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Some Exchange Rates Are More Stable than Others: Short-Run Evidence from Transition Countries

Aleš Bulíř*

Abstract

The paper investigates empirically the endogenous liquidity nexus of exchange rate determination on a sample of four transition economies. We find evidence in favor of the hypothesis of a nonlinear error correction process vis-à-vis longer-term trend deviations. The results suggest that early and successful exchange-rate market and financial-account liberalization pays off in terms of depth of the market and, hence, faster adjustment of national currencies to short-term shocks to the exchange rate.

JEL Codes: F31, F33, C32.
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Nontechnical Summary

Economists have long observed that exchange rates are more volatile than would seem to be justified by conventional macroeconomic models (Meese and Rogoff, 1983) and that the models' short-term explanatory power is severely limited as compared to models that incorporate elements of microstructure finance (Evans and Lyons, 2002 and Derviz, 2003). In addition, exchange rate developments appear to follow nonlinear trajectories (Sarno, 2003) and nonlinear models in general mimic actual exchange rate developments better than those based on linear assumptions.

This paper tested a nonlinear model of exchange rate determination, with endogenously provided liquidity, on daily and weekly exchange rate data for the U.S. dollar and euro against four Central European currencies for the 1998–2002 period: the Czech and Slovak korunas, Hungarian forint, and Polish zloty. The investigation was done in three steps. First, we found indications of endogenous liquidity in exchange rate returns. Second, we identified transitory movements and trends in the exchange rate data. Finally, we tested whether the speed-of-adjustment of national exchange rates is related to longer-term trend deviations thereof. The underlying hypothesis is that foreign exchange order flow is attracted to a liquid market with a seemingly temporary shock to the exchange rate, reacting disproportionately fast to sizeable deviations from the longer-term exchange rate path as compared to small deviations. Our results are broadly comparable to those for G-7 countries reported Faruqee and Redding (1999).

The results have some attractive policy implications. First, early liberalization of the foreign exchange market and financial-account transactions may pay off in terms of market liquidity and, hence, faster adjustment of the exchange rate to the longer-term trend. However, early liberalization is a necessary condition for liquidity, not a sufficient one, as shown in the paper by some currencies. Second, the endogenous liquidity hypothesis does not imply that the volatility of the nominal exchange rate is necessarily low in a liquid market, we only observe that the self-correcting mechanism is faster than in an illiquid market. The endogenous liquidity hypothesis seems to hold especially well in the full sample of the Czech koruna and shorter samples of the Hungarian forint.

1. Introduction

Economies with liberalized exchange-rate markets and financial-account transactions appear to have a faster speed of adjustment of their national currency exchange rates to trend-based equilibrium values as compared to those currencies that remain regulated. These developments in Central European countries seem to be linked to endogenously provided liquidity. The speed of adjustment appears to be significantly faster vis-à-vis the euro as opposed to the dollar, an intuitive consequence of the region's link to the euro area. In addition, these results seem to be in line with the finding that volatility of exchange rates in the region appears to be only loosely related to the national exchange rate arrangements (Darvas and Szapáry, 2000; Wickham, 2002).

Models of foreign exchange markets with endogenously provided liquidity assume that large deviations from longer-term pricing levels attract marginal liquidity into the market, which in turn speeds up the market's return toward its trend-based value (Sarno, 2003). Hence, markets with a large number of informational traders or with deep order books, ought to have smaller and shorter deviations from the trend of the exchange rate. The fact that endogenous liquidity (the client order flow) can affect market exchange rates is well known, and Faruqee and Redding (1999); Evans and Lyons (2002); Bacchetta and van Wincoop (2003); and Derviz (2003) present useful theoretical models of endogenous liquidity and foreign exchange microstructure that capture this impact. Empirically, the liquidity effect has been shown by Faruqee and Redding (1999) on a sample of G7 currencies, and the order flow effect by Evans and Lyons (2002).

In this paper, we tested the liquidity impact using daily and weekly exchange rate data for the U.S. dollar and euro against four Central European currencies for the 1998–2002 period: the Czech and Slovak korunas, Hungarian forint, and Polish zloty. First, we examined the statistical properties of the series and found excess kurtosis present in the exchange rate returns. Second, we identified transitory movements and trends in the data using the Hodrick–Prescott filter. Finally, we tested the speed-of-adjustment hypothesis within an error-correction framework, finding results comparable to those of Faruqee and Redding (1999).

The effects of endogenous liquidity seem to dominate the effects of the random walk in exchange rate determination. The results also suggest that early foreign-exchange market and financial-account liberalization might explain the resulting, faster-than-average adjustment of some national currencies to short-term deviations from the trend. Liberalized markets with sufficient endogenous liquidity, such as the Czech Republic, are likely to be more stable than those without endogenous liquidity, such as Slovakia, or tightly regulated ones, such as Hungary in the late 1990s.

The paper is organized as follows. After discussing the model and stylized facts, we examine the data properties and outline the empirical techniques. In the following section we present our results. The final section concludes.

2. A Model of Short-Term Exchange Rate Determination

Macroeconomists have long observed that exchange rates—both in nominal and real terms—are more volatile than would seem to be justified by conventional macroeconomic models and that the models' short-term explanatory power is limited (Meese and Rogoff, 1983). Clearly, exchange rates can remain fairly long outside of the "fair" value, be it based on fundamental- or equilibrium-based calculations. In contrast, models augmented with microeconomic variables—such as the order flow of foreign exchange in the field of microstructure finance—seem to do much better in the short term.[1]

2.1 Foreign Exchange Markets in the Short Run

In reality, foreign exchange flows and price developments are linked endogenously. Only a fraction of foreign exchange trading is related to a desire to take a particular asset position, for the purposes of either financial investment or foreign trade activity, and the remainder is initiated by some traders providing liquidity to other traders for an expected profit. The expected profit may be based on a sudden departure from a longer-term level of the exchange rate or an increase in the bid-ask spread. Intuitively, large and sudden deviations from the longer-term level of the exchange rate would either activate limit orders set by the existing traders or attract new traders into the market, accelerating the exchange rate's return back to its longer-term level.

Market participants are likely to place large limit orders only in deep and transparent markets.[2] Taking a clue from the theory of finance: "investors tend to invest in instruments they are aware of" (Merton, 1987). On the one hand, knowing that the market for a given currency is deep, that is, limit orders are placed on exchange rate departures in either direction, agents would presumably be less prone to initiate sharp price adjustment to non-fundamental news. On the other hand, the return to longer-term equilibrium values ought to be faster with a deep market, because those limit orders would be activated or new market makers would enter such a market quickly to take advantage of the profit opportunity.[3] In reality, the former effect is likely to outstrip the latter one: the number of foreign exchange dealers in transition countries has been low and stable in the

[1] Order flow is defined as the net of buyer-initiated and seller-initiated orders. In conventional macroeconomic models actual trades are typically not related to exchange rate movements. Empirical examples of the link between order flow and daily exchange rate returns were documented for the U.S. dollar/German mark and Czech koruna/euro rates by Evans and Lyons (2002) and Derviz (2003), respectively. Unfortunately, the existing order flow data have been collected in an *ad hoc* fashion and are yet to be tested in a long-run sample. For example, the Evans and Lyons paper uses a four-month period in 1996 and Derviz's data run from mid-1999 until end-2001.

[2] For simplicity, we omit the role of quantitative central bank interventions or informal rules and regulatory pressures in this exercise. Regarding the former, the measurable impact of quantitative interventions is hard to gauge as it partly depends on the market structure, see Bofinger and Wollmershaeuser (2001) and Holub (2003). Even in the Czech Republic, where central bank interventions have been sizeable from time to time, the typical monthly intervention has been substantially less than the daily amount of regular foreign exchange market trading. Regarding the latter, a rule of central bank "sterilization" of privatization proceeds into its reserves is an example of nonstandard intervention with a potentially sizable impact. Even less can be said about regulatory pressures, although it is known that some foreign exchange market participants may "restrain" themselves voluntarily if they are worried that the central bank might "punish" them for speculating against the currency.

[3] The empirical literature provides some support for this hypothesis. For example, Devereux and Lane (2003) found that high levels of financial linkages between creditor and developing countries resulted in a lower level of bilateral exchange rate variability.

short run. Controlling for exchange rate regimes, foreign exchange market history, central banks' interventions, both open and hidden, and so on, deep markets ought to exhibit faster adjustment to longer-term levels of exchange rates.

2.2 An Outline of a Theoretical Model

How would such an economy operate? Faruqee and Redding (1999) and Evans and Lyons (2002) employed three-period, short-term models that share some salient features and we will use a simplified version thereof to motivate the empirical work in our paper. Arguably, the Evans and Lyons model is geared toward explaining intra-day movements in the exchange rate, while Faruqee and Redding are concerned with somewhat longer periods. However, the adjustment in both models is effected through similar processes.

Consider a three-period economy in which trading influences the price of an asset (foreign currency). Ignoring the difference between the direct and brokered markets, the economy is populated by two types of agents. First, non-informational ("noise") traders who trade primarily for purposes of financial investment or foreign trade activity and may push the exchange rate away from the best estimate of the liquidation price (${}_iEP_3$) in each period. These traders do not have a longer-term view of the exchange rate and typically buy foreign exchange at a given price. Second, informational traders (sometimes also called market makers), who provide desired liquidity in exchange for an expected profit to be realized in period 3 when the market is liquidated. These traders speculate, with a view of returning the rate to a market clearing level.[4]

The asset will have a liquidation price (P_3) in the final period that is not known with certainty prior to T_3. The asset is being traded in the first two periods, T_1 and T_2, resulting in prices P_1 and P_2, gradually improving the precision of the estimate of P_3, i.e., ${}_1EP_3$ and ${}_2EP_3$. The first- and second-period estimates of P_3 are normally distributed around P_3 with posterior sample variances of σ_1^2 and σ_2^2, respectively, such that $\sigma_1^2 \geq \sigma_2^2$. Naturally, if non-informational traders' orders were completely absent, the second-period price would be equal to the expectation of the third-period price. The resulting time line of exchange rate determination is described in Figure 1.

Figure 1: Time Line of Exchange Rate Determination

Period 1 (T_1)				**Period 2 (T_2)**			**Period 3 (T_3)**
First-period informational traders enter the market and limit orders are set	The net demand of "noise" traders is announced	${}_1EP_3$ is revealed	Public trading results in a rate P_1	Second-period informational traders enter the market and limit orders are set	${}_2EP_3$ is revealed	Public trading results in a rate P_2	The market is liquidated at a price P_3

[4] Which particular instruments are used by individual market participants is a side issue, depending primarily on the structure of individual markets.

The model assumption of gradual price discovery ($\sigma_1^2 \geq \sigma_2^2$) implies that the expected exchange rate will drift toward its longer-term value. However, the speed of the adjustment is endogenous: the bigger the absolute value of the discrepancy between the expected final and realized first-period exchange rates $\left(\left|{}_1EP_3 - P_1\right|\right)$, the deeper the order book will be in the second period and the faster the reversion to P_3.

2.3 An Error-Correction Framework

The principal empirical implication of the Faruqee–Redding model is that the speed of adjustment of exchange rates toward their fundamental values is proportionately higher when the deviation from the trend equilibrium value level is large and, hence, the order book is deep. While the former implication is testable, the latter is not, given the absence of long-run order flow data.

The empirical testing of exchange rate developments (e) consists of two parts. First, we separate longer-term trends (e^f), which are presumably affected by fundamental factors, from the effects of transitory demand (e^n):

$$e = e^f + e^n. \tag{1}$$

The decomposition was done using univariate techniques, namely the Hodrick–Prescott and kernel filters.[5] The filters were used to characterize the trend portion of the exchange rate, and the differences of the filter from the actual exchange rate were used to characterize the transitory demand. The pros and cons of using the above univariate techniques are obvious. On the one hand, we avoid the troublesome joint testing of the endogenous liquidity hypothesis and an *ad hoc* chosen model of long-term exchange rate determination as in Meese and Rogoff (1983). On the other hand, filtering away short-term fluctuations is not equivalent to the fundamental-based exchange rate. Moreover, filtering does not shed light on the nature of the trend process. The underlying trend can be stochastic, supporting an I(1) variable, or deterministic with broken trend and, hence, supporting an I(0) variable. Applying the selected filters we impose a fixed speed of adjustment of the trend to actual developments in all periods and all countries.[6] On balance, however, we see the disadvantages of joint testing as more serious than those of mechanical filtering.

Second, we explore the nature of the short-term fluctuations. The model predicts that the transitory departure from the trend ought to be short-lived and the exchange rate would revert in expectation to the trend value of the exchange rate:

$$\left| {}_tE\left[e_{t+1}^n\left(e_t^n\right)\right] \right| < \left|e_t^n\right|, \tag{2}$$

[5] Only the results using the Hodrick–Prescott filter are shown. We use the common rule for determining the bandwidth parameter λ for the Hodrick–Prescott filter: 100 times its frequency-squared. Hence, λ takes the value 14,400, 260,100, and 6,250,000 for monthly, weekly, and daily data, respectively. The results using the kernel-based smoother (the Epanechnikov kernel) are not materially different from the Hodrick–Prescott filter, including spurious autocorrelation (Cogley and Nason, 1995).

[6] For example, the market may be characterized by two or more adjustment processes: one through private agent interactions and the other through central bank interventions. Presumably, the monetary authority is concerned less with the speed of adjustment of the nominal exchange rate than with finding a particular level of the real exchange rate.

where the expected value of the transitory component of the exchange rate in time t+1 is conditional on its value in time t.

Moreover, the model predicts the nonlinear mean reversion to be positively related to the initial value of the short-term deviation from the trend (e^n):

$$\frac{d\left[e_t^n - {}_tE\left(e_{t+1}^n\right)\right]}{de_t^n} > 0 \tag{3}$$

In other words, the error correction term in the numerator ought to be increasing in the initial departure from the trend.[7]

The presence of the above univariate error-correction mechanism can be tested in a simple equation:

$$\Delta e_t = \alpha e_{t-1}^n + \beta\left(e_{t-1}^n\right)^3 + \sum_1^x \delta_i \Delta e_{t-i} + \varepsilon_t , \tag{4}$$

where $e_t^n = e_t - e_t^J$ is the difference from the Hodrick–Prescott filter; the parameter α captures the proportional part of the mean reversion (error-correction) process; and the parameter β captures the additional mean reversion owing to the endogenous liquidity hypothesis. The advantage of a cubic term—as opposed to a quadratic term—for the additional mean reversion process is that it preserves the sign of the transitory component. In addition, equation 4 can include lagged dependent variables and, hence, it becomes a direct test of the random walk hypothesis of exchange rate determination as compared to the endogenous liquidity hypothesis.

Given that mean reversion entails negative autocorrelation, both the proportional and additional error-correction mechanisms imply negative expected values for α and β. Should the endogenous liquidity hypothesis be rejected, the estimated value of β would have to be zero. Hence, in the subsequent sections we will test both the sign and significance of the α and β coefficients for different country series and time periods. The null hypothesis of the random walk implies that δ_i is equal to zero.

3. Stylized Facts About Foreign Exchange Markets

3.1 The Liberalization Process in Central Europe and Market Liquidity

Although the national exchange rate arrangements differed in the initial stages of transition, the authorities in the Czech Republic, Hungary, Poland, and the Slovak Republic (also known as the Viszegrád Four) eventually adopted managed floating arrangements, gradually building up liquidity in their foreign exchange markets (Table 1).[8] Notwithstanding the similarity in the choice of the arrangement during 2000–2002, the volume of trading on those markets differed

[7] Granger and Siklos (1997) addressed a similar problem, namely that the cointegrating relationship can be nonlinear in nature and, moreover, depend on the nature of the policy regime in place. See also the review of nonlinear adjustment models in Sarno (2003).

[8] This classification is based on self-assessment by individual countries. Reinhart and Rogoff (2002) or Bofinger and Wollmershaeuser (2001) have shown that the actual behavior of exchange rates can be consistent with regimes different from those declared officially.

substantially (Figure 2).[9,10] This is certainly related to the history of the market: for example, in the mid-1990s the Czech market was exposed to large privatization-related foreign exchange trading with substantial exchange rate risk, necessitating early development of the derivative market, as compared to the less exposed market in Slovakia or the tightly controlled Hungarian market. In terms of market size, the Czech koruna was the most heavily traded, while the market for the Slovak koruna was the thinnest. The Polish market was large in levels, but in per capita terms it was comparable to the Slovak market.[11]

Our sample countries differed also in terms of capital-flow liberalization. In terms of institutional control, Czech financial-account transactions were freed in the second half of the 1990s and the koruna was floated subsequently. These developments resulted in large daily turnovers and fast-growing derivative markets. Slovakia followed a similar path, although with much slower turnover growth. In contrast, Hungary and Poland maintained narrow-band crawling peg regimes and tight capital-flow controls until the fall of 2001, with correspondingly small turnover. Another measure of institutional liberalization and market depth is the volume of the eurobond market in the national currency. While the Czech koruna-denominated market has had around 40 new issues per year for the past six years (with outstanding eurobonds of more than U.S. dollar 3 billion), forint-denominated eurobonds were first issued in 2001 (with a current outstanding volume of about U.S. dollar 100 million). No Slovak-koruna denominated eurobonds have been issued yet, showing Slovakia's comparatively slow progress in foreign exchange market development, and no summary data are available for Poland.

Unfortunately, none of the above indicators provide an explicit measure of foreign exchange market liquidity. For example, trading volumes are only imperfect substitutes, dependable on the chosen structure of the market: *ceteris paribus*, a market with multiple market makers is going to record much higher trading volume than a broker-based market (Bofinger and Wollmershaeuser, 2001). Hence, high trading volumes in mid-1998 in the Czech Republic or end-2002 in Hungary may indicate periods of market turmoil rather than overall high liquidity. Similarly, the number of tradable instruments is only an indication of potential liquidity and, for example, the volumes on the domestic-currency eurobond markets have been rather thin.

The available data provide an indication of market liquidity, not a precise measure thereof. Nevertheless, one expects the Czech market to be the most liquid from the outset, irrespective of the subsequent decline in turnover. The remaining markets are less liquid, although the Hungarian market has picked up substantially following the floating of the forint. In terms of the empirical model defined above, the estimates of the nonlinear speed of adjustment, β, ought to be high in absolute terms for the Czech koruna for the early data and lower in the later data, possibly accompanied by increasing absolute value αs. In the remaining countries, the estimated absolute-value βs might be increasing in time as their currencies were floated.

[9] Both the daily returns and the deviations from the Hodrick–Prescott filter exhibit a fair amount of spatial correlation (as expected, the dollar rate pair-wise correlations are larger than those of the euro-rate). This confirms the intuitive assumption that all Central European currencies were subject to similar exchange rate shocks.

[10] See Darvas and Szapáry (2000) for an argument of exchange rate arrangement irrelevancy for exchange rate volatility.

[11] Poland's population in the late 1990s was almost 40 million, as compared to about 10 million for the Czech Republic and Hungary and 5 million for Slovakia. See also some relative comparisons at Bofinger and Wollmershaeuser (2001).

Table 1: Exchange Rate Regimes and Capital Account Developments in Selected Countries, 1991–2002[1]

Country	Period	Regime	Band	Periodic de-valuations	Capital account developments
Czech Republic	January 1993–May 1997	Peg	+/- 0.5 percent, changed to +/- 7.5 percent in February 1996	No	Fast deregulation from the mid 1990s, very liberal regime (1999 Foreign Exchange Act)
	From May 27, 1997	Float (with discretionary interventions)	No	…	
Hungary[2]	January 1991–February 1995	Soft peg	Gradually changed from nil to +/- 2.25 percent in December 1994	Yes	Closely regulated until June 2001 Government decree
	March 1995–September 2001	Crawling peg (pre-announced)	+/- 2.25 percent, change to +/- 15 percent in May 2001	Yes	
	From October 1, 2001	Float (with discretionary interventions)	+/- 15 percent	Yes	
Poland[3]	October 1991–March 2000	Crawling peg (pre-announced)	Gradually changed from nil to +/- 15 percent in December 1994	Yes	Slow progress until the late 1990s, major liberalization in January 2000
	From April 12, 2000	Independent float	No	…	
Slovak Republic	January 1993–September 1998	Peg	Extended in several steps from +/- 1.5 percent in January 1996 to +/- 7 percent in January 1997	No	Similar steps if somewhat slower pace than in the Czech Republic (1999 Amendment to Foreign Exchange Act)
	From October 1, 1998	Float (with discretionary interventions)	No	…	

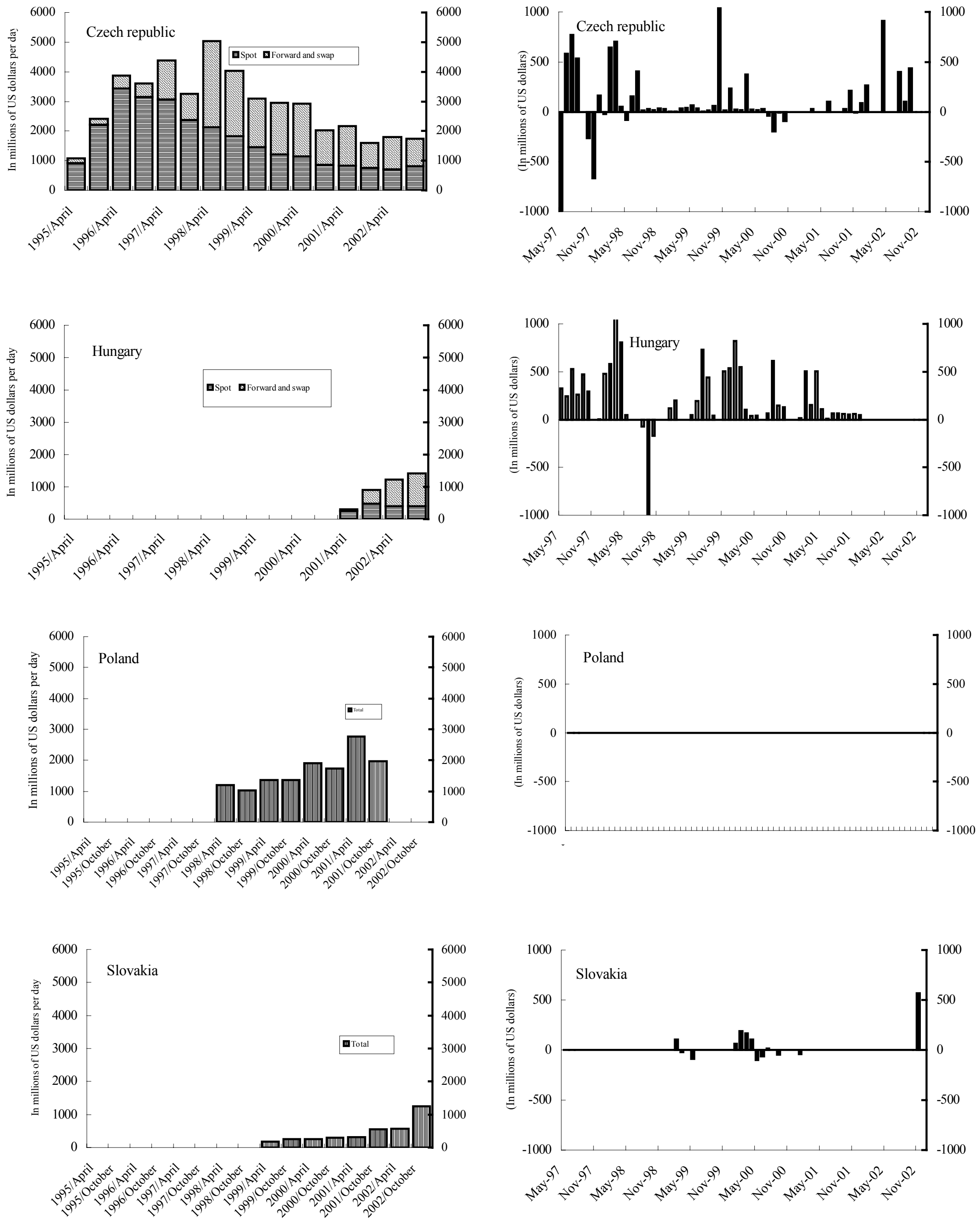
Sources: Darvas and Szapáry (2000); Schardax (2002); and central banks' official web sites.

Notes: [1)] This table contains an official classification of the national exchange rate arrangement as declared by the national monetary authorities. However, Reinhart and Rogoff (2002) have shown that *de jure* as opposed to *de facto* exchange rate arrangements can differ. For example, the early Czech, Hungarian, and Slovak pegs are re-classified as crawls, Poland's crawling peg during 1991–95 is re-classified as a dual market rate arrangement, and so on.

[2)] Step devaluation (9 percent) in March 1995. Schardax (2002) argued that the widening of the intervention band, followed in October 2001 by the adoption of an inflation targeting regime, is equivalent to a managed floating regime.

[3)] Step devaluation (6 percent) in December 1995.

Figure 2: The Czech Republic, Hungary, Poland[1], and Slovakia: Average Daily Foreign Exchange Trading Data[2] and Monthly Foreign Exchange Interventions[3], Various Dates



Sources: National central bank web sites.

Notes: [1] No interventions during 2000-2002; older data are not publicly available.

[2] Left column.

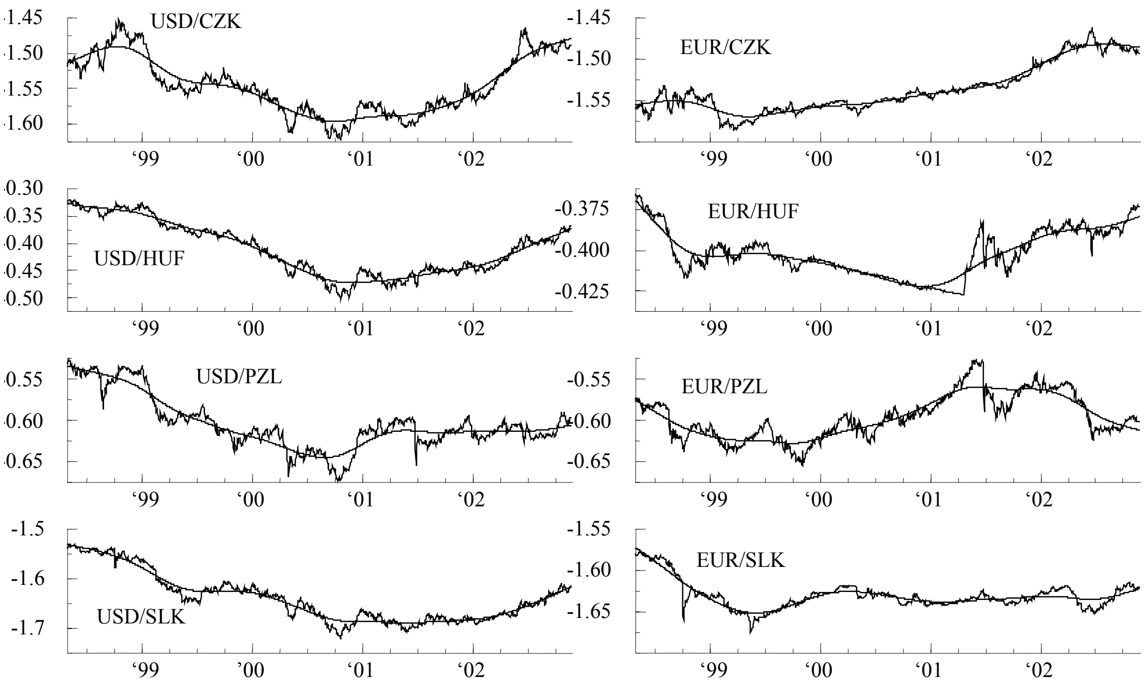
[3] Right column.

3.2 National Exchange Rates Developments

End-period daily, weekly, and monthly exchange rate data from May 1998 until December 2002 were collected. The data are expressed in foreign-currency units (U.S. dollar and euro) per unit of domestic currency and logarithms were taken of all series. All the series were found to be nonstationary, that is, of order I(1). The estimated coefficient of the lagged variable in the augmented Dickey–Fuller (ADF) regressions was typically around 0.99, indicating that the series follows a random walk. The only exception was the weekly data for the Slovak koruna, which appear to be stationary, although the estimated coefficient of the lagged variable remained high at around 0.90–0.91.

We observed that the nominal volatility in our sample does not seem to be linked directly to the exchange rate arrangement in place (Figure 3). For example, the zloty/euro deviations from the trend under the crawling peg arrangement appear to be much larger than Czech koruna/euro deviations from the trend under the floating arrangement. Indeed, the only period of markedly low volatility in our sample was the one-year Hungarian experiment with a tight crawling peg from about December 1999 until November 2000. Overall, the variation of standard deviations of the filtered series are of similar order as the variation of the corresponding standard deviations of the differenced series (for example, the full-sample estimated standard deviations of the daily and weekly euro rates for the Czech and Slovak korunas and the Hungarian forint are practically identical).

Figure 3: The Czech Republic, Hungary, Poland, and Slovakia: Daily U.S. Dollar and Euro Exchange Rates and the Hodrick–Prescott Filter, May 1998–December 2002[1]



Source: Czech National Bank; author’s calculations.
Notes: [1] The Hodrick–Prescott filter parameter λ has been set to 6,250,000. All data are in logs.

The dependent variable for the error-correction mechanism suggested above is the differenced series, and its properties are shown in Table 2.[12] All differenced series were found to be stationary, that is, of order I(0), and the estimated coefficients of the lagged variable in the ADF test were invariantly smaller than 0.20. A few observations are worth noting. First, the volatility of the euro exchange rates is typically lower than that of dollar rates, as most foreign exchange trading has been done in the German mark and subsequently in euro, an intuitive consequence of the region's trading and investment link to Western Europe. As a result, the exchange rate vis-à-vis the dollar is often little more than a cross rate. The Polish zloty is an outlier—the estimated volatility vis-à-vis the dollar and euro are essentially identical.

Table 2: Dollar and Euro Exchange Rate Daily Returns: Sample Moments and Kurtosis tests, May 1998–December 2002

	Czech Republic		**Hungary**		**Poland**		**Slovakia**	
	Dollar	**Euro**	**Dollar**	**Euro**	**Dollar**	**Euro**	**Dollar**	**Euro**
Daily data *(n=1,165)*								
Mean	2.548E-5	5.738E-5	-3.837E-5	-6.472E-5	-5.570E-5	-2.381E-5	-6.586E-5	-3.397E-5
Std. Deviation	0.00334	0.00202	0.00290	0.00159	0.00315	0.00349	0.00317	0.00174
Skewness[1]	0.0585	-0.3171	0.0185	-1.3463	-0.5352	-0.6227	-0.3409	-5.1425
Excess kurtosis[2]	1.3052	5.3962	1.7769	18.792	10.359	8.7817	4.5902	88.746
Normality test[3]	61.99***	503.9***	103.4***	1514.8***	1514.4***	841.8***	393.2***	1066.2***
Weekly data *(n=234)*								
Mean	8.710E-5	2.843E-4	-2.235E-4	-2.638E-5	-3.148E-4	-1.118E-5	3.609E-4	-1.638E-4
Std. Deviation	0.00798	0.00432	0.00658	0.00320	0.00649	0.00647	0.00679	0.00418
Skewness[1]	0.1389	-0.9793	0.2010	-1.0245	-0.9177	-1.2818	0.1892	-2.2137
Excess kurtosis[2]	3.6668	6.0544	0.00561	8.3269	2.9452	3.9448	0.00260	15.786
Normality test[3]	73.65***	78.81***	1.642	129.1***	29.94***	43.83***	1.460	99.25***
Monthly data *(n=56)*								
Mean	5.301E-4	1.193E-3	-7.982E-4	1.351E-4	-1.159E-3	4.950E-4	-1.370E-3	-7.066E-4
Std. Deviation	0.01644	0.00947	0.01121	0.00661	0.01456	0.01383	0.01127	0.00694
Skewness[1]	0.3886	-0.4582	0.3912	0.6595	-0.4535	-1.2049	0.0888	-0.3084
Excess kurtosis[2]	0.8399	2.2613	0.11842	3.2121	1.90459	1.47035	-0.34811	-0.10951
Normality test[3]	4.417*	13.54***	1.748	18.77***	10.751**	16.49***	0.091	1.05

Source: Czech National Bank; author's calculations.

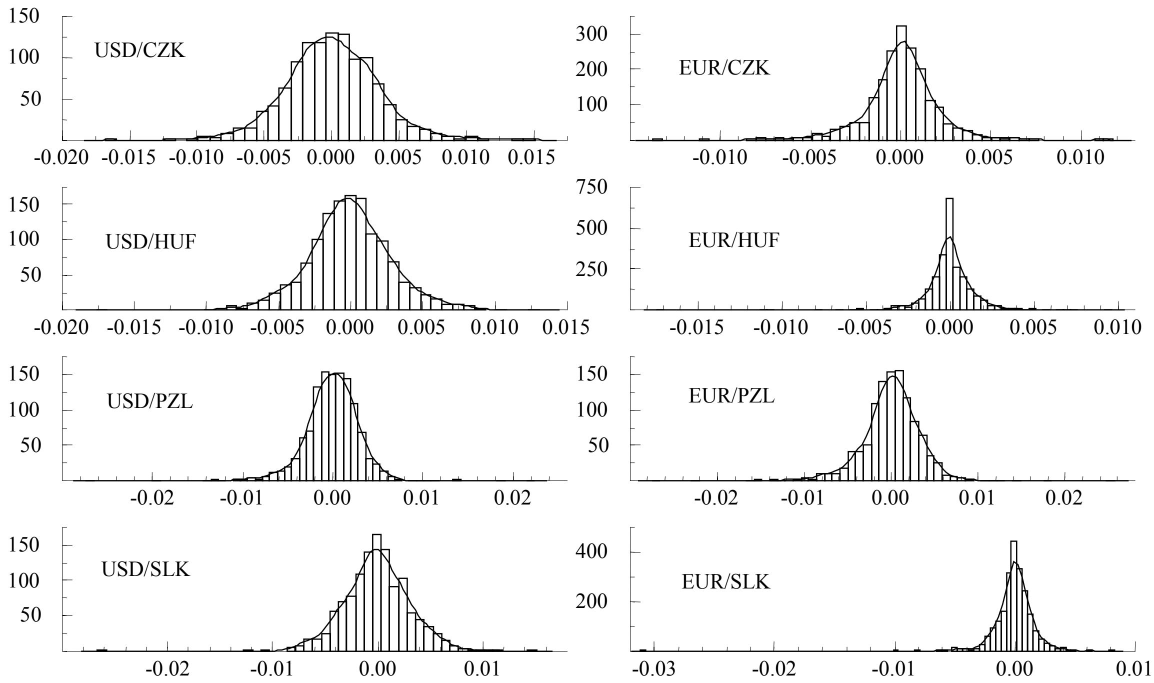
Notes: [1] The skewness statistic θ_1 of a variable x is calculated as $\theta_1 = \frac{E[(x-\mu)^3]}{(Var[x])^{3/2}}$, where μ is the sample mean and the standard deviation is based on 1/T. In normal distributions with no skewness the θ_1 statistic would be zero.

[2] The excess kurtosis statistic θ_2 of a variable x is calculated as $\theta_2 = \frac{E[(x-\mu)^4]}{(Var[x])^2} - 3$, where μ is the sample mean and the standard deviation is based on 1/T. In normal distributions with no excess kurtosis the θ_2 statistic would be zero.

[3] The normality test is distributed as $\chi^2(2)$. The significance level of the rejection of the hypothesis that the variable under consideration is distributed as a normal variable at 1, 5, and 10 percent are denoted by '***', '**', and '*', respectively.

[12] We confirmed the presence of cointegrating relationships between the raw exchange rate and its Hodrick–Prescott filter using the Johansen and Juselius (1990) technique and observed that the speed of adjustment is fairly slow. Both results are intuitive: first, the Hodrick–Prescott filter is a dynamic measure of central tendency and, second, the actual exchange rates remained persistently above or blow its trend.

Figure 4: The Czech Republic, Hungary, Poland, and Slovakia: Estimated Density and Histograms of Daily U.S. Dollar and Euro Exchange Rate Returns, May 1998–December 2002[1]



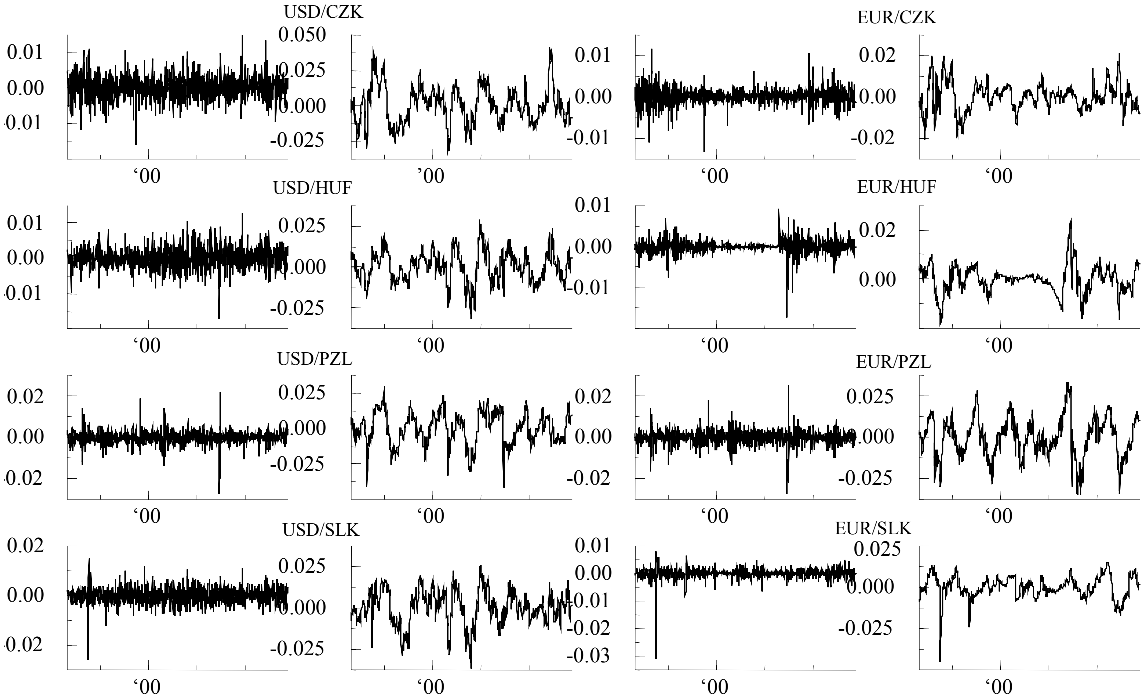
Source: Czech National Bank; author's calculations.
Note: [1] All data are in logs and in first differences.

Second, all daily and most weekly and monthly series fail the normality test at the 1 percent significance level, primarily on account of excess kurtosis present in the data.[13] The distributions of the differenced series have "fat tails," that is, we observe a number of surprisingly high daily returns (Figure 4). See Wickham (2002) for a discussion of the statistical properties of exchange rate series in both industrial and emerging market economies. In reality, skewness and excess kurtosis are typical properties of most high-frequency financial data.

We also observe that throughout our sample the national currencies experienced protracted departures from the trend (Figure 5). Typically, each currency's exchange rate had 20 or more periods when it remained above or below the trend (Hodrick–Prescott filter) for at least 5 trading days, consistent with a slow speed of adjustment in the cointegrating relationship that was observed between the raw series and its Hodrick–Prescott filter. Upon observation, quite a few of those events appear to be consistent with the hypothesis of endogenous liquidity provision—a gradual buildup of deviations from the trend followed, from a certain moment, with a swift adjustment back to the trend.

[13] The outliers are the weekly dollar series for Hungary and Slovakia and also the monthly euro series for Slovakia. As discussed earlier, however, the dollar returns are mostly a product of cross-rate developments. In addition, we have only 234 and 56 observations for the weekly and monthly series, respectively.

Figure 5: The Czech Republic, Hungary, Poland, and Slovakia: Daily U.S. Dollar and Euro Exchange Rate Returns and Deviations from the Hodrick–Prescott Filter, May 1998–December 2002[1]



Source: Czech National Bank; author's calculations.

Note: [1] First differences and differences from the Hodrick–Prescott filter. All data are in logs.

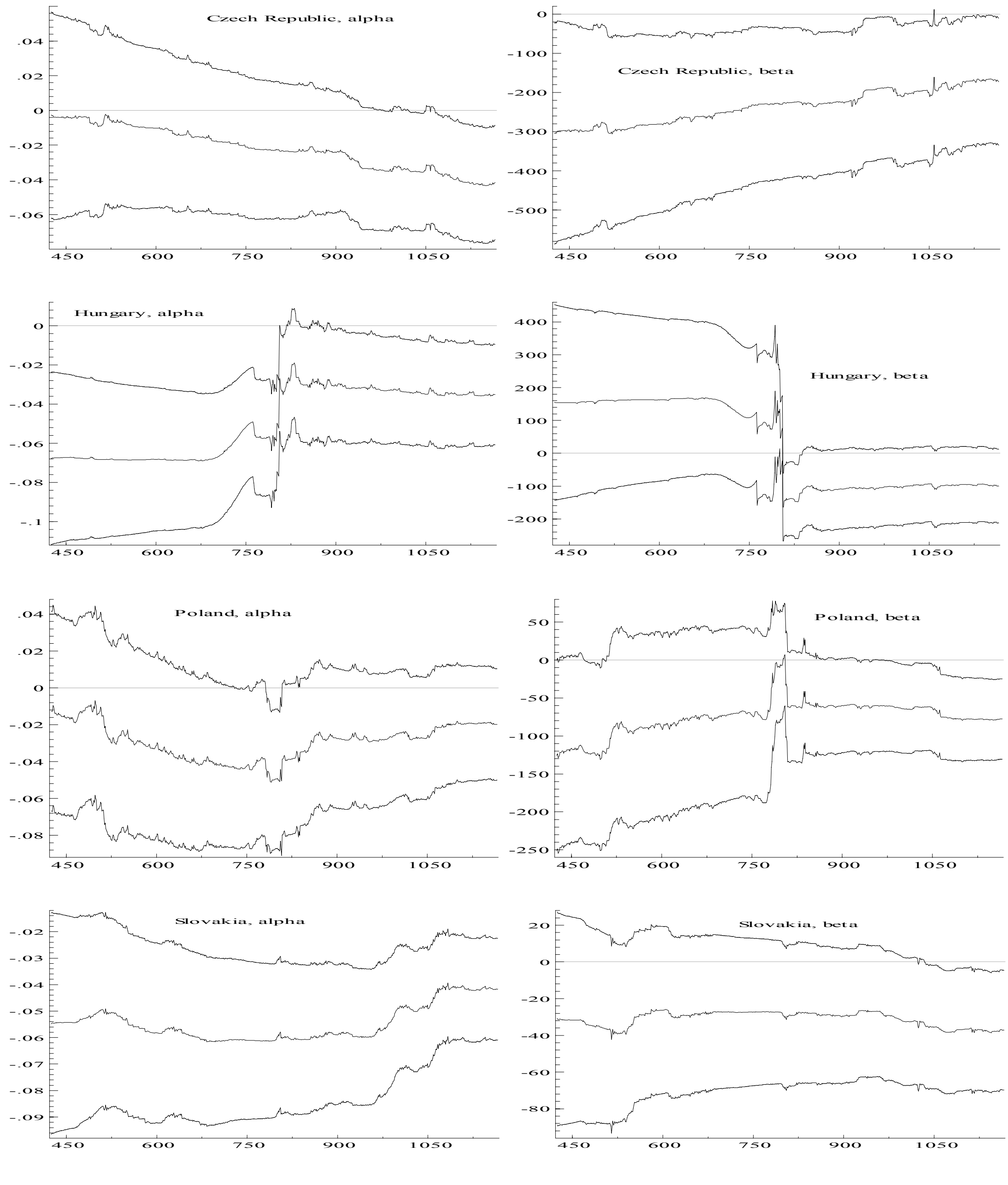
4. Results

The results of estimating equation (4) by ordinary least squares for both the U.S. dollar and euro exchange rates in each of the four countries (the Czech Republic, Hungary, Poland, and Slovakia) are shown in Tables 3–6. In addition to presenting the results for the daily and weekly series in Tables 3–4 and 5–6, respectively, we decided to take into account the different exchange rate regimes. Equation 4 is thus estimated for three sample periods: (i) May 1998–December 2002, corresponding to the period of the float of the Czech koruna; (ii) May 2000–December 2002, corresponding to the period of the float of the Polish zloty and Slovak koruna; and (iii) October 2001–December 2002, corresponding to the period of the float of the Hungarian forint.

4.1 Daily Data

The results for the daily data are shown in Table 3. As is well known, the Hodrick–Prescott filter can create spurious serial correlation in the detrended data (Cogley and Nason, 1995). Indeed, a few of the Durbin–Watson statistics are low, typically in tandem with the Pagan's error autocorrelation test. Also, the usual tests of heteroskedasticity of residuals are uncomfortably high. Therefore, we present the parameter estimates with standard errors corrected for both autocorrelation and heteroskedasticity (HACSE).[14] The stability of estimated parameters of interest—α and β—was tested in recursive regressions and the results for the euro rates are shown in Figure 6. The fact that these regressions explain only a limited portion of the variance of the Δe_t variable (1–7 percent) should not be seen as surprising—the order flow drives daily exchange rate fluctuations, not endogenous liquidity processes.

Figure 6: Recursive estimates of* α *and* β *and their two standard errors



Source: Own estimates of equation (4); daily data, initialization after 400 observations.

[14] As shown by Faruqee and Redding (1999), Monte Carlo experiments with calibrated data replicating the fourth moment of the distribution of dependent variables yield fairly similar α and β coefficients as compared to the ordinary least squares regressions.

Table 3: Daily Data Without Lagged Dependent Variables[1]

	Czech Republic		Hungary		Poland		Slovakia	
	Dollar	**Euro**	**Dollar**	**Euro**	**Dollar**	**Euro**	**Dollar**	**Euro**
Sample period : May 1998–December 2002 (n=1141)								
α	-0.01456	-0.02652	-0.03098**	-0.02643**	0.02559	-0.00258	-0.03981***	-0.02996***
HACSE	(0.0125)	(0.0175)	(0.0148)	(0.0197)	(0.0165)	(0.0169)	(0.0155)	(0.0097)
β	-32.398*	-134.83*	-60.062	-82.542	-176.82***	-85.676**	-16.872	-10.188
HACSE	(16.79)	(69.57)	(45.65)	(161.8)	(47.11)	(43.52)	(36.41)	(25.51)
DW	1.93	2.03	1.93	2.00	1.89	2.04	1.89	1.55
Autocor.	2.0061*	2.2105*	1.5362	3.6337***	1.8494*	1.8948*	2.0595*	11.607***
R^2	0.018	0.027	0.025	0.021	0.061	0.029	0.023	0.017
Sample period : May 2000–December 2002 (n=654)								
α	-0.03225**	-0.05655**	-0.02980	-0.01632	0.02215	0.00375	-0.04837*	-0.01693
HACSE	(0.0163)	(0.0228)	(0.0198)	(0.0252)	(0.0223)	(0.0215)	(0.0202)	(0.0145)
β	-9.357	-17.680	-61.777	-141.46	-177.93**	-84.406*	4.6477	-51.647
HACSE	(17.33)	(121.2)	(50.54)	(178.5)	(73.34)	(48.64)	(44.49)	(98.50)
DW	1.91	1.98	1.92	2.00	1.85	1.99	1.95	1.76
Autocor.	1.6491	1.3998	0.7338	3.5054***	1.3541	0.8074	1.0760	1.9763*
R^2	0.021	0.029	0.028	0.027	0.065	0.030	0.024	0.011
Sample period : October 2001–December 2002 (n=299)								
α	-0.03620	-0.06352	-0.03036	-0.02350	-0.00590	0.00976	-0.06766	0.00838
HACSE	(0.0367)	(0.0409)	(0.0329)	(0.0231)	(0.0333)	(0.0225)	(0.0428)	(0.0194)
β	-1.0951	-0.60180	-144.96	-552.04**	-319.79*	-138.55***	-631.45**	-178.87
HACSE	(29.37)	(173.0)	(124.4)	(249.9)	(180.4)	(50.89)	(299.3)	(116.8)
DW	1.84	1.95	2.00	1.90	1.83	1.98	1.85	1.69
Autocor.	2.2701	1.6882	0.3126	0.9763	1.8579	2.0071*	1.2118	1.3030
R^2	0.019	0.030	0.032	0.046	0.035	0.044	0.071	0.013

Source: Author's calculations.

Note: [1] Estimation by ordinary least squares, robust standard errors corrected for heteroskedasticity and autocorrelation are reported in parentheses as 'HACSE'. The Durbin–Watson and error autocorrelation statistics are labeled as 'DW' and 'Autocor.', respectively. The coefficient of determination is R^2. The significance level at 1, 5, and 10 percent are denoted by '***', '**', and '*', respectively. The parameter significance is based on robust standard errors.

Most point estimates of the mean reversion processes α and β have the expected signs. We have found only four cases in Poland, and one in Slovakia, where the proportionate mean reversion process does not hold (α>0). The implication of positive αs is that the exchange rate tends to drift away from the trend indefinitely. In the cases with positive αs, however, the nonlinear process (β) overcompensates for the linear one: when the deviation becomes "large," the exchange rate adjusts swiftly. In contrast, the additional mean reversion process fails only once for the dollar exchange rate in Slovakia (β>0). Quantitatively, the point estimates of the proportionate mean reversion process (α) indicate that it would take between 7 and 60 days (up to 12 weeks) for the exchange rate to return to its trend in the absence of the nonlinear process. As expected, the speed of adjustment in the euro segment of the national foreign exchange markets was the fastest in the Czech Republic and the slowest in Poland and Slovakia.

Accounting for nonlinear mean reversion removes some of the excess kurtosis from the data. Comparing the excess kurtosis statistics from Table 2 with the excess kurtosis statistics estimated for the residuals from Tables 3 to 6, their values are declining across currencies and periods, although the statistics itself rarely become insignificant.

To test for the presence of random walk and to examine the robustness of our results to residual autocorrelation, we re-estimated equation 4 with 24 lagged dependent variables included in each regression (Table 4). Regarding the former question, we find rather limited evidence that random walk is present. On the one hand, the estimated parameters of one-period lagged dependent variables (Δe_{t-1}) are not statistically different from zero, thus suggesting consistency with the random walk hypothesis. On the other hand, and more importantly, the estimated parameters at other lags (Δe_{t-1-i}) are significant, although at different lag lengths for various currencies. As a result, the explained variance in the model (4) rises sharply. Regarding the latter issue, residual autocorrelation became less of a problem in most regressions, although a few stubbornly correlated residuals remained (e.g., the forint/euro series).

In the augmented equation (4) both the coefficients on the proportional and nonlinear mean reversion increased in magnitude and for many of them we can reject the null hypothesis of α or β, or both, being equal to zero. For example, seven out of the twelve estimated β parameters for the euro exchange rates are statistically significant and all are negative and large. Only a handful of the estimated α parameters are positive and none of them is significantly different from zero. Thus, while we cannot reject the random walk hypothesis completely, it appears to be dominated by the endogenous liquidity processes.

Second, the estimated βs increase (and often become statistically significant) as we shorten the sample—with the notable exception of the Czech koruna data—perhaps as a manifestation of additional interest in those currencies from the side of international foreign exchange market makers. These findings are consistent with our observations regarding market depth. Why, then, the decline in βs in the Czech koruna and post-2000 Hungarian forint results? We see two complementary explanations. First, the volume of Czech koruna foreign exchange trades peaked at U.S. dollar 5 billion per day in late 1998/early 1999 and declined to less than one half thereafter (Figure 2). Second, we interpret declining βs as a sign of a gradually maturing market: from the recursive estimates we see that as the absolute-value βs decline, the absolute-value αs increase (Figure 6). In other words, these markets no longer require a sizable exchange rate shock to start an endogenous-liquidity-driven return back to the trend.

Table 4: Daily Data With Lagged Dependent Variables[1]

	Czech Republic		Hungary		Poland		Slovakia	
	Dollar	**Euro**	**Dollar**	**Euro**	**Dollar**	**Euro**	**Dollar**	**Euro**
	Sample period : May 1998–December 2002 (n=1141)							
α	-0.02603**	-0.04146**	-0.04714***	-0.03504*	0.01136	-0.01984	-0.05347***	-0.04170***
HACSE	(0.0121)	(0.0169)	(0.0142)	(0.0190)	(0.0174)	(0.0162)	(0.0152)	(0.0093)
β	-44.565***	-172.48**	-77.618	-99.9559	-191.97***	-77.6933*	-16.402	-37.155***
HACSE	(16.86)	(78.50)	(47.29)	(158.7)	(57.20)	(42.17)	(36.34)	(14.44)
DW	2.00	2.00	2.00	2.00	1.99	2.00	2.00	2.01
Autocor.	1.0769	0.8182	0.5858	3.8167***	0.7565	2.0378**	0.1900	2.8731*
R^2	0.047	0.056	0.052	0.077	0.081	0.076	0.053	0.101
	Sample period : May 2000–December 2002 (n=654)							
α	-0.04241***	-0.08747***	-0.04340**	-0.02055	0.01112	-0.00909	-0.06740***	-0.03568**
HACSE	(0.0154)	(0.0241)	(0.0189)	(0.0241)	(0.0236)	(0.0199)	(0.0195)	(0.0154)
β	-23.125	-52.889	-71.2372	-149.06	-188.91*	-86.951*	4.9190	-10.313
HACSE	(19.00)	(124.7)	(52.54)	(176.2)	(91.74)	(48.71)	(43.27)	(97.73)
DW	2.00	2.00	2.00	2.01	1.99	1.98	2.00	2.01
Autocor.	0.5460	0.9448	0.8339	5.6865***	6.0222***	4.1287	0.5122	2.2694**
R^2	0.062	0.068	0.055	0.012	0.094	0.109	0.068	0.079
	Sample period : October 2001–December 2002 (n=299)							
α	-0.04081	-0.09434**	-0.03560	-0.02515	-0.02948	0.00146	-0.13085***	0.00510
HACSE	(0.0328)	(0.0455)	(0.0346)	(0.0261)	(0.0377)	(0.0191)	(0.0504)	(0.0178)
β	-18.282	-49.546	-199.03	-526.91**	-312.92	-166.67***	-759.54**	-190.80*
HACSE	(31.94)	(175.7)	(145.4)	(254.9)	(237.3)	(49.00)	(348.4)	(118.6)
DW	1.99	2.00	2.00	2.03	2.01	2.05	2.00	2.00
Autocor.	0.4484	0.4273	0.3729	4.5363***	1.4056	1.9929	0.0672	0.7709
R^2	0.100	0.100	0.079	0.127	0.133	0.176	0.162	0.096

Source: Author's calculations.

Note: [1] 24 lagged values of the dependent variable are included in the regression, but are not reported. Estimation by ordinary least squares, robust standard errors corrected for heteroskedasticity and autocorrelation are reported in parentheses as 'HACSE'. The Durbin–Watson and error autocorrelation statistics are labeled as 'DW' and 'Autocor.', respectively. The coefficient of determination is R^2. The significance level at 1, 5, and 10 percent are denoted by '***', '**', and '*', respectively. The parameter significance is based on robust standard errors.

4.2 Weekly Data

We re-estimated equation 4 with the weekly data with a dual objective in mind (Tables 5 and 6). First, we can investigate the time horizon of the endogenous liquidity hypothesis. For example, the speed of adjustment may be better captured in lower-frequency data.[15] Second, we can explore the robustness of the daily data results.

We find little difference in the overall explanatory power of the model with the weekly data as compared to the daily data. It appears that the nonlinear error correction mechanism is at play in the same countries as with the daily data and there is not much information one can use for deciding between the daily-data and weekly-data results. As expected, the weekly data are less choppy than the daily data and, as a result, the adjusted R^2 is somewhat higher, explaining between 5 and 10 percent of the variance of the dependent variable.

Similar to the daily data estimates, we have some reservations with respect to the serially correlated residuals in the model without lagged dependent variables (Table 5) and, hence, we focus primarily on the full-blown estimates with 18 lagged dependent variables (Table 6). All α parameters have the expected negative signs and only four out of the 16 point estimates in the two sample periods are not statistically different from zero (all either in the shortest subsample or in the U.S. dollar data). The estimates indicate that it would take between 2 and 15 weeks for the exchange rate to return to its long-term value in the absence of the nonlinear process. The size of the estimated β parameters and their statistical significance appears to be comparable to the daily data models. The findings with respect to the random walk hypothesis seem to be comparable to the daily data results. The endogenous liquidity hypothesis seems to be validated especially in the full sample of the Czech koruna and shorter samples of the Hungarian forint.

For comparison, the median Faruqee–Redding estimate of α for weekly data (0.18) in the G-7 countries is quite close to our median estimate of α (0.15). Our results imply, however, a less forceful nonlinear adjustment: our estimates of β are a multiple of that in Faruqee and Redding. This is a fairly intuitive result—it takes a larger departure from the trend in transition economies, as compared to the G-7 countries, to attract endogenous liquidity to the market.

[15] We re-estimated equation 4 also with the monthly data and obtained results comparable to the weekly and daily data. Given the rapidly declining number of degrees of freedom in our regressions—only for the Czech Republic do we have 44 observations with a consistent exchange rate arrangement—we refrain from presenting the estimates. Needless to say, the endogenous liquidity hypothesis holds for the Czech Republic with the monthly data as well.

Table 5: Weekly Data Without Lagged Dependent Variables[1]

	Czech Republic		Hungary		Poland		Slovakia	
	Dollar	**Euro**	**Dollar**	**Euro**	**Dollar**	**Euro**	**Dollar**	**Euro**
				Sample period : May 1998–December 2002 (n=216)				
α	-0.03932	-0.06595	-0.10352	-0.02543	-0.08316*	-0.02849	-0.09276	-0.01522
HACSE	(0.0524)	(0.0542)	(0.0630)	(0.0580)	(0.0497)	(0.0633)	(0.0572)	(0.0341)
β	-62.088*	-123.32	-117.30	-308.80	-33.319	-110.33	-17.700	-180.12***
HACSE	(37.04)	(131.5)	(73.32)	(271.7)	(64.03)	(120.4)	(66.60)	(58.99)
DW	1.83	1.69	1.88	2.03	1.86	1.97	1.77	1.86
Autocor.	1.7807*	1.1746	2.6335**	1.6027	1.7495*	2.5691**	3.4597***	1.5652
R^2	0.0892	0.047	0.086	0.059	0.051	0.055	0.051	0.047
				Sample period : May 2000–December 2002 (n=135)				
α	-0.04097	-0.02185	-0.07385	-0.01678	-0.09572	-0.02186	-0.13105*	-0.08499
HACSE	(0.0794)	(0.0721)	(0.0823)	(0.0736)	(0.0643)	(0.0752)	(0.0770)	(0.0607)
β	-85.983*	-317.37**	-129.37	-620.72**	1.8589	-108.75	-34.442	-102.57
HACSE	(49.90)	(161.6)	(85.12)	(311.9)	(60.36)	(135.5)	(76.58)	(248.5)
DW	1.89	1.80	1.98	2.06	1.84	2.00	1.78	1.85
Autocor.	1.4643	0.9688	1.9404	2.0512*	0.9668	2.1794**	2.2045	1.5802
R^2	0.062	0.049	0.072	0.101	0.044	0.058	0.068	0.057

Source: Author's calculations.

Note: [1] Estimation by ordinary least squares, robust standard errors corrected for heteroskedasticity and autocorrelation are reported in parentheses as 'HACSE'. The Durbin–Watson and error autocorrelation statistics are labeled as 'DW' and 'Autocor.', respectively. The coefficient of determination is R^2. The significance level at 1, 5, and 10 percent are denoted by '***', '**', and '*', respectively. The parameter significance is based on robust standard errors.

Table 6. Weekly Data With Lagged Dependent Variables[1]

	Czech Republic		Hungary		Poland		Slovakia	
	Dollar	**Euro**	**Dollar**	**Euro**	**Dollar**	**Euro**	**Dollar**	**Euro**
	Sample period : May 1998–December 2002 (n=216)							
α	-0.14599**	-0.12525**	-0.20284***	-0.08138	-0.21714**	-0.18393***	-0.16471***	-0.09569**
HACSE	(0.0462)	(0.0544)	(0.0590)	(0.0566)	(0.0968)	(0.0706)	(0.0521)	(0.0442)
β	-57.678	-261.50	-85.643	-683.54***	31.272	-83.9694	-40.280	-204.11*
HACSE	(44.92)	(169.8)	(78.69)	(262.3)	(70.09)	(119.9)	(61.47)	(110.4)
DW	1.99	2.00	1.97	2.06	1.99	2.00	1.99	2.0
Autocor.	1.0444	1.7225	1.0720	1.9527*	0.9712	2.2851**	0.9875	7.3536***
R^2	0.204	0.217	0.197	0.232	0.138	0.209	0.188	0.154
	Sample period : May 2000–December 2002 (n=135)							
α	-0.19665***	-0.09379	-0.36069**	-0.04481	-0.23740	-0.14425*	-0.30068***	-0.15068*
HACSE	(0.0564)	(0.0716)	(0.1074)	(0.0894)	(0.1533)	(0.0795)	(0.0894)	(0.0795)
β	-140.24***	-312.49*	-64.6889	-994.60***	54.021	-110.76	-39.103	-382.82
HACSE	(50.75)	(186.2)	(84.93)	(323.5)	(89.08)	(127.3)	(84.71)	(247.7)
DW	1.99	2.01	1.96	2.11	1.89	1.96	1.97	2.01
Autocor.	1.1319	4.0984***	1.0813	4.3844***	2.8731***	1.3880	1.1936	1.5713
R^2	0.262	0.190	0.225	0.302	0.135	0.284	0.207	0.242

Source: Author's calculations.

Note: [1] 18 lagged values of the dependent variable are included in the regression, but are not reported. Estimation by ordinary least squares, robust standard errors corrected for heteroskedasticity and autocorrelation are reported in parentheses as 'HACSE'. The Durbin–Watson and error autocorrelation statistics are labeled as 'DW' and 'Autocor.', respectively. The coefficient of determination is R2. The significance level at 1, 5, and 10 percent are denoted by '***', '**', and '*', respectively. The parameter significance is based on robust standard errors.

4.3 Policy Implications

Our results have some attractive policy implications. First, the results suggest that early liberalization of the foreign exchange market and financial-account transactions may pay off in terms of market liquidity and, hence, faster adjustment of the exchange rate. More stable currencies, that is, currencies with shorter deviations from the trend, are those that have more liquid markets. However, early liberalization is a necessary condition, not a sufficient one. For example, although the Slovak koruna was floated soon after the Czech koruna, the Slovak market remained thin and the koruna turnover on the foreign exchange market low (Figure 2). Similarly, the floating of the Hungarian forint in 2001 and the rapid growth in forint turnover appear to be associated with fast-growing endogenous liquidity. Moreover, the results for the Czech koruna suggest that a deep, maturing market is likely to be characterized by a gradually increasing weight of the linear adjustment processes. Comparatively, the speed of proportional adjustment of the Czech koruna and, more recently, Hungarian forint is gradually becoming comparable to that of the G-7 countries, although the nonlinear portion remains slower.

Second, the endogenous liquidity effect dominates empirically the autoregressive processes of the random walk. In summary, endogenous liquidity does not imply that the volatility of the nominal exchange rate is lower in a liquid market, it only implies that the self-correcting mechanism is fast. The endogenous liquidity hypothesis seems to hold especially well in the full sample of the Czech koruna and shorter samples of the Hungarian forint.

Third, we conjecture that national exchange rate arrangements have a peculiar impact on the behavior of the exchange rate. For example, the Polish zloty is an outlier—in terms of similar volatility in both the dollar and euro exchange rates, wrong signs of the proportional mean reversion process (α), and comparatively large βs in the daily regressions. The results suggest that the zloty exchange rate adjusts only if the departure from the trend is large and that in the case of smaller deviations the rate tends to drift further from the trend, because a limited amount of endogenous liquidity is attracted into the market. We find similar results for the euro rate of the Slovak koruna in the 2001-2002 period.

5. Concluding Remarks

Utilizing the models of Faruqee and Redding (1999) and Evans and Lyons (2002), we test endogenous liquidity provision in four Central European foreign exchange markets. In this model, order flow is attracted to a market with a seemingly temporary shock to the exchange rate, reacting disproportionately fast to sizeable deviations from the longer-term exchange rate path. The testable equation contains a univariate error-correction mechanism that captures both the linear and nonlinear parts of the mean reversion process, where the latter part is driven by the endogenous liquidity hypothesis. In addition, the model can be extended to allow testing of the random walk hypothesis.

The empirical part of the paper validated the theoretical priors and found that the depth of the market may have implications for the speed of adjustment of the error-correction mechanism. We found the strongest impact of the endogenous liquidity hypothesis in the Czech koruna market and, more recently, also in the Hungarian forint market. Quantitatively, the results are comparable to those of Faruqee and Redding for G-7 countries. The model was tested using weekly and daily exchange rate data for the U.S. dollar and euro against four Central European currencies for the 1998–2002 sample period, the Czech and Slovak korunas, Hungarian forint, and Polish zloty.

Liquid currencies, that is, those with a sufficient number of market makers and deep order books, tend to have shorter deviations from the trend exchange rate. The results seem to suggest a connection between the early liberalization of the foreign-exchange market and financial-account transactions in the Czech Republic and the resulting, faster-than-average adjustment of the Czech koruna to short-term deviations from the trend. A similar link between market deepening and the speed of adjustment of the exchange rate is observable in the post-2000 Hungarian data. The endogenous liquidity hypothesis seems to dominate the effects of the random walk hypothesis of exchange rate determination. We found limited clues, however, as to what is the appropriate frequency for testing the endogenous liquidity hypothesis—we have to wait for longer time series under consistent exchange rate arrangements. Further empirical extensions of the model may want to consider, *inter alia*, the impact of central bank interventions on the speed of exchange rate adjustment.

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