange rate convergence criterion requires participation in ERM II and maintaining exchange rate stability against the euro.

[4] See Fidrmuc and Korhonen (2004) for an overview of this literature.

Our research approach is twofold:

- First, we identify the stylized facts concerning exchange rate volatility in CEE currencies. We decompose exchange rate volatility into a long-run trend and a transitory component using the Component-GARCH model developed by Engle and Lee (1993). The decomposition, originally proposed by Beveridge and Nelson (1981) for the analysis of business cycles, has been found useful in the analysis of exchange rate volatility (Black and McMillan, 2004; and Byrne and Davis, 2005). The two components of volatility are typically interpreted as driven by different factors: the long-run trend of volatility as reflecting shocks to economic fundamentals, and transitory volatility as driven by market sentiment and short-term position-taking. We include an asymmetric term in the model to test for differences in volatility associated with exchange rate depreciations and appreciations.

- Next, we examine principal components and pairwise correlations between currencies for evidence of common volatility trends. To check for robustness of the results based on principal component analysis, we test for the presence of volatility spillovers across currency markets. Volatility spillovers have also been known as "meteor showers," after the original paper by Engle, Ito, and Lin (1990), which found evidence of volatility spillovers in well-integrated currency markets for major currencies.[5] We would expect to find volatility spillovers in CEE currency markets, particularly in light of the finding by Kóbor and Székely (2004) that there is strong correlation between the Hungarian forint and the Polish zloty, and between the Czech koruna and the Slovak koruna during high-volatility periods.[6]

We find that the volatility patterns in CEE currency markets are broadly similar to those observed in other mature and emerging market currency markets (Byrne and Davis, 2005; Guimarães and Karacadag, 2004; and Black and McMillan, 2004). The long-run volatility component outweighs the transitory component, suggesting that exchange rate volatility is mainly driven by shocks to economic fundamentals rather than shifts in market sentiment. The degree of persistence in the exchange rate volatility of CEE currencies is fairly high, often exceeding that in mature currencies, but has been declining over time. There is evidence of asymmetric effects in the volatility of CEE currencies: depreciations are often associated with higher volatility than appreciations.

The principal component and correlation analyses confirm on-going economic and financial convergence of CEE countries and the eurozone: a common long-run volatility trend in CEE currencies is found to be correlated with the long-run volatility trend in the euro for the

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[5] See also Melvin and Melvin (2003).

[6] A finding of volatility spillovers across CEE currency markets would increase the likelihood of "bandwagon" effects and contagion in a financial crisis. Measurement of contagion effects and prediction of financial crises in the CEE region are beyond the scope of this paper, which focuses on optimum currency area issues. For a review of the empirical literature on contagion in emerging markets, see Dungey and others (2004).

period from January 1997 to June 2005. This result is consistent with econometric evidence presented in Horváth's (2005) cross-country study, which uses a simpler statistical measure of volatility. As expected, the degree of commonality is less than what Black and McMillan (2004) found for major industrial countries in Europe before the introduction of the euro. Among the Central and Eastern European currencies, volatility in the Slovak koruna appears to be most closely related to that in the euro, while volatility in the Polish zloty the least.

The transmission of volatility shocks within the region appears to have changed over time. The incidence of spillovers from one currency's volatility into the volatility of other currencies has declined over time, possibly reflecting increased country differentiation on the part of investors. Only the Hungarian forint has consistently remained an important source of intraregional volatility shocks throughout the last decade. In addition, we find ample evidence of spillovers of volatility into means, which suggests strong cross-linkages among regional currency markets and some degree of predictability in returns, possibly owing to limited liquidity in CEE currency markets.

The rest of the paper is organized as follows: Section II describes the methodological approach used in the study, focusing on the description of the Component-GARCH model and Wald tests for volatility spillovers. The section also describes the data set. Section III discusses the findings of the study: the relative importance of the long-run and transitory components of volatility, common trends in these components, and volatility spillovers. Section IV concludes.

II. Methodology and Data

A. Volatility Decomposition and Common Trends

Our analysis of exchange rate volatility is cast within the generalized autoregressive conditional heteroskedasticity (GARCH) class of models introduced by Engle (1982) and Bollerslev (1986). These models have been designed to capture the volatility clustering observed in financial time series, including exchange rates. GARCH models focus on the conditional variance of the underlying series by identifying and measuring the degree of autocorrelation in second moments.

We use a specification known as Component-GARCH (CGARCH), which decomposes volatility into two components—a stochastic long-run trend and short-run deviations from that trend. The model is described by the following set of equations:

$$x_t = a_0 + a_1 x_{t-1} + \varepsilon_t + b_1 \varepsilon_{t-1}, \qquad \varepsilon_t \mid I_{t-1} \sim N(0, h_t^2), \tag{1}$$

$$h_t^2 = q_t + \alpha_1(\varepsilon_{t-1}^2 - q_{t-1}) + \gamma(\varepsilon_{t-1}^2 - q_{t-1})D_{t-1} + \beta_1(h_{t-1}^2 - q_{t-1}), \tag{2}$$

$$q_t = \omega + \rho q_{t-1} + \varphi(\varepsilon_{t-1}^2 - h_{t-1}^2), \tag{3}$$

where $D_t = 1$ for $\varepsilon_t < 0$, $D_t = 0$ otherwise. Equation (1) is the mean equation, where x_t is the log-difference and hence the continuously compounded rate of return of daily exchange rates. The term ε_t reflects any unexpected appreciation or depreciation, which is assumed to be uncorrelated and conditionally normally distributed, given I_{t-1}, the information set available at time t-1. The mean equation also includes AR(1) and MA(1) terms.[7]

Our main interest lies in the conditional variance in equations (2) and (3). By analogy with the GARCH(1,1) setup, this equation models the conditional variance (h_t^2) as a linear function of a time-dependent intercept, the lag in the squared realized residual (the so-called ARCH term), an asymmetric term that augments the ARCH term whenever a lagged residual is negative, and the lagged conditional variance (labeled the GARCH term). The ARCH, GARCH, and asymmetric terms are all specified as deviations from the long-run trend of conditional variance.

The model allows for asymmetric effects on volatility of currency appreciation and depreciation. In line with the literature (Engle and Lee, 1993; and Byrne and Davis, 2005; for example), we include an asymmetric term in the model—through a dummy variable (D_t) that takes the value '1' for negative realized residuals.

The distinctive feature of the CGARCH setup is equation (3), which explicitly models the time-varying long-run component of conditional variance. This component consists of a time-invariant permanent level (ω), an AR term (ρ), and the so-called forecast error (φ), which is the difference between the lag in the squared realized residual and the forecast from the model (based on information available at time t-2). The long-run component is allowed to vary over time in response to the forecast error, but, as equation (3) shows, it converges to the time-invariant unconditional level, provided $|\rho| < 1$. Given that the long-run component is fully accounted for by q_t, the short-run component of conditional variance is described by the right-hand side of a rearranged version of equation (2):

$$h_t^2 - q_t = \alpha_1(\varepsilon_{t-1}^2 - q_{t-1}) + \gamma(\varepsilon_{t-1}^2 - q_{t-1})D_{t-1} + \beta_1(h_{t-1}^2 - q_{t-1}). \tag{4}$$

The (unconditional) expectation of each of the three terms in the short-run component is equal to zero, implying that transitory volatility will converge to zero over time and aggregate volatility converges to its long-run trend. The condition for these volatility dynamics to hold is that the short-run component of volatility converge faster than the long-run component: $(\alpha_1 + \beta_1) < \rho$. The sum of the coefficients α_1 and β_1 is also referred to as the half-life of (positive) shocks and is used as a measure of volatility persistence.

A number of restrictions need to be satisfied in this model to ensure that the conditional variance is nonnegative for out-of-sample forecasts: (i) $1 > \rho > (\alpha_1 + \beta_1) > 0$, (ii) $\beta_1 > \varphi > 0$, and (iii) $\alpha_1, \omega > 0$. In addition to specifying the relative speed of convergence of the volatility components, restriction (i) rules out a random walk for the long-run component. In practice, a unit root is frequently observed in the data. This finding does not invalidate estimation results, but calls for caution when using parameter estimates for forecasting purposes to

[7] We determine the appropriate lag structure of the mean equation for each currency based on the Schwarz Information Criterion (BIC) and other regression diagnostics. We have tested for higher-order AR, MA, and ARMA effects in each currency model, but the best fit resulted universally from an AR(1), MA(1), or AR(0) structure.

avoid obtaining negative estimates of variance. Restrictions (ii) and (iii) impose strict positivity on all regression parameters, except for the asymmetric term.

Engle and Lee (1993) show that the CGARCH setup is essentially a GARCH(2,2) model. Such a more general model is less restrictive than a GARCH(1,1) specification, and in the case of over-specification, reduces to the simpler GARCH(1,1) setup. Conditions for this are as follows: (i) $\rho = \varphi = 0$, or (ii) $\alpha_1 = \beta_1 = 0$. If both ρ and φ are equal to zero, as in (i), the CGARCH model will reduce to the standard GARCH(1,1) setup with a constant long-run volatility trend and only short-run dynamics around this trend. If (ii) holds, with both α_1 and β_1 equal to zero, the resulting specification will differ from the standard GARCH model in that it takes into account only the long-run component of volatility, allowing it to vary over time.

We implement the CGARCH model in a univariate manner, that is, for individual currencies. An alternative, multivariate GARCH (MGARCH) approach would have the advantage that it can explicitly account for cross-currency spillovers in the volatility equation, but at the cost of not being robust to the ordering of series or requiring restrictions inconsistent with the purposes of our study.[8] We estimate the model using the quasi-maximum likelihood method and compute Bollerslev-Wooldridge robust standard errors. The robust errors tend to be larger than non-robust errors and present an appropriately more rigorous basis for hypothesis testing (Bollerslev and Wooldridge, 1992).

In the next step, we use principal component analysis to identify common trends in the long-run and short-run volatility components for CEE currencies and the euro. We also examine pairwise correlations as a cross-check and a guide for interpreting the results of the principal component analysis.

B. Volatility Spillovers

As a robustness check of the results concerning common trends in volatility, we also test for cross-currency volatility spillovers. The presence of meteor showers or volatility spillovers across currency markets (Engle, Ito, and Lin, 1990) would be consistent with rising financial and economic integration and would imply a greater likelihood of bandwagon effects and contagion across these markets. To identify volatility spillover effects, we include the lagged variance series of another currency in the variance equation for the trend or transitory

[8] In its most general and flexible specification, the so-called VEC model of Bollerslev, Engle and Wooldridge (1988), working with six series would require the estimation of so many parameters, even without the CGARCH enhancement, that the significance of the parameter estimates would be severely reduced. The problem of a lack of degrees of freedom can be overcome in more restricted multivariate specifications, such as the BEKK model proposed by Engle and Kroner (1995). However, the resulting specification is unlikely to be robust to the ordering of the series, and the number of parameters to be estimated still remains large. Severely restricted specifications, such as the constant conditional correlation model by Bollerslev (1990), sufficiently restrict the number of parameters, but the assumption of constant correlations would be hard to defend in our study. Allowing correlations to change over time, as in the dynamic conditional correlation models by Engle (2002) and Tse and Tsui (2002), imposes identical dynamics on all conditional correlations, which is also inappropriate in the context of our study. For more details, see the survey paper by Bauwens, Laurent, and Rombouts (2003).

component of volatility. For spillover effects into the long-run component of conditional variance, we adjust equation (3) by including the lagged conditional variance:

$$q_t = \omega + \rho q_{t-1} + \varphi(\varepsilon_{t-1}^2 - h_{t-1}^2) + \delta_{k,j} h^2_{k,t-1}. \quad (5)$$

Likewise, causality in the transitory component of conditional variance is tested by modifying equation (2):

$$h_t^2 = q_t + (\alpha_1 + \gamma D_{t-1})\times(\varepsilon_{t-1}^2 - q_{t-1}) + \beta_1(h_{t-1}^2 - q_{t-1}) + \delta_{k,j}h^2_{k,t-1}, \quad (6)$$

where D_t is defined as before: $D_t = 1$ for $\varepsilon_t < 0$, $D_t = 0$ otherwise.

Besides testing for meteor showers, we test for spillovers of volatility into means, whereby higher volatility in one market might lead to a change in the level of the exchange rate in the same or another market. Evidence of such volatility-mean spillovers would imply existence of a time-varying risk premium and predictability in exchange rates, which would be inconsistent with the market efficiency hypothesis (Fama, 1970 and 1991). To test for causality in mean, we change the mean equation (1) by including the lagged conditional standard deviation of either the same or a different currency:

$$x_{j,t} = a_{j,0} + a_{j,1}x_{j,t-1} + \varepsilon_{j,t} + b_{j,1}\varepsilon_{j,t-1} + \delta_{k,j}h_{k,t-1}. \quad (7)$$

The setup with the conditional standard deviation included in the mean equation bears close resemblance to the ARCH-in-mean (ARCH-M) specification, often used to test for the presence of time-varying risk premia in financial markets. A significant coefficient will suggest that the level of volatility has an impact on the price of the currency in question, but, given that we use lagged variance series, such a finding will clearly imply return predictability.

In each case, we perform a Wald test for the significance of $\delta_{k,j}$. For the causality-in-mean setup, we can test for significant spillover effects from all of the six currencies to a given currency, because currencies may be affected by their own lagged volatility (as in the original ARCH-M specification). When we test for causality in variance, however, we can only include the lagged conditional variance of another currency, as the own lagged conditional variance is by definition already included in both parts of the variance equation of the CGARCH model.

C. Data

Our focus is on CEE currencies and the euro. The currency series consist of daily closing prices for the Czech koruna (CZK), the Hungarian forint (HUF), the Polish zloty (PLN), the Slovenian tolar (SIT), the Slovak koruna (SKK), and the euro (EUR), all of which are quoted as U.S. dollar rates (USD). The data source is WM/Reuters, as reported by Datastream. Prior to 1999, the EUR series is reconstructed through the DEM/USD rate, which is divided by 1.95583, the fixed DEM/EUR conversion rate. The exchange rate data are shown in Figure 1.

The sampling period covers the time period during which CEE countries had an exchange rate regime flexible enough to render the analysis meaningful (Borghijs and Kuijs, 2004). For the Czech Republic, the sample period starts in February 1996, when the authorities replaced the exchange-rate peg with a band of ± 7.5 percent. The sample period for Hungary starts in March 1995, when the ± 2.25 percent exchange-rate band was introduced.[9] Poland introduced a crawling exchange-rate band in May 1995. In the Slovak Republic, the crawling band was widened to ± 7 percent in early 1997. There are two exceptions to this sample dating approach: the EUR has been flexible during the whole period in question, while the SIT has been significantly managed for most of the time. The series for these two currencies thus start in January 1993. All six series end in June 2005. The sample period common to all six currencies is January 1997–June 2005. We also compare estimates for the earlier part of the sample period (January 1997–June 2001) and the later part (July 2001–June 2005).

All data series display a unit root, as shown in Table 1. Hence, we transform them into log-differences and obtain continuously compounded exchange-rate returns in percentage terms: $x_t = 100[\ln(S_t) - \ln(S_{t-1})]$, where S_t is the spot rate.

III. Volatility Dynamics in Central and Eastern European Currency Markets

A. Is Volatility in Central and Eastern European Currencies of Long-Run or Transitory Nature?

Using the CGARCH model described in Section II, we decompose exchange rate volatility into a long-run component and a transitory component (Table 2 and Figure 2). In the long-run component of volatility, we find a positive and highly significant constant (ω) for all currencies. The AR coefficient of permanent volatility (ρ) is large and highly significant for all currencies in all periods.[10] Its size exceeds that of the coefficients in the transitory component ($\alpha_1 + \beta_1$) in all instances, implying the model is stable. The degree of volatility persistence found in CEE currencies is higher than that in the euro, but broadly in line with what has been found for other industrial economies (Byrne and Davis, 2005; Black and McMillan, 2004). The coefficient of the forecast error (φ), which shows how shocks affect the permanent component of volatility, is positive in all regressions and generally significant. An interesting pattern is that, for most currencies, the AR coefficient of long-run volatility is smaller in the late period than in the early period, implying that over time long-run volatility tends to revert to its time-invariant level faster, possibly because in the later period the exchange rates were allowed to fluctuate more freely. The signs and relative magnitudes of coefficients confirm that the CGARCH model is well specified and is an appropriate framework for analyzing volatility patterns in CEE currencies.

As expected, the combined coefficient for the short-run component of volatility ($\alpha_1 + \beta_1$) is positive and smaller than that for the long-run component (ρ). In a few instances, we find a

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[9] While the degree of flexibility is still limited in this regime, a widening of the band to ± 15 percent occurred only in 2001, which would have reduced our observation period so significantly that a comparison with the other countries would have been difficult.

[10] In several instances, we find that the coefficient of the autoregressive term in the trend equation is equal or very close to one. As discussed above, this suggests that the long-run component follows a random walk and that out-of-sample forecasting needs to be handled with care to ensure the non-negativity of variance estimates.

significant negative coefficient for the ARCH term (α_1), but even in these cases, the sum of the coefficients on the ARCH and GARCH terms (α_1 and β_1) is still positive. In cases where both α_1 and β_1 are insignificant, a Wald test generally cannot reject the hypothesis that both coefficients are jointly equal to zero. Together with significant coefficients on the forecast error in the long-run component, this implies that in those instances (specifically, the HUF and the SKK in the early period) shocks to the exchange rate were mostly of a long-run nature. The opposite holds for the CZK and the PLN in the early period, where shocks to volatility appear largely transitory, as the coefficient on the forecast error in these cases is insignificant.

For most currencies, short-run volatility is hardly persistent. This is reflected in the relatively short half-life of these shocks—about one day. Notable exceptions are the CZK and the PLN in the early period and the HUF in the late period. Higher persistence of short-run shocks in these cases (half-life exceeding 5 days) reflects episodes of turbulence in currency markets. Short-run volatility persistence has declined for the CZK and the PLN since then—their half-life was less than one day in the later part of the sample.

We find significant negative asymmetric effects (γ) for several CEE currencies, particularly the HUF in the late period and the PLN in the early period. Since the exchange rates are defined as domestic currency per U.S. dollar, a negative coefficient implies higher volatility in cases of currency depreciation. This would suggest that long and short positions in these currencies were not evenly enough distributed so that the market as a whole had a unidirectional view on the currency. This finding is in line with the literature: Byrne and Davis (2005), for instance, find a similar effect of unexpected depreciations for the Japanese yen and the Canadian dollar, while Guimarães and Karacadag (2004) find significant asymmetric effects for the Mexican peso and the Turkish lira.[11]

For all currencies and periods, the short-run component of volatility is much smaller than the long-run component (Figure 2 and Table 3). This suggests that transitory shifts in financial market sentiment tend to be less important determinants of exchange rate volatility than shocks to the underlying fundamentals. A similar pattern has been observed in currencies of industrial countries (Black and McMillan, 2004; and Byrne and Davis, 2005). Yet, relative to its lower mean level, the transitory component is in all cases much more volatile than the long-run trend level of volatility, as one would expect. For several currencies—the CZK and the PLN in the early period, and the HUF in the late period, the standard deviation of the short-run component exceeds that of the long-run component, reflecting periods of temporary turbulence in these markets. When scaled by means, the standard deviations of the short-run component exceed those for the long-run component (the third column in Table 3). We now turn to the analysis of comovement in the long-run and short-run components of volatility.

[11] Like Byrne and Davis (2005), we find that the asymmetric effect is insignificant for the euro. Its inclusion weakens the overall fit and stability properties of the model, and hence we exclude the asymmetric effect from the baseline specification for the euro. Our results are robust to the inclusion of the asymmetric effect.

B. Is There a Common Volatility Trend in Central and East European Currencies and the Euro?

Next, we explore the degree of similarity in the volatility trends of CEE currencies and the euro. Principal component analysis of the long-run volatility components suggests a fairly high degree of comovement between CEE currencies and the euro. In particular, for the recent period, the weights on the first component are similar in sign and absolute value for CEE currencies and the euro, which can be interpreted as evidence of a common underlying trend in CEE currencies and the euro (Table 4 and Figure 3).

The degree of similarity in the long-run volatility trends of CEE currencies and the euro is somewhat less than what Black and McMillan (2004) found for major industrial countries prior to the introduction of the euro. In their paper, deviations between the weights on the principal components for different European currencies (the French franc, Deutsche mark, Italian lira, and British pound) are smaller than what we find for CEE currencies and the euro. However, a broad measure of commonality in volatility trends, the share of variation explained by the first principal component, is quite similar to that observed for mature European currencies (about 55 percent) (Table 4). Country-specific findings of the principal component analysis are:

- In the early half of the sample, the common factor for the HUF, PLN, SIT, and the EUR almost entirely leaves out the SKK, which instead is the sole driver behind the second principal component, and the third component picks up a significant portion of volatility in the CZK. In the later half of the sample period, the weights on the first principal component are more evenly distributed among CEE currencies, so all these currencies appear to share a common long-run volatility trend. Notably, the SKK is not an outlier anymore, and instead shares a common component with other regional currencies.

- The degree of commonality in the long-run trends of the PLN and other CEE currencies is weaker than in the long-run trends of these other CEE currencies. The PLN appears to react differently to shocks than other CEE currencies, consistent with the findings in Borghijs and Kuijs (2004), who show that the shock-absorbing role of the PLN differs from that of other regional currencies. The second component is strongly correlated with the PLN and HUF, suggesting close linkages in these currencies.

- Interestingly, both in the early and later period, a significant portion of volatility in the CZK can be explained by factors other than those influencing the other CEE currencies. This finding might reflect the role of the CZK as a funding currency for investments in other CEE currencies and the high liquidity of the Czech koruna market, the development of which has been facilitated by a relatively more rapid liberalization of capital controls in the Czech Republic than in other countries in the region.

Pairwise correlations for the long-run volatility component broadly confirm the findings of the principal component analysis (Table 5). Bilateral correlations of the SKK and the CZK with the EUR increase in the second half of the sample, while those of the PLN, the HUF, and the SIT decline, so that on the whole the cross-country differences in the degree of correlation with the EUR are smaller in the later period. As expected, pairwise correlations between the long-run volatility of CEE currencies and the euro appear weaker that those observed between currencies of major industrial countries in Europe. Black and McMillan (2004), for example, find correlation between the French franc and the Deutsche mark of 0.90. By comparison, correlations between CEE currencies and the euro hardly exceed 0.60 for the three major CEE currencies—the PLN, the CZK, and the HUF—although they are higher for smaller, and more managed, currencies such as the SIT and the SKK. Within the CEE region, we find strong correlations between the CZK, PLN, HUF, and SIT in the early period and between the PLN and the HUF, and the CZK and the SKK in the later period.

The principal component and correlation analyses for the short-run volatility component suggest that these components have less in common than the long-run components (Tables 6-7). The dispersion and overall variability of weights for the short-run component are significantly higher than for the long-run component. This is not surprising, as the short-run component of volatility reflects transitory and unsystematic disturbances, and is in line with findings for major industrial countries reported by Black and McMillan (2004). Only the SKK and the EUR show a significant common trend in all periods, but even that relationship is not stable as reflected in the changing signs in the early and late periods in the weights on the first principal component as well as in the correlations. In the early period, the HUF also shares a common factor with the EUR, and in the late period the same is true for the CZK. Despite the variability in the relationship of the short-run volatilities for individual currencies, as a group, they show that common factors increasingly drive transitory volatility. This is reflected in the higher proportion of variance accounted for by the first principal component in the late period compared to the early period (40 percent versus less than 30 percent, respectively) and the cumulative proportion of variance explained by the first three principal components (76 percent versus 66 percent, respectively).

C. How Significant Are Volatility Spillovers Between CEE Currency Markets and the Euro-Dollar Market?

The Wald tests for volatility spillovers confirm the findings of the principal component and correlation analyses. The tests show that long-run volatility spillovers from the EUR to the CZK and the SKK become significant in the second half of the sample (Table 8)—these are the two currencies for which the principal component and correlation analyses show that the long-run volatility trends have become more similar to those in the euro over time. Likewise, spillovers from the EUR to other currencies (particularly, the PLN) are weaker in the later part of the sample, in line with the principal component and correlation analyses. Tests for volatility spillovers also show that volatility in CEE currencies is increasingly driven by common shocks affecting the region as a whole. The transmission of these shocks within the region appears limited: the number of significant intraregional spillovers in both long-run and short-run volatility has declined over time (Tables 8-9). While we find significant volatility spillovers between most CEE currencies in the early period, only the HUF remains an important source of volatility spillovers to the PLN and the SKK in the later period.

Intraregional spillovers of volatility into means have become more frequent over time, implying that the degree of integration of CEE currency markets has increased (Table 10). These results also suggest a relatively high degree of predictability in returns on CEE currencies, possibly reflecting limited efficiency and liquidity of these markets. Interestingly, although the PLN does not affect any other currencies in the early period, it has a significant impact on most of them in the late period, which could be indicative of the increased importance of the PLN in the region as the country has removed the remaining capital controls. The mean return on the CZK is found to respond strongly to volatility in the SKK in the later period, confirming strong links between these two currencies. The volatility in the EUR also has a significant effect on the mean returns of CEE currencies in a few instances.

IV. Conclusions

This paper seeks to complement the existing analyses of common trends in economic activity in CEE countries and the eurozone with an analysis of common trends in exchange rate volatility of the respective national currencies. We find that volatility dynamics of CEE currencies and the euro are indeed similar, suggesting growing economic and financial integration of the region with the eurozone. The degree of convergence implied by the commonality of the long-run volatility trends is smaller than what had been achieved between mature European economies prior to the introduction of the euro, which is broadly in line with conclusions of earlier studies on this topic (Fidrmuc and Korhonen, 2004; and Horváth, 2005).

The degree of commonality in the long-run volatility trends of the CEE currencies and the euro varies across countries. Volatility in the Slovak koruna appears to be most closely related to that in the euro. The degree of similarity is smaller for the Czech koruna, the Hungarian forint, and the Slovenian tolar, while volatility in the Polish zloty is correlated with the euro the least. The finding that the long-run volatility trend in the Polish zloty differs somewhat from that in other CEE currencies is consistent with Borghijs and Kuijs's (2004) conclusion that in Poland the exchange rate plays a more significant role as a shock absorber than in other economies in the region, possibly reflecting the relatively large size of the Polish economy and the smaller degree of trade openness, compared to the neighboring countries. The degree of commonality in the long-run volatility trend of the Slovak koruna and other CEE currencies and the euro appears to have increased in recent years, suggesting increasing regional integration of the Slovak economy. Lastly, volatility in the Czech koruna appears to be driven to a certain extent by factors other than those influencing other CEE currencies, possibly reflecting the role of the Czech koruna as a funding currency for investments in other CEE currencies.

We find mixed evidence on the role of regional currency markets in transmitting volatility across regional currency markets. Intraregional volatility spillovers appear to have diminished over time, although the Hungarian forint remains a source of volatility shocks in regional currency markets. In contrast, prevailing spillovers of volatility into means suggest cross-linkages among regional currency markets and some degree of predictability in returns, possibly because of limited liquidity in these markets.

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Figure 1. Daily Dollar Exchange Rates, January 1997-June 2005

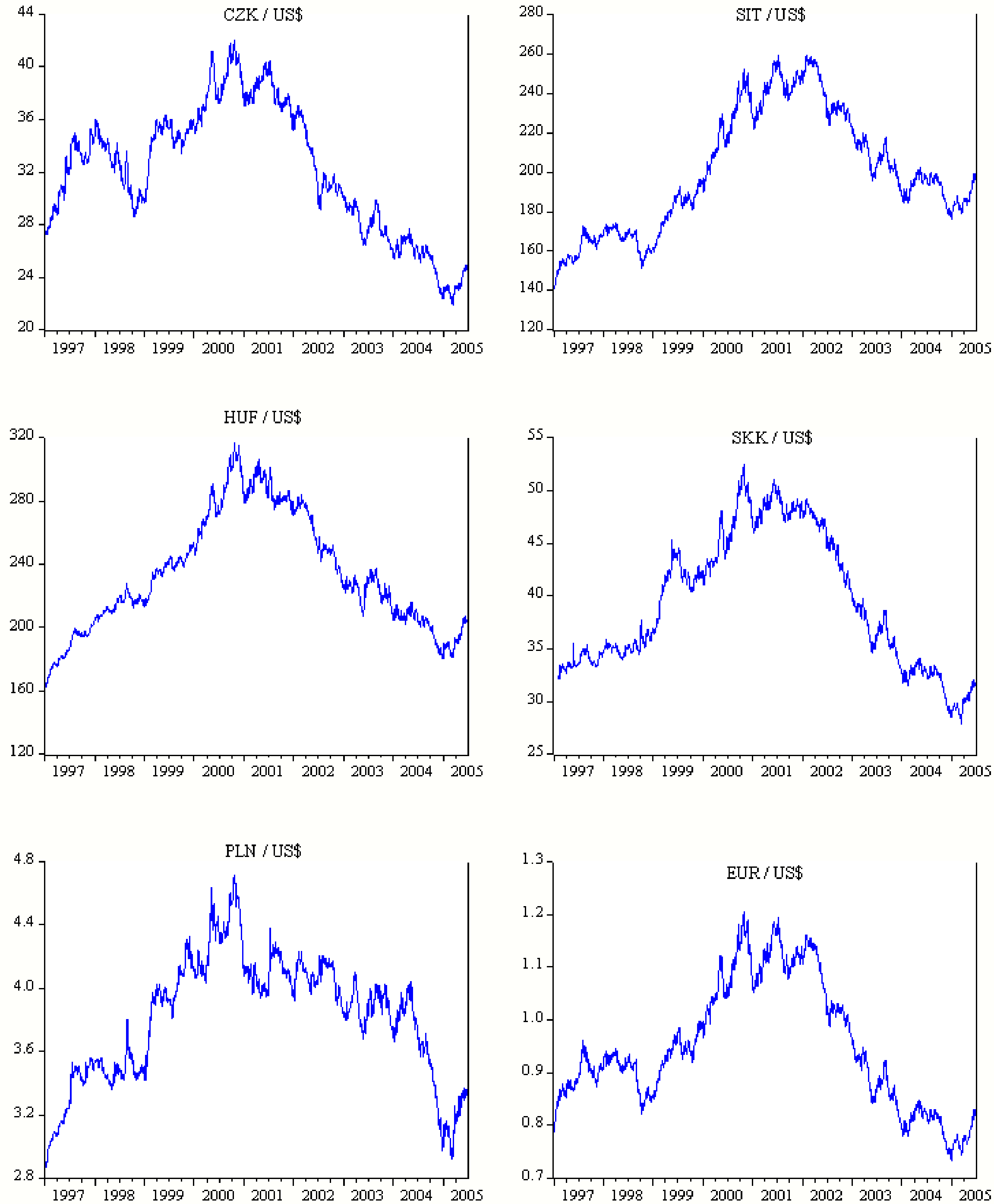

Figure 2. Conditional Variance of Daily Dollar Exchange Rates, January 1997-June2005

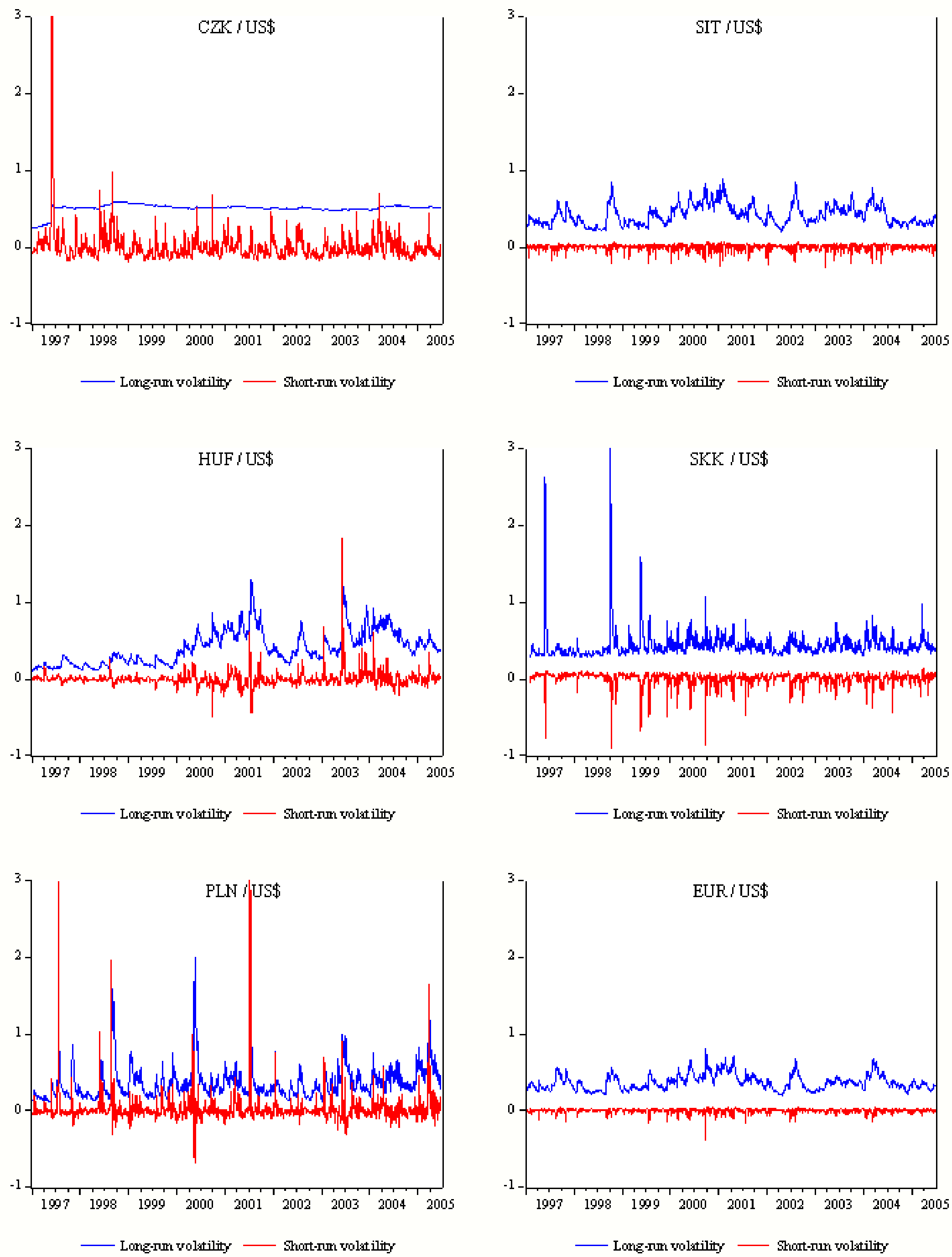

Figure 3. Eigenvalues for the First Principal Component, January 1997-June 2005

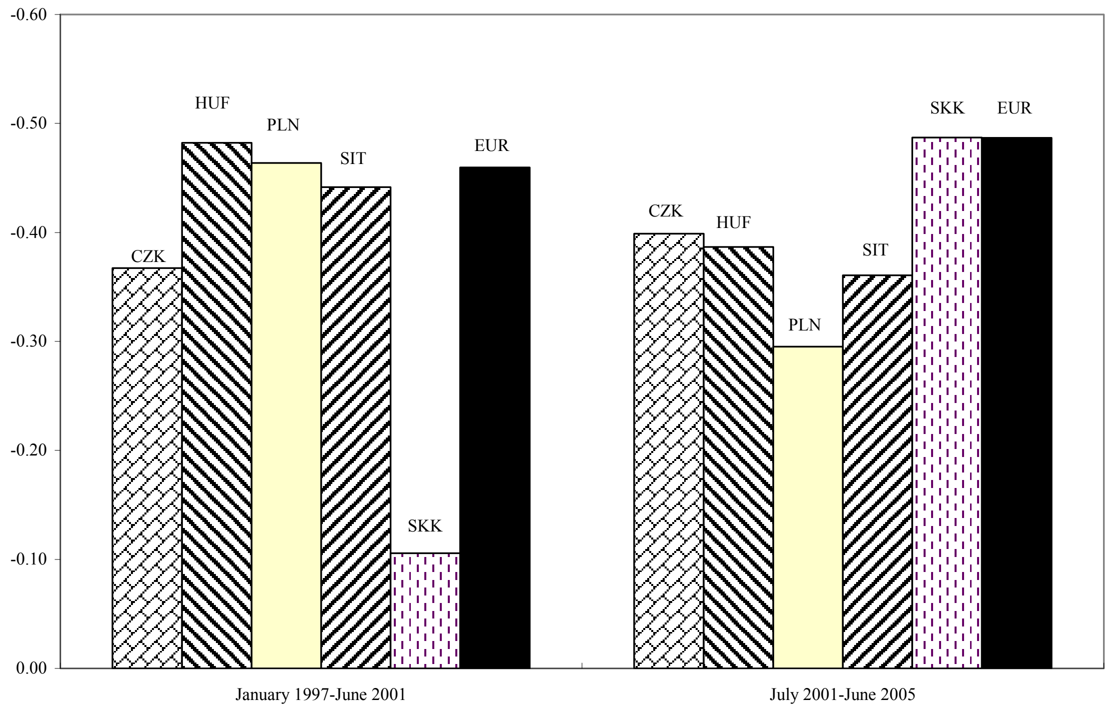

Table 1. Unit Root Tests

		CZK	HUF	PLN	SIT	SKK	EUR
Levels							
Test:	ADF	-0,96	-0,38	-0,90	-0,93	-1,25	-1,06
	PP	-1,02	-0,34	-0,81	-0,89	-1,25	-1,07
First Differences							
Test:	ADF	-49,92 ***	-52,94 ***	-47,92 ***	-58,29 ***	-45,73 ***	-56,88 ***
	PP	-49,89 ***	-52,90 ***	-47,78 ***	-58,31 ***	-45,73 ***	-56,88 ***

Source: Authors' estimates.

For both the Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests, the null hypothesis is for the existence of a unit root.

*, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent levels, respectively.

Table 2. Asymmetric Component GARCH Estimates

The Full Period: January 1997-June 2005

		CZK	HUF	PLN	SIT	SKK	EUR
Trend Intercept	ω	0.558 ***	0.357 **	0.349 ***	0.427 ***	0.434 ***	0.371 ***
		(5.60)	(2.09)	(5.80)	(8.00)	(13.10)	(10.59)
Trend AR Term	ρ	0.998 ***	0.996 ***	0.968 ***	0.985 ***	0.830 ***	0.983 ***
		(1,101.3)	(383.9)	(78.55)	(151.5)	(5.00)	(127.3)
Forecast Error	φ	0.003 *	0.040 ***	0.098 ***	0.041 ***	0.076	0.037 **
		(1.84)	(5.38)	(4.32)	(4.67)	(0.36)	(3.25)
ARCH Term	α_1	0.080 **	0.077 **	0.164 ***	-0.063 ***	-0.005	-0.036 *
		(2.27)	(2.34)	(3.31)	(-2.60)	(-0.02)	(-1.85)
Asymm. Term	γ	0.014	-0.126 ***	-0.240 ***	0.018	-0.100 **	
		(0.34)	(-2.95)	(-3.91)	(0.68)	(-2.19)	
GARCH Term	β_1	0.790 ***	0.774 ***	0.636 ***	0.504	0.768 **	0.681 *
		(12.16)	(8.56)	(5.63)	(1.44)	(1.94)	(1.89)
	$\alpha_1 + \beta_1$	0.87	0.85	0.80	0.44	0.76	0.64
Half-life (days) for	$\alpha_1 + \beta_1$	5.0	4.3	3.1	0.8	2.6	1.6

The Early Period: January 1997-June 2001

		CZK	HUF	PLN	SIT	SKK	EUR
Trend Intercept	ω	0.882 *	-1.197	0.065	0.438 ***	0.475 ***	0.375 ***
		(1.71)	(-0.10)	(0.09)	(4.99)	(5.45)	(7.10)
Trend AR Term	ρ	0.999 ***	1.000 ***	1.000 ***	0.989 ***	0.692 ***	0.989 ***
		(1,167.3)	(258.2)	(867.2)	(146.3)	(5.76)	(167.9)
Forecast Error	φ	0.003	0.040 ***	0.006	0.040 ***	0.203 **	0.026 ***
		(1.39)	(4.88)	(1.50)	(3.64)	(2.14)	(3.37)
ARCH Term	α_1	0.151 **	0.081	0.253 ***	-0.056 *	-0.111	0.013
		(2.39)	(1.59)	(4.16)	(-1.91)	(-1.23)	(0.48)
Asymm. Term	γ	-0.064	-0.086	-0.180 ***	0.048	-0.099	
		(-1.32)	(-1.27)	(-2.82)	(1.37)	(-1.51)	
GARCH Term	β_1	0.740 ***	-0.192	0.657 ***	0.595	-0.349	-0.309
		(9.09)	(-0.47)	(8.84)	(1.07)	(-1.47)	(-0.23)
	$\alpha_1 + \beta_1$	0.89	-0.11	0.91	0.54	-0.46	-0.30
Half-life (days) for	$\alpha_1 + \beta_1$	6.0	0.3	7.4	1.1	0.9	0.6

The Late Period: July 2001-June 2005

		CZK	HUF	PLN	SIT	SKK	EUR
Trend Intercept	ω	0.450 ***	0.474 ***	0.379 ***	0.405 ***	0.409 ***	0.347 ***
		(13.84)	(7.68)	(7.33)	(10.98)	(10.03)	(10.23)
Trend AR Term	ρ	0.875 ***	0.965 ***	0.946 ***	0.944 ***	0.955 ***	0.965 ***
		(8.00)	(52.99)	(48.13)	(27.08)	(30.64)	(44.69)
Forecast Error	φ	0.064	0.036 **	0.090 ***	0.060 *	0.060 *	0.043 **
		(1.15)	(2.00)	(3.84)	(1.72)	(1.68)	(2.22)
ARCH Term	α_1	-0.024	0.137 **	-0.078	-0.127 ***	-0.051	-0.092 ***
		(-0.32)	(2.43)	(-1.30)	(-3.68)	(-0.94)	(-3.22)
Asymm. Term	γ	-0.055	-0.264 ***	0.001	0.037	-0.070 *	
		(-0.75)	(-4.51)	(0.01)	(1.07)	(-1.71)	
GARCH Term	β_1	0.252	0.737 ***	0.130	0.694 ***	0.751 ***	0.566 *
		(0.36)	(10.34)	(0.23)	(2.67)	(3.54)	(1.85)
	$\alpha_1 + \beta_1$	0.23	0.87	0.05	0.57	0.70	0.47
Half-life (days) for	$\alpha_1 + \beta_1$	0.5	5.2	0.2	1.2	1.9	0.9

Source: Authors' estimates.

Bollerslev-Wooldridge robust t-statistics in parentheses.

*, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent levels, respectively.

Table 3. Comparison of Long-Run and Short-Run Volatility Components

The Full Period: January 1997-June 2005

	St. Dev. of Long-Run Component / St. Dev. of Short-Run Component	Mean of Long-Run Component / Mean of Short-Run Component	(St. Dev. / Mean) of L-R Comp. / (St. Dev. / Mean) of S-R Comp.
CZK	0.18	70	0.0025
HUF	2.22	368	0.0060
PLN	1.27	19	0.0653
SIT	3.45	785	0.0044
SKK	2.07	159	0.0130
EUR	3.54	70,259	0.0001

The Early Period: January 1997-June 2001

	St. Dev. of Long-Run Component / St. Dev. of Short-Run Component	Mean of Long-Run Component / Mean of Short-Run Component	(St. Dev. / Mean) of L-R Comp. / (St. Dev. / Mean) of S-R Comp.
CZK	0.17	33	0.0051
HUF	7.06	155	0.0454
PLN	0.31	7	0.0445
SIT	5.53	208	0.0266
SKK	1.85	282	0.0066
EUR	10.11	6,041	0.0017

The Late Period: July 2001-June 2005

	St. Dev. of Long-Run Component / St. Dev. of Short-Run Component	Mean of Long-Run Component / Mean of Short-Run Component	(St. Dev. / Mean) of L-R Comp. / (St. Dev. / Mean) of S-R Comp.
CZK	2.15	6,217	0.0003
HUF	0.71	16	0.0460
PLN	4.17	1,673	0.0025
SIT	1.34	1,750	0.0008
SKK	1.55	606	0.0026
EUR	1.52	808	0.0019

Source: Authors' estimates.

Table 4. Principal Components of Long-Run Volatility

The Full Period: January 1997-June 2005

	Component 1	Component 2	Component 3	Component 4	Component 5	Component 6
Eigenvalues:	2.61	1.13	0.90	0.72	0.46	0.18
Variance Proportion	44%	19%	15%	12%	8%	3%
Cumulative Proportion	44%	62%	77%	89%	97%	100%
Eigenvectors:						
CZK	-0.17	-0.74	0.19	-0.63	-0.06	-0.04
HUF	-0.48	0.12	-0.31	-0.03	-0.81	-0.06
PLN	-0.27	-0.48	-0.65	0.42	0.32	-0.03
SIT	-0.53	0.30	0.09	-0.18	0.37	-0.67
SKK	-0.31	-0.27	0.66	0.61	-0.14	-0.03
EUR	-0.55	0.22	0.07	-0.16	0.28	0.74

The Early Period: January 1997-June 2001

	Component 1	Component 2	Component 3	Component 4	Component 5	Component 6
Eigenvalues:	3.66	0.97	0.95	0.24	0.12	0.06
Variance Proportion	61%	16%	16%	4%	2%	1%
Cumulative Proportion	61%	77%	93%	97%	99%	100%
Eigenvectors:						
CZK	-0.37	-0.08	-0.69	0.31	0.44	-0.32
HUF	-0.48	-0.07	0.15	-0.66	-0.11	-0.54
PLN	-0.46	-0.13	-0.39	-0.13	-0.51	0.59
SIT	-0.44	0.03	0.41	0.67	-0.36	-0.24
SKK	-0.11	0.99	-0.11	-0.06	-0.03	0.01
EUR	-0.46	0.01	0.42	-0.05	0.64	0.46

The Late Period: July 2001-June 2005

	Component 1	Component 2	Component 3	Component 4	Component 5	Component 6
Eigenvalues:	3.32	1.28	0.64	0.32	0.30	0.14
Variance Proportion	55%	21%	11%	5%	5%	2%
Cumulative Proportion	55%	77%	87%	93%	98%	100%
Eigenvectors:						
CZK	-0.40	-0.17	-0.75	0.23	0.45	0.00
HUF	-0.39	0.48	0.34	0.62	0.08	-0.34
PLN	-0.30	0.68	-0.02	-0.55	0.24	0.30
SIT	-0.36	-0.48	0.50	-0.31	0.50	-0.20
SKK	-0.49	-0.06	-0.20	-0.33	-0.61	-0.48
EUR	-0.49	-0.23	0.18	0.21	-0.33	0.72

Source: Authors' estimates.

Table 5. Correlations of Long-Run Volatility Component

The Full Period: January 1997-June 2005

	CZK	HUF	PLN	SIT	SKK	EUR
CZK	1	0.09	0.21	0.07	0.20	0.13
HUF		1	0.32	0.54	0.20	0.59
PLN			1	0.16	0.14	0.22
SIT				1	0.29	0.82
SKK					1	0.32
EUR						1

The Early Period: January 1997-June 2001

	CZK	HUF	PLN	SIT	SKK	EUR
CZK	1	0.51	0.84	0.36	0.13	0.37
HUF		1	0.78	0.74	0.11	0.85
PLN			1	0.59	0.10	0.60
SIT				1	0.15	0.86
SKK					1	0.14
EUR						1

The Late Period: July 2001-June 2005

	CZK	HUF	PLN	SIT	SKK	EUR
CZK	1	0.30	0.25	0.38	0.64	0.58
HUF		1	0.67	0.24	0.49	0.52
PLN			1	0.01	0.42	0.24
SIT				1	0.51	0.69
SKK					1	0.77
EUR						1

Source: Authors' estimates.

Table 6. Principal Components of Short-Run Volatility

The Full Period: January 1997-June 2005

	Component 1	Component 2	Component 3	Component 4	Component 5	Component 6
Eigenvalues:	1.90	1.34	0.96	0.74	0.70	0.36
Variance Proportion	32%	22%	16%	12%	12%	6%
Cumulative Proportion	32%	54%	70%	82%	94%	100%
Eigenvectors:						
CZK	-0.39	0.11	0.64	-0.52	-0.36	-0.16
HUF	0.09	0.69	0.11	-0.28	0.62	0.22
PLN	-0.03	0.68	-0.01	0.49	-0.54	0.07
SIT	0.35	-0.13	0.74	0.50	0.24	-0.13
SKK	0.60	0.18	-0.14	-0.28	-0.15	-0.70
EUR	0.60	-0.11	0.12	-0.29	-0.34	0.65

The Early Period: January 1997-June 2001

	Component 1	Component 2	Component 3	Component 4	Component 5	Component 6
Eigenvalues:	1.73	1.25	0.98	0.84	0.77	0.43
Variance Proportion	29%	21%	16%	14%	13%	7%
Cumulative Proportion	29%	50%	66%	80%	93%	100%
Eigenvectors:						
CZK	-0.26	0.37	0.64	-0.52	0.30	0.16
HUF	-0.46	0.01	-0.59	-0.40	0.35	-0.40
PLN	-0.10	0.67	-0.19	-0.13	-0.70	-0.07
SIT	-0.12	-0.64	0.16	-0.52	-0.53	-0.06
SKK	0.54	0.05	-0.39	-0.49	0.11	0.55
EUR	-0.63	-0.10	-0.18	0.21	-0.06	0.71

The Late Period: July 2001-June 2005

	Component 1	Component 2	Component 3	Component 4	Component 5	Component 6
Eigenvalues:	2.38	1.16	1.04	0.71	0.47	0.25
Variance Proportion	40%	19%	17%	12%	8%	4%
Cumulative Proportion	40%	59%	76%	88%	96%	100%
Eigenvectors:						
CZK	-0.48	-0.06	-0.36	-0.16	-0.78	0.00
HUF	0.10	-0.69	-0.45	0.55	0.08	0.09
PLN	-0.25	0.63	-0.13	0.72	0.02	0.01
SIT	-0.19	-0.31	0.81	0.35	-0.31	0.03
SKK	-0.57	-0.18	0.00	-0.05	0.38	-0.71
EUR	-0.57	-0.08	0.02	-0.15	0.38	0.70

Source: Authors' estimates.

Table 7. Correlations of Short-Run Volatility Component

The Full Period: January 1997-June 2005

	CZK	HUF	PLN	SIT	SKK	EUR
CZK	1	0.04	0.05	-0.07	-0.33	-0.23
HUF		1	0.28	0.01	0.19	-0.03
PLN			1	-0.06	0.07	-0.09
SIT				1	0.17	0.31
SKK					1	0.57
EUR						1

The Early Period: January 1997-June 2001

	CZK	HUF	PLN	SIT	SKK	EUR
CZK	1	0.07	0.13	-0.04	-0.19	0.07
HUF		1	0.07	0.04	-0.10	0.39
PLN			1	-0.20	-0.01	0.05
SIT				1	-0.06	0.11
SKK					1	-0.46
EUR						1

The Late Period: July 2001-June 2005

	CZK	HUF	PLN	SIT	SKK	EUR
CZK	1	0.00	0.21	0.00	0.54	0.54
HUF		1	-0.22	-0.05	-0.02	-0.11
PLN			1	-0.04	0.19	0.22
SIT				1	0.24	0.21
SKK					1	0.75
EUR						1

Source: Authors' estimates.

Table 8. Wald Tests for Causality in Long-Run Volatility Component

The Full Period: January 1997-June 2005

	Dependent variable:				
	CZK	HUF	PLN	SIT	SKK
Explanatory variables:					
CZK	---	12.93 ***	5.65 **	1.08	0.32
HUF	0.34	---	7.34 ***	2.82 *	3.38
PLN	0.12	3.13 *	---	3.15 *	2.08
SIT	3.41 *	0.48	1.98	---	3.96
SKK	0.10	0.22	0.07	0.01	---
EUR	0.06	1.32	0.93	1.68	2.34

The Early Period: January 1997-June 2001

	Dependent variable:				
	CZK	HUF	PLN	SIT	SKK
Explanatory variables:					
CZK	---	1.12	8.24 ***	0.05	0.15
HUF	0.38	---	6.72 ***	2.85 *	0.18
PLN	0.01	7.93 ***	---	4.89 **	2.68
SIT	0.00	2.15	0.05	---	1.03
SKK	0.08	0.00	0.23	0.01	---
EUR	0.00	2.39	2.83 *	4.59 **	1.22

The Late Period: July 2001-June 2005

	Dependent variable:				
	CZK	HUF	PLN	SIT	SKK
Explanatory variables:					
CZK	---	0.01	2.11	0.01	0.09
HUF	2.52	---	1.85	1.38	3.32 *
PLN	1.30	0.85	---	0.47	2.06
SIT	0.10	0.74	1.20	---	0.73
SKK	0.94	0.01	2.20	0.14	---
EUR	4.67 **	1.22	1.05	0.06	5.00 **

Source: Authors' estimates.

*, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent levels, respectively.

Table 9. Wald Tests for Causality in Short-Run Volatility Component

The Full Period: January 1997-June 2005

	Dependent variable:				
	CZK	HUF	PLN	SIT	SKK
Explanatory variables:					
CZK	---	0.01	2.18	49.51 ***	0.87
HUF	0.00	---	8.95 ***	28.06 ***	1.58
PLN	2.27	1.20	---	0.62	1.55
SIT	0.79	0.84	1.68	---	9.02 ***
SKK	0.32	0.69	7.52 ***	0.95	---
EUR	3.43	0.11	3.69 *	2.06	10.45 ***

The Early Period: January 1997-June 2001

	Dependent variable:				
	CZK	HUF	PLN	SIT	SKK
Explanatory variables:					
CZK	---	0.00	1.54	19.32 ***	0.13
HUF	7.48 ***	---	9.62 ***	34.54 ***	12.59 ***
PLN	0.46	6.79 ***	---	2.01	4.24 **
SIT	0.53	0.23	3.28 *	---	4.67 **
SKK	0.79	0.24	1.75	1.85	---
EUR	3.26 *	6.73 ***	4.13 **	49.48 ***	2.90 *

The Late Period: July 2001-June 2005

	Dependent variable:				
	CZK	HUF	PLN	SIT	SKK
Explanatory variables:					
CZK	---	0.09	1.42	0.50	2.76 *
HUF	1.57	---	4.10 **	0.72	4.39 **
PLN	0.86	0.09	---	0.74	2.28
SIT	3.59 *	2.27	0.74	---	2.19
SKK	2.70	0.58	2.62	2.26	---
EUR	4.78 **	1.88	0.85	144.02 ***	0.97

Source: Authors' estimates.

*, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent levels, respectively.

Table 10. Wald Tests for Causality in Mean

The Full Period: January 1997-June 2005

	Dependent variable:				
	CZK	HUF	PLN	SIT	SKK
Explanatory variables:					
CZK	0.45	55.75 ***	7.24 ***	0.37	0.03
HUF	0.51	7.44 ***	1.25	0.35	1.22
PLN	0.00	2.80 *	3.34 *	0.19	0.19
SIT	0.07	0.02	1.50	1.31	0.00
SKK	0.00	0.23	0.01	8.18 ***	5.55 **
EUR	0.01	2.84 *	2.26	1.01	0.26

The Early Period: January 1997-June 2001

	Dependent variable:				
	CZK	HUF	PLN	SIT	SKK
Explanatory variables:					
CZK	0.02	0.00	2.67	0.77	0.08
HUF	0.11	0.32	3.52 *	0.02	0.00
PLN	0.01	0.51	4.42 **	0.74	2.50
SIT	3.86 **	0.23	0.89	0.58	0.09
SKK	0.03	0.41	0.01	3.40 *	0.34
EUR	0.01	0.27	9.33 ***	3.03 *	0.49

The Late Period: July 2001-June 2005

	Dependent variable:				
	CZK	HUF	PLN	SIT	SKK
Explanatory variables:					
CZK	2.57	0.49	1.43	9.19 ***	0.98
HUF	0.18	0.25	1.01	4.11 **	0.04
PLN	5.89 **	6.66 ***	4.39 **	1.54	4.18 **
SIT	0.68	1.20	2.71	1.25	0.03
SKK	557.00 ***	3.61 *	2.47	12.70 ***	3.37 *
EUR	2.46	2.76 *	1.36	0.14	1.66

Source: Authors' estimates.
*, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent levels, respectively.