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Yield Curve Dynamics and Spillovers in Central and Eastern European Countries

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Yield Curve Dynamics and Spillovers in Central and Eastern European Countries[1]

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Abstract

This paper applies the models used to study yield curve dynamics and spillovers in the U.S. and other countries to Central and Eastern European countries (CEE countries). Using the Diebold, Rudebusch, and Aruoba (2006) dynamic version of the Nelson-Siegel representation of the yield curve, the paper finds that the two-way relationship between macroeconomic and financial variables in the CEE countries is similar to the one in mature economies. However, inflation shocks have very little persistence in the CEE countries, owing to the strong convergence trends in these countries—which tend to re-anchor expectations faster. Increased convergence in policies and market integration over time are associated with a stronger correlation between the levels of the yield curves, while the curves slopes are more driven by idiosyncratic factors. Shifts in the euro yield curve are transmitted both to interest rates and inflation expectations in the CEE countries—and transmission is stronger after 2004.

JEL Classification Numbers: E43, E44, F42, G15

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I. INTRODUCTION

Market participants, central bankers, and academics are increasingly using yield curves to extract information about expectations and forecast macroeconomic variables. This has stimulated recent advances in modeling macro-financial linkages, as well as spillovers between increasingly integrated (global) bond markets. These advances have mostly focused on mature markets' yield curves. In this paper, we use some of these modeling strategies to study the macro-finance linkages and cross-border spillovers in a group of emerging markets, the Czech Republic, Hungary, and Poland, referred subsequently as the "Central and Eastern European" (CEE) countries.

Diebold, Piazzesi, and Rudebusch (2005) summarize recent advances in modeling macro-finance linkages to understand the dynamics of yield curves, and note that a relatively high share of yield curve variance can be explained by macroeconomic variables. Models from both the finance tradition, which are based on a no-arbitrage framework, and the Nelson and Siegel (1987) approach, have shown promising results when combined with macroeconomic variables. Another strand of the literature, which studies linkages across financial markets, points to the dominant role played by spillovers from the U.S. to foreign markets, even when controlling for the role of macroeconomic factors (Ehrmann and Fratzscher, 2005; and Bayoumi and Swiston, 2007).

For a number of reasons CEE countries provide an interesting case to study whether some of the results found for the mature markets still hold in emerging markets. First, since the mid-1990s, these countries have developed relatively large and liquid fixed-income markets with strong participation from institutional investors from Western Europe and other countries (Roldós, 2004). Second, their yield curves have displayed a fair amount of volatility, as a result of investors "convergence plays" in these markets and the difficulties associated with macroeconomic management on the road to accession to the EU in 2004 and the forthcoming adoption of the euro by the CEE countries (Schadler, 2005).

In the spirit of Diebold, Rudebusch, and Aruoba (2006), this paper uses a dynamic version of the Nelson-Siegel (NS) representation of the yield curve, but with a different estimation strategy and incorporating open economy variables. Moreover, the paper employs generalized impulse responses of the NS parameters—representing the level, slope and curvature of the curve—that have been re-aggregated to study shifts in the full yield curve following macroeconomic and financial shocks. The analysis provides insights into the two-way dynamic relationship between the yield curve and macroeconomic variables such as inflation expectations, business cycle developments, and real exchange rates in the CEE countries. Spillovers from the yield curves of the three countries noted above (CEE), and in particular from the euro area, are characterized with impulse responses as well as with variance decompositions—which allow for an examination of the importance of global versus domestic factors in the dynamics of CEE countries yield curves.

The paper is organized as follows. The next section briefly reviews the related literature and summarizes the main stylized facts of yield curves in these countries. Section III describes the empirical strategy used to study both macro-finance linkages and spillovers, and Section IV presents the empirical results. Building on the empirical framework developed in these last two sections, Section V examines the potential macroeconomic effect of Poland's euro adoption announcement on September 10, 2008 and its potential spillover effects. A final section concludes.

II. Background

A. Literature Review

The recent literature on macro-finance linkages and yield curve modeling attempts to understand the relationship between macroeconomic variables, and yield curve factors. This literature has evolved in two strands (Diebold, Piazzesi, and Rudebusch, 2005). The first one starts from typical affine, no-arbitrage, latent (unobservable) factor models, and adds macroeconomic variables that may explain the evolution of such factors.[2] The second strand does not impose no-arbitrage restrictions and focuses on the dynamic relationship between Nelson-Siegel (1987) representations of the yield curve and macroeconomic variables. Most studies have examined U.S. government bond markets, but a few have studied European countries. Recently, some studies have started to look at the spillovers across mature economies' bond markets—which are increasingly important with the integration and globalization of bond markets.

The integration of macroeconomic variables into affine (linear) no-arbitrage models has had some empirical successes. Piazzesi (2005) incorporates the Fed's target rate into an otherwise latent factor model, and shows that pricing errors are reduced. Ang and Piazzesi (2003) add measures of inflation and real activity to a vector auto-regression (VAR) with the latent factors for U.S. yield curve data, and find that up to 85 percent of the forecast variance at short-and medium-term maturities can be explained by macroeconomic factors—with inflation being the dominant macroeconomic factor. More recent studies have focused on improving the misfit of the long end of the term structure. Rudebusch and Wu (2004) combine a no-arbitrage model of the term structure with a standard new-Keynesian macro model with an inflation target, while Dewachter and Lyrio (2006) introduce a filtered measure of long-run inflation expectations. Both studies find that inflation expectations determine the level of the yield curve, while cyclical factors (including the monetary authority response) determine the slope of the yield curve.

The second strand relates the Nelson-Siegel (NS) representation of the yield curve to macro variables, without imposing the no-arbitrage conditions.[3] Diebold and Li (2006) show that the three time-varying NS parameters may be interpreted as factors corresponding to the level,

[2] Affine (linear), no-arbitrage models are the preferred ones in finance. For a survey of the voluminous literature, see Piazzesi (2003)

[3] Diebold and others (2006) argue that, to the extent that the no-arbitrage condition is satisfied in the data for the U.S., it would also be satisfied with the flexible specification of the NS yield curves. In other countries, transitory arbitrage opportunities may persist for some time, so imposition of the restriction could also lead to misspecification.

slope, and curvature of the yield curve, and Diebold and others (2006) use VAR analysis and variance decompositions to show strong evidence of bidirectional causality between macroeconomic and yield factors. When bidirectional feedback is considered, half the variance of long yields can be attributed to macro factors. In a standard VAR with macro variables and three yields (1, 12, and 60 month maturities), Evans and Marshall (2007) also find that macro impulses account for a larger fraction of the variance of long-term yields when an interest rate-smoothing equation is incorporated. They also find that, while technology and preference shocks have a significant impact on yields, fiscal policy shocks are not an important source of interest rate variability.[4]

A few studies from European countries have found similar linkages between macro and yield factors, and a nascent literature on spillovers has started to study linkages across major government bond markets.[5] In particular, Diebold and others (2007b) extract global and country-specific yield factors from the term structure of government bonds in the U.S., Germany, Japan, and the U.K. The estimated (unobservable) global factors explain more than half the variance of yields, with a bigger impact on levels than on the slope of individual country yield curves. This reflects the convergence of inflation levels, as well as the lower correlation of business cycles and monetary policies across the major industrialized countries.

B. Stylized Facts on CEE Countries

Most of the studies summarized above focus on the U.S. yield curve, which displays on average a positively sloped yield curve (between 5.5 percent at the short end and 7 percent at the long, over the last 20 years; see Diebold and Li, (2006)). Similarly, Lemke (2008) finds an average upward-sloping yield curve for the euro (between 3 and 4½ percent; see his Figure 4). In contrast, the CEE countries have experienced negatively sloped yield curves for many years.[6] The accession to the European Union (EU) in May 2004 marked the end of the first phase of these countries' integration with Western Europe (Schadler, 2005) and is likely to have changed the linkages between the two groups of countries. A closer look at these countries' yield curves is thus warranted.[7]

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[4] In previous work, Evans and Marshall (1998) had shown that monetary policy shocks raise the level of the yield curve, and reduce its slope and curvature.

[5] See, for instance, Hordahl and Tristani (2007), and Bayoumi and Swiston (2007).

[6] The data cutoff date for this section and the empirical estimation that follows below was end-May 2008. Thus, the extreme turmoil in financial markets in the second half of 2008 was excluded from the sample period. Kladívko (2008) documents the high level of volatility in the Czech treasury bond market and how it differed from the less turbulent period.

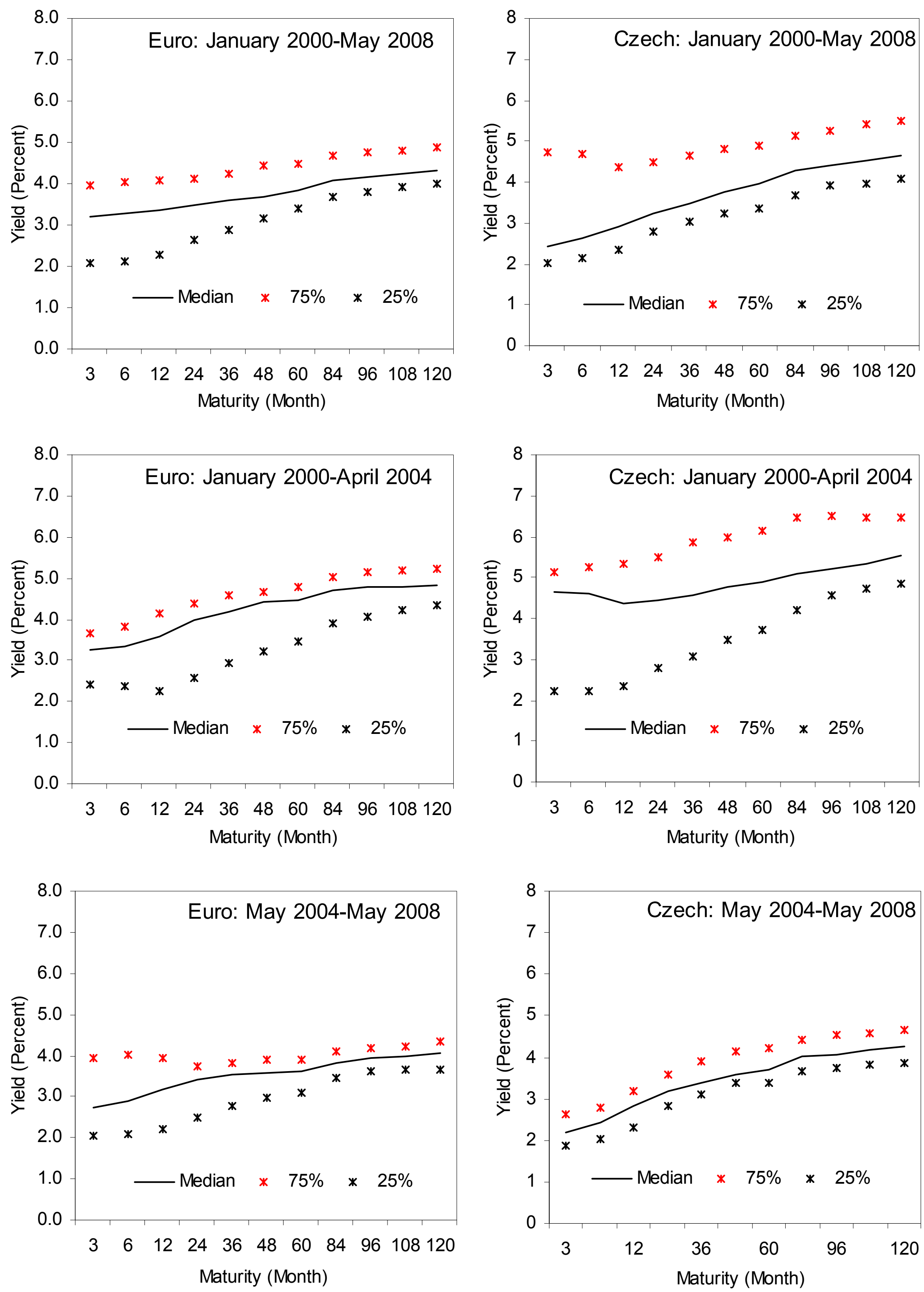
[7] Yield data for our study cover the period 2000-08. The yield curve data are based on zero coupon yields for government bonds with maturities of 3, 6, 9, 12, 24, 36, 48, 60, 84, 96, and 120 months and is from Bloomberg. Bloomberg data for the euro area are a composite of sovereign yields of Spain, Austria, Germany, France, Ireland, Finland, and the Netherlands.

Figures 1 and 2 display the median yield curves for the euro area and the three CEE countries, together with the 25 and 75 percentiles, for the whole sample and for two subsamples (January 2000-April 2004 and May 2004-May 2008). The euro median curve is upward sloping in the first half of the sample, an easing period during which the curve shifted down and became very steep (as illustrated by the greater distance of the 25 percentiles, as opposed to the 75 percentiles, from the median). In the second subsample, the curve became flatter and much more tightly anchored at the long end, at a lower level of 3½ -–4 percent.

The median yield curve for the whole sample is upward sloping for the Czech Republic, downward sloping for Hungary, and flat for Poland.[8] However, the sample median hides a fair amount of volatility and changing patterns for each country, and the lower panels provide evidence of the change in regime experienced around 2003-04. Low inflation and an appreciating koruna have contributed to a relatively low and slightly upward-sloping yield curve in the Czech Republic, especially after sharp policy rate cuts in 2003 and 2005 left the Czech National Bank rate below the euro rate. Hungary's yield curve remained inverted during most of our sample period, with brief periods of flat curves in 2003, 2005, and 2007. Poland's steep, negatively sloped yield curve at the beginning of the century un-inverted in 2003, and has since been flat or positively sloped—except over the last few months of our sample. A more detailed description of the main policy developments driving each country's yield curve dynamics is presented in Appendix I.

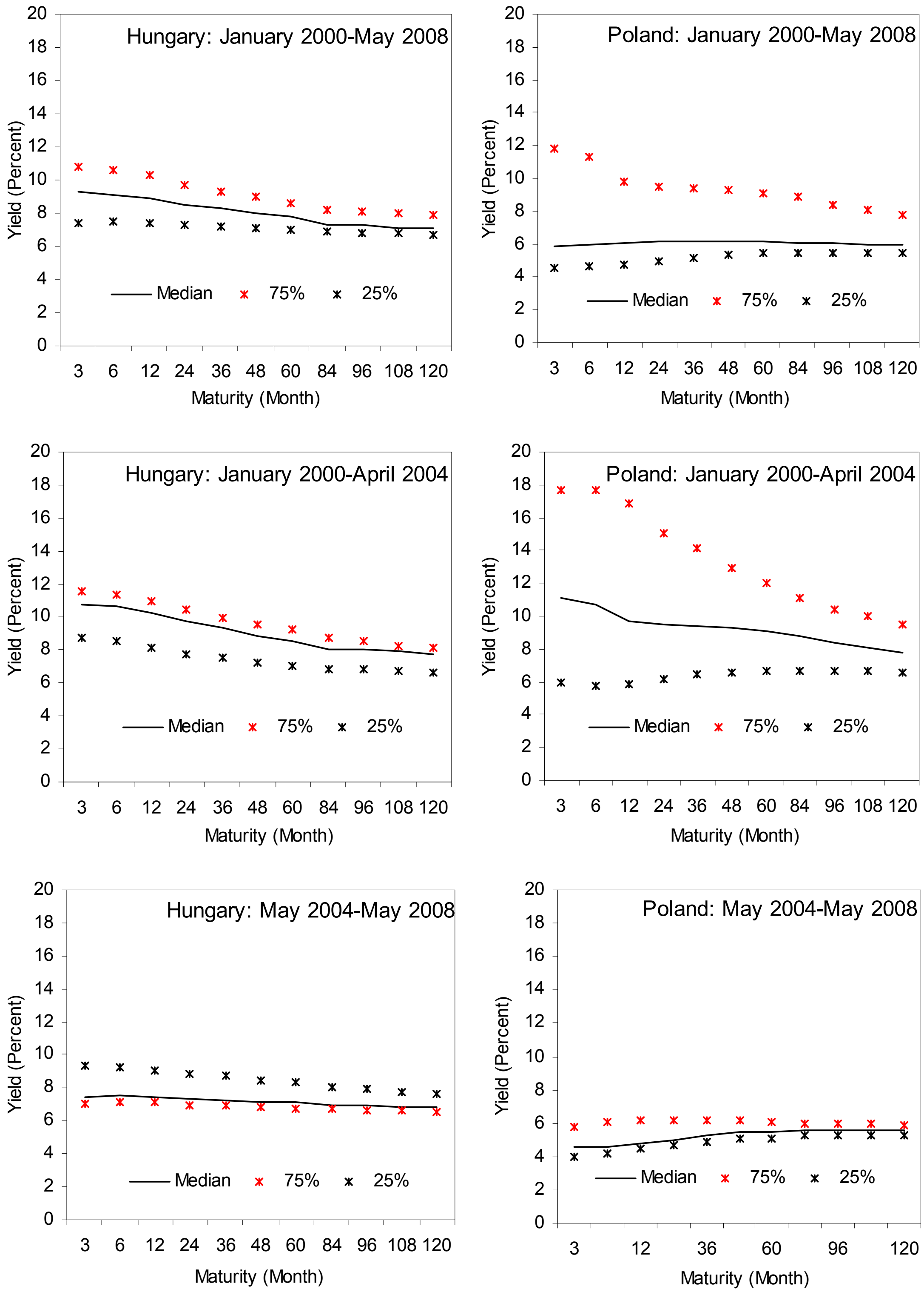
[8] The 25-75 percentiles document the well-known stylized fact of higher volatility in the short-end of the curve compared to the long end; they also show that the distribution of yields tend to be asymmetric, with relatively fat right tails.

Figure 1. Euro Area and the Czech Republic: Median Yield Curve, 2000–08



Source: Bloomberg.

Figure 2. Hungary and Poland: Median Yield Curve, 2000–08



Source: Bloomberg.

III. Modeling the Yield Curve and Spillovers

A two-pronged modeling strategy is followed to assess linkages between macro and yield factors, as well as spillovers across countries. First, Nelson-Siegel factors are related to macro variables, to confirm that the typical results obtained for developed economies still hold for these emerging economies. Second, the nature and importance of cross-country spillovers and regional or global factors in domestic yield curves are assessed with the use of the euro area yield curves.

A. The Nelson-Siegel Representation

Following Diebold and Li (2006), this paper models the yield curve using a variation of the three-component exponential approximation proposed by Nelson and Siegel (NS, 1987):

$$y_t(\tau) = \beta_{1t} + \beta_{2t}\left(\frac{1 - e^{-\lambda_t \tau}}{\lambda \tau}\right) + \beta_{3t}\left(\frac{1 - e^{-\lambda_t \tau}}{\lambda \tau} - e^{-\lambda_t \tau}\right),$$

where $y_t(\tau)$ denotes the yield at maturity τ (month) and the subindex *t* refers to the time period. As noted above, the time-varying parameters of the NS curve can be interpreted as factors representing the level (β_{1t}), slope (β_{2t}), and curvature (β_{3t}) of the yield curve. The corresponding loadings are a constant (one) and the terms in square brackets that vary with maturity. Diebold and Li (2006) justify this interpretation of the NS factors on the grounds that

- β_{1t} is the factor whose loading does not change with maturity and thus affects all yields by the same amount;
- β_{2t} is the factor whose loading equals one at zero maturity and declines to zero as the maturity increases;[9] and
- β_{3t} is the factor whose loading displays a humped shape as it starts at zero at zero maturity, increases at intermediate maturities, and falls back to zero at longer maturities.

As noted by Dewachter and Lyrio (2006, p. 121) for the U.S., "the level effect can be linked to long-run inflation expectations... the slope factor correlates well with predictable inflation and business cycle components and... the curvature effect is related... to real interest rate

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[9] In this study, as in Diebold and Li (2006), the slope of the yield curve is taken to be the difference of $y_t(3)$ and $y_t(120)$, which approximately equals β_{2t} .

movements not related to standard macroeconomic conditions." The parameter λ determines the speed of the (exponential) decay of the approximation, with smaller (larger) values associated with slow (fast) decay rates that are better able to approximate the longer (shorter) end of the yield curve.

B. Macro-Financial Linkages

The empirical relationship between the yield curve and the economy is characterized using the generalized vector autoregressive (GVAR) analysis of Koop, Pesaran, and Potter (1996). The GVAR does not impose restrictions on the estimated covariance matrix and has the advantage of fully capturing the true historic dynamic linkages among the variables. A further advantage is that generalized impulse response functions are unique, that is, order invariant. While GVAR analysis neither identifies nor recovers structural shocks—the focus of other VAR analysis—it is in the spirit of the theoretical analysis of Sims (1980).

For each country, and in the spirit of Diebold, Rudebusch, and Aruoba (2006), the following vector autoregressive (VAR) model,

$$X_t = \mathrm{C}(L)^{-1}\mu_t$$

$$X_t = \left[i_t^*, \beta_{1,t}, \beta_{2,t}, \beta_{3,t}, IP_t, REER_t, \pi_t, i_t^P \right]',$$

is estimated. As in Diebold, Rudebusch, and Aruoba (2006), the NS factors (the $\beta_i's$) are included, together with standard macro variables: industrial production (*IP*), inflation (π), and the domestic policy interest rate (i^P). Taking into account the small, open economy characteristics of the countries under study, and anticipating the analysis of spillovers in the next section, the model also includes the euro area policy interest rate (i^*) and the real effective exchange rate (*REER*). With the exception of the $\beta_i's$ and interest rates, all variables are expressed as 12-month year-on-year changes. C(*L*) is a lag polynomial matrix with *L*=3.[10] The shocks μ are assumed to be distributed multivariate normal, N(0, Ω), and thus X_t is also distributed multivariate normal, $N\left(0, \mathrm{C}(L)^{-1} \times \Omega \times \mathrm{C}(L)^{-1}{}'\right)$. A brief description of the GVAR methodology under these assumptions is provided in Appendix II.

This specification and the GVAR methodology allow us to study a number of linkages between macro and financial variables. Typical yield curve analysis focuses on parallel shifts

[10] To determine the number of lags in the VAR models, standard lag-length tests—comparing the cost of increasing the lag length (reduced degrees of freedom) with the benefit (greater information extraction from the data—were computed. Using a maximum lag length of eight for all countries, the Akaike's information criterion test suggested using eight lags and the Schwarz test suggested using one. Models with eights lags were unstable and those with one lag were not able to ensure well-behaved residuals. The results discussed below stem from an intermediate number of lags: three. Results with four lags are qualitatively similar to those discussed here.

of the curve, steepening or flattening (tilting) movements, or a combination of the two. In this context, and to facilitate the interpretation of the results, the paper innovates; rather than report the individual impulse responses of macro shocks on the NS factors, the movements in the whole yield curve are constructed from the individual responses of the $\beta_{i,t}$'s. Shocks to the NS factors are interpreted as financial shocks, and their impact on macro variables—in the spirit of Estrella and Mishkin (1997)—is also analyzed.

C. Spillovers

To assess the influence of regional and global bond market factors on the CEE countries' yield curves, this study does two things. First, it estimates the principal components driving each NS factor for the three CEE countries, and then for the largest bond markets (U.S., Germany, and Japan).[11] Simple correlations between these regional and global components and each country's NS factors allow the study to establish the degree of co-movement in yield curve dynamics across these markets. Second, spillovers and linkages of the yield curves in the three CEE countries and the euro area are characterized by the generalized impulse responses and variance decompositions of a GVAR model comprising the three NS $\beta_{i,t}$'s for each of the CEE countries *plus* the three $\beta_{i,t}$'s of the euro area. Specifically, using the notation above, this paper estimates the following GVAR:

$$\tilde{X}_t = \tilde{C}(L)^{-1}\tilde{\mu}_t$$

$$\tilde{X}_t = \left[\beta_{1,t}^{euro}, \beta_{2,t}^{euro}, \beta_{3,t}^{euro}, \beta_{1,t}^{CZE}, \beta_{2,t}^{CZE}, \beta_{3,t}^{CZE}, \beta_{1,t}^{HUN}, \beta_{2,t}^{HUN}, \beta_{3,t}^{HUN}, \beta_{1,t}^{POL}, \beta_{2,t}^{POL}, \beta_{3,t}^{POL}\right]'.$$

with the NS factors that summarize the yield curve dynamics for the three CEE countries and the euro area.

[11] The (first) regional component of $B_i^{REG}{}' = \left[\beta_{i,t}^{CZE}, \beta_{i,t}^{HUN}, \beta_{i,t}^{POL}\right]$ for i=1, 2, and 3 was obtained as $P_i^{REG} = W_i^{REG}{}' \times B_i^{REG}$, with weights, $W_i^{REG}{}' = \left[w_i^{CZE}, w_i^{HUN}, w_i^{POL}\right]$ determined by the following solution: $\max S(W_i^{REG}) = W_i^{REG}{}' \times \Sigma_i^{REG} \times W_i^{REG}$ subject to $W_i^{REG}{}' \times W_i^{REG} = 1$ for i=1, 2, and 3, where Σ_i^{REG} denotes the (sample) variance-covariance matrix of B_i^{REG}. The solution corresponds to the (normalized) eigenvector associated with the largest eigenvalue of Σ_i^{REG} (Campbell, Lo, and Mckinley, 1997, pp. 236–7). In addition, the study considers the (first) global component—namely, $P_i^{GLO} = W_i^{GLO}{}' \times B_i^{GLO}$, where $B_i^{GLO} = \left[\beta_{i,t}^{DEU}, \beta_{i,t}^{US}, \beta_{i,t}^{JPN}\right]'$ and $\beta_{i,t}^{DEU}, \beta_{i,t}^{US}, \beta_{i,t}^{JPN}$ denote the $\beta_{i,t}$'s for Germany, the U.S. and Japan.

IV. Empirical Results[12]

This section presents the paper's empirical findings. First, the NS parameters ($\beta_{i,t}\text{'}s$) are estimated and contrasted with their yield curve and macroeconomic counterparts. Second, spillover effects are analyzed in two alternative reduced-form models that include the NS $\beta_{i,t}\text{'}s$ estimates. A brief discussion of regional and global components of the yield curve in these countries complements the discussion of spillovers.

A. The Estimated $\beta_{i,t}$'s and Macroeconomic Variables

The estimated $\beta_{i,t}\text{'}s$ capture key features of the yield curves in the CEE countries.[13] Consistent with the Diebold and Li (2006) interpretation and the available literature, the estimated $\beta_{i,t}\text{'}s$ are highly correlated with the empirical counterparts of the level, slope, and curvature of the yield curve. Specifically, the correlation of β_{1t} and the long rate $y_t(120)$ exceeds 0.8 in all three countries (Table 1). The correlation between β_{2t} and the slope of the yield curve—measured as the short rate $y_t(3)$ minus the long rate $y_t(120)$—is also strong, and the same applies to the correlation between β_{3t} and the curvature of the yield curve—measured, as in DL, as $2\times y_t(24)$ minus { $y_t(120)+y_t(3)$ }.

[12] As noted above, the cut-off date for the empirical analysis was end-May 2008.

[13] The estimation strategy employed here follows Diebold and Li (2006). For details see Appendix III.

Table 1. Correlation between the $\beta_{i,t}$ and Yield Curve

	Long-term Yield Level (y(120))	Slope (y(3) - y(120))	Curvature, (2y(24) - y(120)-y(3)
β_1			
The Czech Republic	0.83		
Hungary	0.98		
Poland	0.99		
β_2			
The Czech Republic		0.75	
Hungary		0.98	
Poland		0.99	
β_3			
The Czech Republic			0.91
Hungary			0.95
Poland			0.77

Sources: IFS, National authorities and authors' calculations.
1/ Based on consumer survey by the European Commission.

Descriptive statistics for the NS factors during the sample period (2000–08) show that the mean level factor (β_1) exceeded that of the euro area by 100 basis points or more (Table 2). The smallest margin over the euro area was found in the Czech Republic, which has also experienced the most variability. The converse was found to be true for Hungary, whose margin over the euro β_1 exceeded 150 basis points and whose standard deviation was about half of that in the other CEE countries. A noteworthy difference in these countries emerges from the estimated slope of the yield curve, β_2. Reflecting differing disinflation policies, the yield curve has been inverted—that is, short rates have exceeded long rates—on average in Hungary and Poland. The extreme value of 13.4 percentage points in Poland corresponds to November, 2000 that is, the height of the previous cycle of monetary policy tightening. As discussed above, the yield curve remains inverted in Hungary, while the yield curve has become more normalized in Poland. The Czech Republic has experienced a normal-sloping yield curve throughout the sample period. Differences also emerge when comparing curvature factors ($\hat{\beta}_3$). Hungary and Poland have seen positive curvature that, all else equal, imparts a hump to the yield curve; the Czech Republic has not.

Table 2. Descriptive Statistics for the Estimated Nelson-Siegel Factors

	β_1				β_2				β_3			
	CZE	HUN	POL	EUR	CZE	HUN	POL	EUR	CZE	HUN	POL	EUR
Mean	6.08	6.53	6.36	4.92	-2.97	2.65	2.14	-1.75	-1.28	1.83	1.98	-1.68
Std. Dev.	1.34	0.53	1.11	0.70	1.08	1.88	4.82	1.13	2.21	1.86	2.98	1.68
Median	6.35	6.53	6.02	5.18	-2.82	2.68	0.06	-1.88	-1.12	1.58	1.36	-1.95
Maximum	9.28	7.71	9.51	6.06	-1.07	6.62	13.40	-0.07	4.38	7.49	13.40	1.49
Minimum	3.81	4.99	4.73	3.50	-5.14	-0.66	-2.13	-3.59	-6.04	-1.74	-2.09	-4.98

Source: Authors' calculations.

The estimated $\beta_{i,t}$'s for the CEE countries correlate with their corresponding macroeconomic counterparts in a manner consistent with studies for other countries. Simple correlations suggest the following (Table 3):

- The level factor β_1 is positively correlated with inflation in Hungary and Poland, and, to a lesser extent, in the Czech Republic. It is also correlated with the European Commission's survey-based inflation expectations (for the next 12 months), except in Hungary.

- The slope factor β_2 is negatively correlated, as expected, with industrial production and capacity utilization in Hungary and Poland but not in the Czech Republic.

- The slope factor β_2 is positively correlated with the real effective exchange rate (REER) in all three countries. This suggests that, to the extent that the longer end of the yield curve is tied down by inflation expectations, increases in the shorter end of the curve attract capital inflows that put upward pressure on the nominal and the real effective exchange rates. Other studies on yield curve dynamics do not focus on this relationship, but the result is consistent with the literature on failures of the uncovered interest parity and the so-called Fama-puzzle (Fama, 1984).

- The curvature factor β_3 is positively correlated with interest rates in Hungary and Poland but not in the Czech Republic, where the correlation has the opposite sign.

Table 3. Correlation between $\beta_{i,t}$ and Macroeconomic Variables

	Inflation			Real Activity			Exchange Rate 2/		Monetary Stance	
	CPI (y-o-y)	Core CPI (y-o-y)	Inflation Expectations 1/	Industrial Production	Industrial Production (y-o-y)	Capacity Utilization	Real Effective Exchange Rate (y-o-y)	Real Effective Exchange Rate (std.dev.)	Nominal Interest Rate	Nominal Interest Rate (std.dev.)
β_1										
The Czech Republic	0.22	0.23	0.56							
Hungary	0.45	n.a.	0.01							
Poland	0.57	0.72	0.58							
β_2										
The Czech Republic				0.33	0.01	0.42	0.45			
Hungary				-0.54	-0.13	-0.34	0.30			
Poland				-0.65	-0.58	-0.64	0.40			
β_3										
The Czech Republic							-0.03	-0.364	0.06	-0.59
Hungary							-0.23	-0.082	0.25	0.01
Poland							0.25	-0.069	0.72	0.09

Sources: IFS, National authorities and authors' calculations.

1/ Based on consumer survey by the European Commission.

2/ At monthly frequency, movements in the real effective exchange rate reflect primarily movements in the nominal exchange rate.

B. Macro-Financial Linkages

The generalized impulse responses that characterize the two-way dynamic relationship between macroeconomic variables and the NS yield factors within each of the CEE countries are discussed next.

Impulse Responses: Macro Shocks on Yield Curves

As noted in the previous subsections, Figure 3 provides the response of each country yield curve—that is, the combined impact on the three $\beta_{i,t}$'s—shocks in the main macro variables for each country. To simplify the figures, the yield curve responses are shown for horizons of zero (that is, the same-month impact effects), three, and 12 months.[14] The main results are the following:

- **Shocks to euro area policy rate** tend to result in a relatively large hump-shaped upward shift in the yield curve at the three-month horizon. This probably reflects a rise in intermediate yields as markets anticipate similar tightening moves by the CEE central banks. The upward shifts in the whole yield curve (not just the CEE short rates) persist at least until after a year—with the exception of Hungary.

- **Shocks to economic activity** have negligible effects on the yield curve. This outcome is consistent with the view that monetary authorities in inflation-targeting countries would not respond directly to output shocks—in contrast to the U.S. experience.

- **Shocks to inflation** systematically shift up the yield curves at the three-month horizon. The Czech Republic experiences a parallel upward shift in the yield curve, while in Hungary and Poland there is a tilting of the curve. The effects die out after one year, except in Hungary, where the curve tilts for a longer period of time.

- **Shocks to the REER** (an unexpected strengthening of the exchange rate) are associated with a small but persistent downward shift in the entire yield curve, with a somewhat larger response in Hungary. This result suggests that markets expect the central bank to "lean against the wind" in response to an exchange rate appreciation by lowering the short-term interest rates, and that the response of future short rates would accelerate the convergence to lower inflation /interest rates. The responses are relatively small, but more pronounced in Hungary, where monetary policy attaches

[14] For the discussion of the yield curve impulse responses below, it is useful to note that impulse responses depict the effect on the variable of interest compared with its baseline. Thus, a negative response of the yield curve does not imply that yields turn negative, but rather that yields fall below their baseline.

- more weight to an exchange rate objective, in the context of an exchange rate band that was in place until January 2008.[15]

- **Shocks to domestic policy rates** are, not surprisingly, rapidly transmitted to the shorter end of the yield curve, reflecting the liquidity effect (Evans and Marshall, 1998). The three yield curves shift up in a humped fashion and become inverted only after three months in the Czech Republic and Hungary. The response of Poland's yield curve differs in two respects: the entire yield curve shifts upward in a parallel fashion, and the effect persists longer than in the other countries.

Impulse Responses: Yield Factor Shocks on Yield Curves

In broad terms, the generalized impulse response of the yield curve in the Czech Republic and Poland have been found to be qualitatively similar—likely reflecting their inflation-targeting regimes—and to contrast with the responses in Hungary (Figure 4):

- **Shocks to the level factor (β_1)** are associated with an upward shift of the curve, combined in the short run with a positive tilt. While the shift in the level of the curve vanishes within a year in Hungary, it is more persistent in Poland and to a lesser extent in the Czech Republic.

- **Shocks to the slope (β_2)** are associated with a "pure" negative tilt in the yield curve, that is an inversion where the change at the long end of the curve, dips below zero, (it falls below the baseline). This is consistent with markets expecting long-run inflation to decline following the shock. The different monetary/exchange rate regime in Hungary suggests that shocks to short-term interest rates are associated with exchange rate moves, and do not have much long-term effect on inflation expectations—remain well anchored at the same level.

- **Shocks to the curvature (β_3)** are associated with a hump-shaped yield curve on impact, but the hump vanishes fairly quickly thereafter. Convexity/curvature opportunities are generally exploited by sophisticated fixed-income traders who are likely to arbitrage them away relatively quickly.

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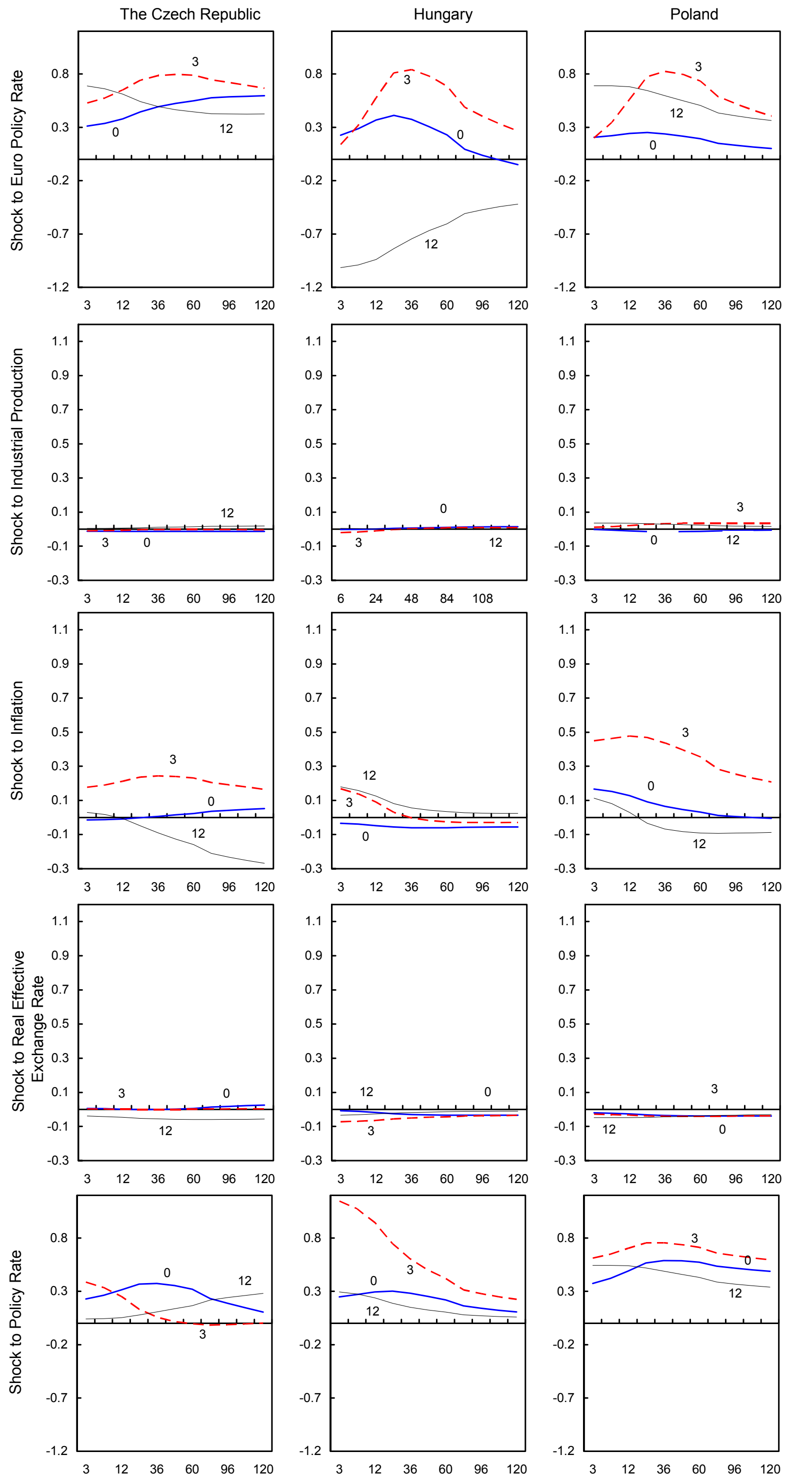
[15] The IMF's *Annual Report on Exchange Arrangements and Exchange Restrictions*, AREAER (1997) classified Hungary's exchange rate regime as a pegged exchange rate within a horizontal band (+/- 15 percent around the central parity).

Figure 3. Central European Countries: Generalized Impulse Response of Yield Curves to 1 percent Macroeconomic Shocks at 0, 3, and 12 Month Horizons, 2000–08



Source: Authors' calculations.

Based on three country VAR models with x = [i*, β_1, β_2, β_3, IP, REER, π, i^p] and L = 3.

Impulse Responses: Yield Shocks on Macro Variables

As in other studies,[16] the responses of the macro variables to shocks in the $\beta_{i,t}$ '*s* are small and measured imprecisely—with a bit more precision for those of the REERs. The macro responses in Czech Republic and Poland are more in line with those of other studies than those for Hungary (Figure 5):

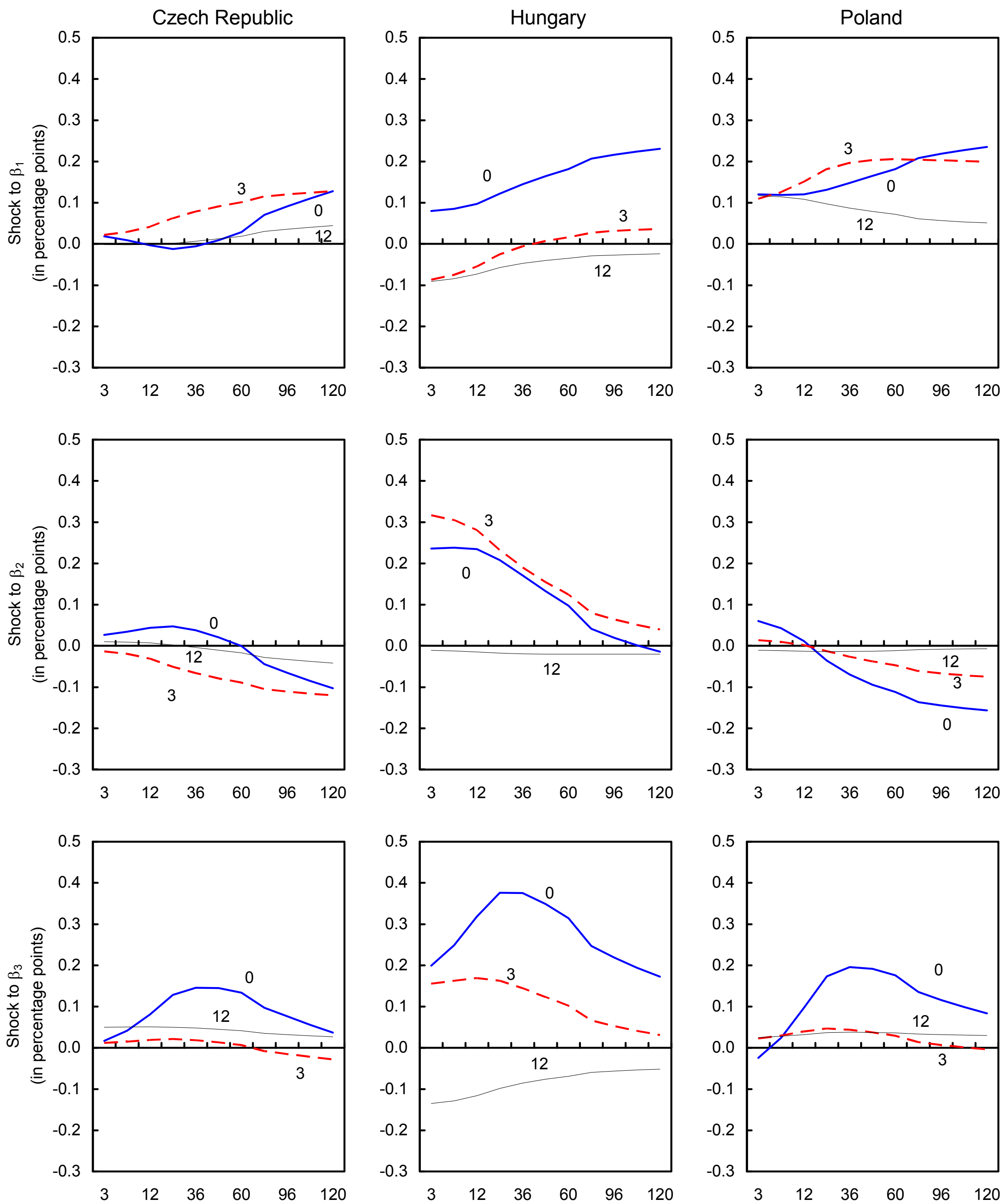
- β_1 **(level) shocks** are associated with an increase in actual inflation and domestic policy rates in Poland and, to a lesser extent, in the Czech Republic. The responses are the reverse in Hungary. The REER tends to appreciate (quickly) in the Czech Republic and to depreciate in Poland and Hungary.

- β_2 **(slope) shocks** are associated with domestic interest rate hikes—quickly reversed in the Czech Republic—and a decrease in economic activity in the short run in all CEE countries (Figure 5). As noted by Diebold, Rudebusch, and Arouba (2006), this relationship can be interpreted in two ways. First, the monetary authority may be reacting to yields. Second, given the gap between the announcement of the macroeconomic news and the policy rate decision, the market yields rise in anticipation of the policy rate increase following the release of macroeconomic data. The negative impact of a yield curve inversion on economic activity has been well documented for both industrialized (Estrella and Mishkin, 1997) and emerging market countries (Mehl, 2006). The REER tends to appreciate in those countries where the short end of the yield curve increases—in Hungary and Poland, and to depreciate where the effect on the short end is smaller, namely, in the Czech Republic.

- β_3 **(curvature)** ***shocks*** are associated with a rise in policy rates in all CEE countries (Figure 7), possibly reflecting the uncertainty of future tightening actions. There is no clear pattern in the response of the other macro variables, suggesting that changes in curvature are purely financial market events.

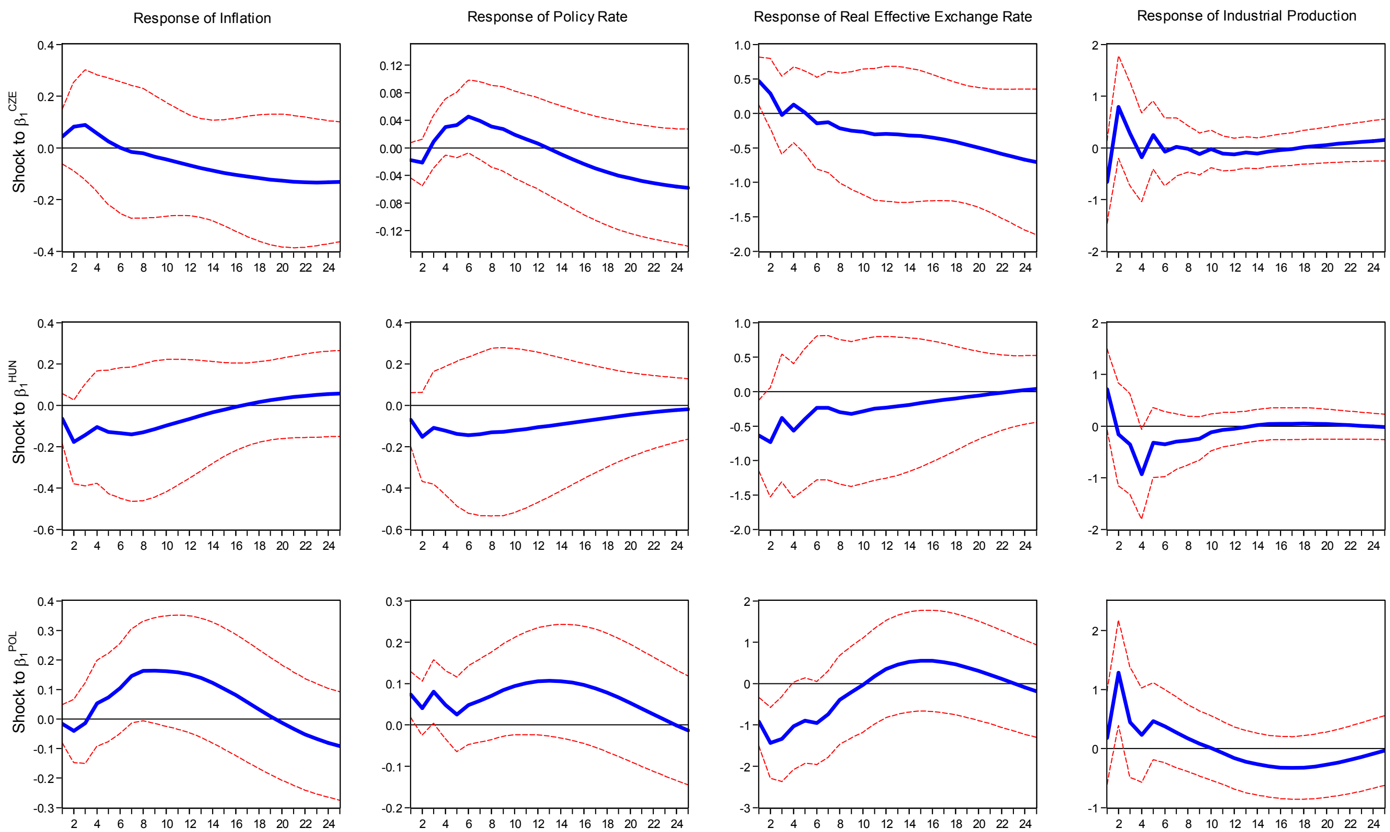
[16] Diebold, Piazzesi and Rudebusch (2005) note that for the U.S. the influences from macro variables to the yield curve are stronger than the ones from yield curves to macro variables.

Figure 4. Central European Countries: Impulse Response of Yield Curves to Own Yield Factor Shocks, 2000–08



Source: Authors' calculations.

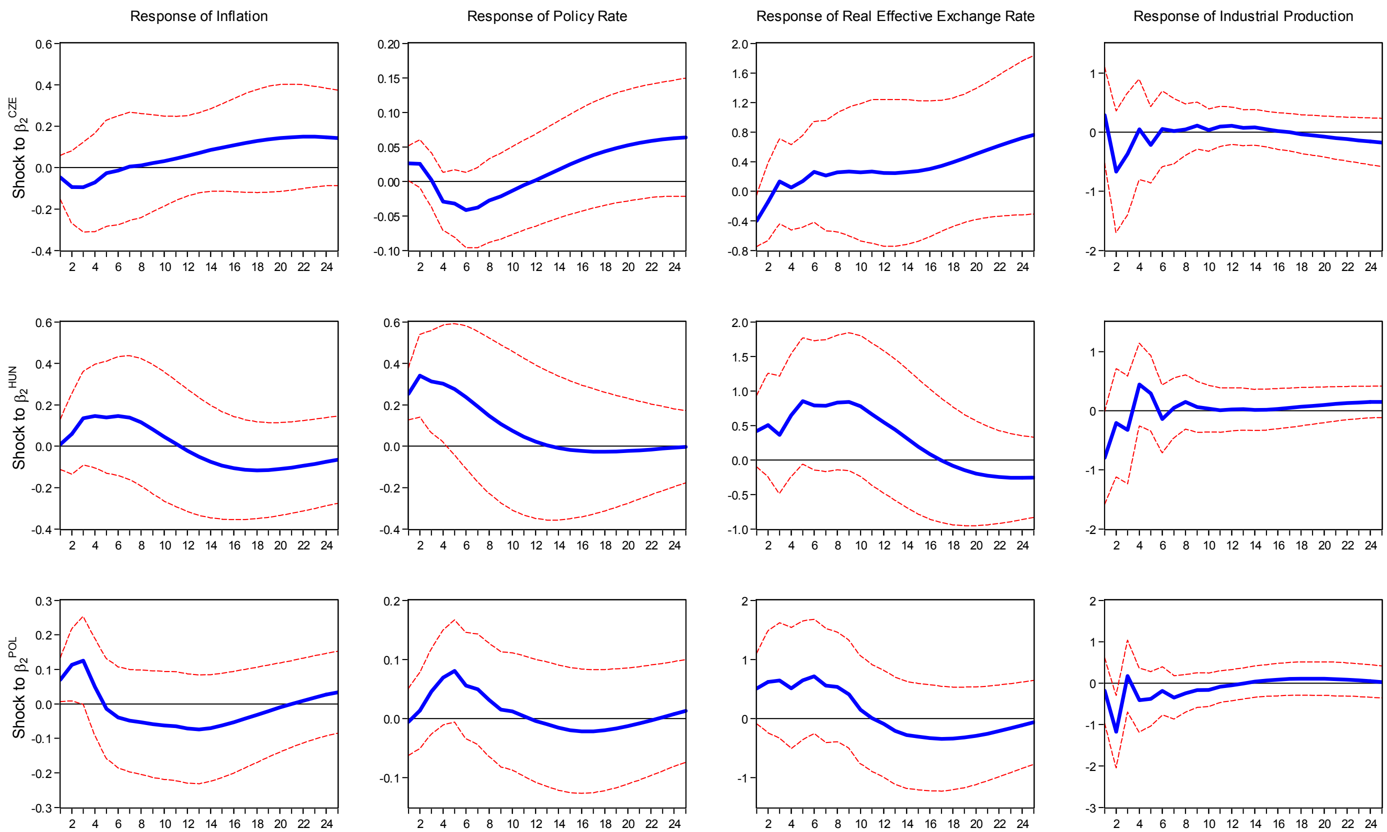
Figure 5. Response of Macroeconomic Variables to Shocks to β_1, 2000–08
(in percentage points)



Source: Authors' calculations.

Based on three country VAR models with x = [i*, β_1, β_2, β_3, IP, REER, π, i^p] and L = 3.

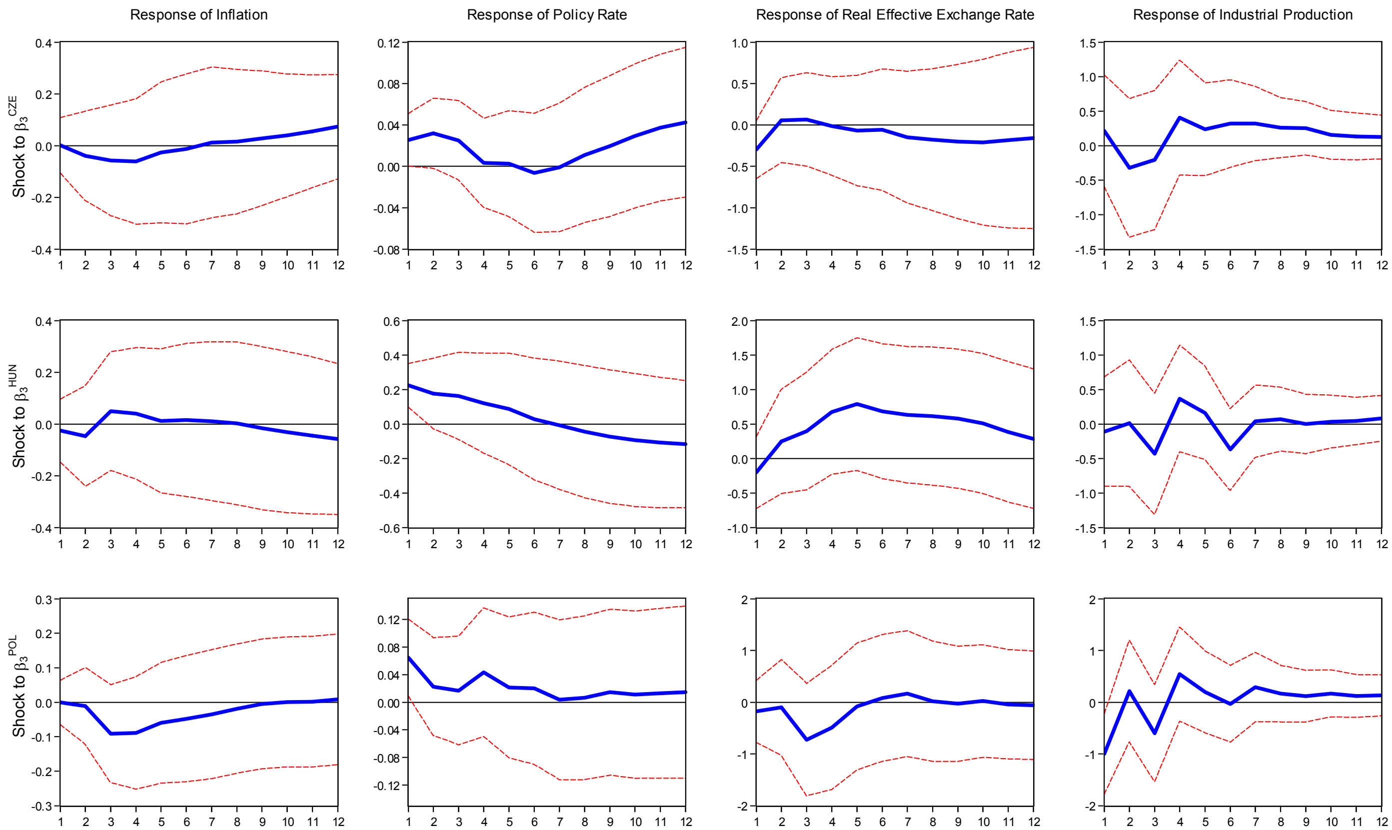
Figure 6. Response of Macroeconomic Variables to Shocks to β_2, 2000–08
(in percentage points)



Source: Authors' calculations.

Based on three country VAR models with x = [i*, β_1, β_2, β_3, IP, REER, π, i^p] and L = 3.

Figure 7. Response of Macroeconomic Variables to Shocks to β_3, 2000–08
(in percentage points)



Source: Authors' calculations.

Based on three country VAR models with x = [i*, β_1, β_2, β_3, IP, REER, π, i^p] and L = 3.

C. Cross-Border Spillovers

This subsection deals with spillover effects from the euro area on the CEE countries and across the region. As in the previous section, the empirical evidence stems from GVAR analysis. However, and in light of the high degree of correlation with the euro area yield curve factors (Figure 7), this analysis has been supplemented by a prior examination of regional and global components of the yield curve using principal components.

Regional and Global Components of the Yield Curve

Principal components. The evidence from principle component analysis suggests that roughly two-thirds of the variance in β_{1t} and at least half of the variance in β_{2t} and β_{3t} for the Czech Republic, Hungary, and Poland can be associated with a regional component (Table 4, top panel). While this points to a substantial common regional component for long-run inflation expectations (β_{1t}), the importance of this component is smaller than the share of variation explained by the global level factor in the G-3 countries (75 percent; see Table 4, bottom panel). It is also less than that for the U.S., Germany, Japan, and the U.K., (Diebold, Li, and Yue, 2007), where over 90 percent of the level factor variation has been associated with the first principal component. In contrast, principle component analysis suggests that the regional component for the slope (β_{2t}) and curvature (β_{3t}) of the yield curve in CEE countries align closely with those in Diebold, Li, and Yue (2007) but a bit lower than the G-3 countries.

Although it is tempting to associate the somewhat smaller explanatory power of the regional component as evidence of less financial integration in the CEE region than in global bond markets, there are at least two reasons to be cautious with such an inference. First, the CEE countries have higher volatility in the local (idiosyncratic) components. Second, although the results for the CEE countries suggest that the regional component has not increased following EU accession, a similar decline in the more recent subsample is apparent in the components of the G-3 yield factors.

More accurate inferences about the degree of spillovers and integration can be obtained from direct correlations between the NS factors of the CEE countries and those of the euro area or a global factor. In particular, differences among the CEE countries emerge when examining the correlation of the $\beta_{i,t}$'s with the euro area or global component counterparts. Consider the level factor (β_{1t}). Correlations of the level factor of the Czech Republic and Poland with that of the euro area (or the global component) are quite high and statistically significant, and they are substantially higher than that of Hungary (Table 5).[17] However, as a result of increased integration or a better convergence of policies, Hungary's correlation has increased substantially (and become statistically significant) since joining the EU (though it still remains about half as strong as that in the other two countries).

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[17] Also, the higher correlation of Poland with the global factor is likely to have been influenced by the larger role of non-European investors in Poland.

Differences are sharper when examining the correlation of the slope (β_{2t}) and curvature (β_{3t}) of the yield curve. Specifically, in the Czech Republic, the correlation with the euro or the global factor has increased since EU accession, and has been consistently larger than in Hungary and Poland. In contrast, the correlations of β_{2t} and β_{3t} in the latter countries have turned negative since joining the EU. This reflects, inter alia, the disinversion of the yield curves in Hungary and Poland, coupled with the European Central Bank's (ECB) tightening since 2004.

Figure 8. CEEC Yield Curve Factors and the Euro Area Yield Curve Factors and Correlations Over Rolling 3-Year Horizon, 2000–08

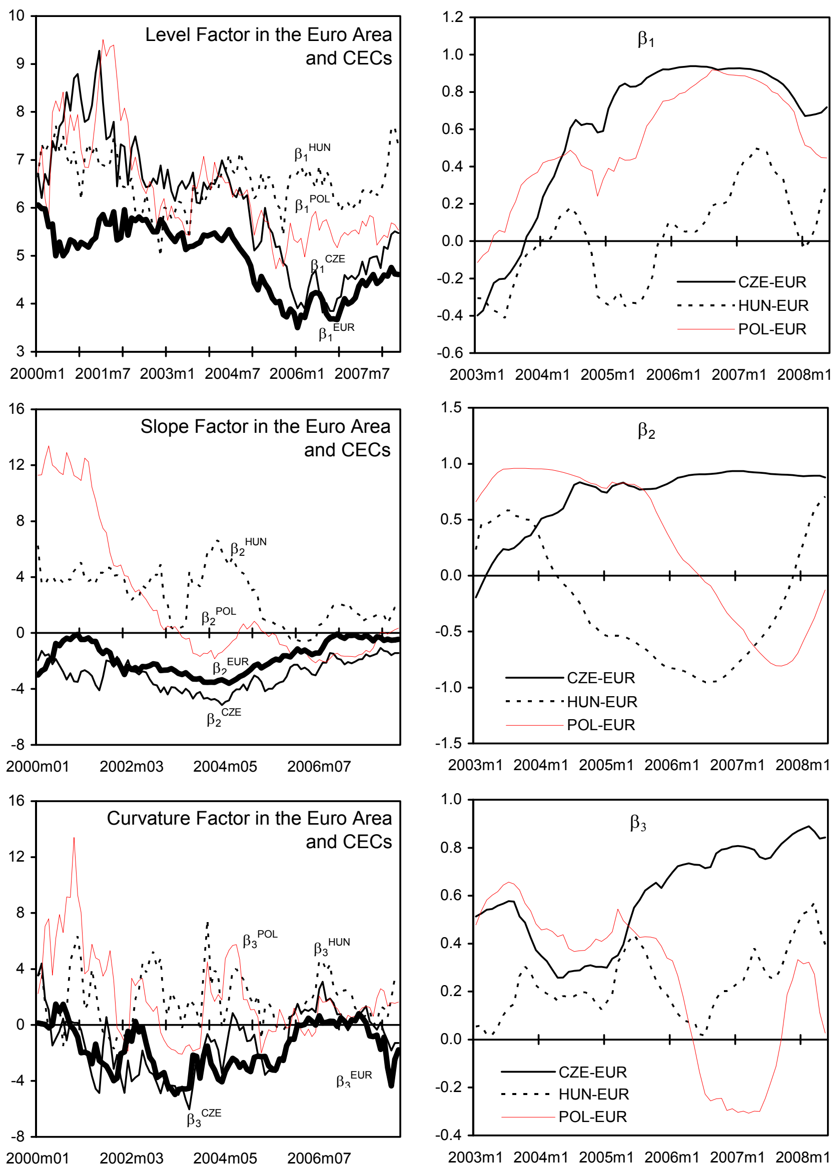


Table 4. Principal Components of Yield Factors

Eigenvalues				Eigenvalues			
Number	Value	Proportion	Cumulative Proportion	Number	Value	Proportion	Cumulative Proportion
2000-08				2004-08			
Panel A. CEC-3							
β1							
1	2.03	0.68	0.68	1	1.99	0.66	0.66
2	0.78	0.26	0.94	2	0.66	0.22	0.88
3	0.19	0.06	1.00	3	0.35	0.12	1.00
β2							
1	1.49	0.50	0.50	1	1.77	0.59	0.59
2	1.18	0.39	0.89	2	0.71	0.24	0.83
3	0.33	0.11	1.00	3	0.51	0.17	1.00
β3							
1	1.56	0.52	0.52	1	1.76	0.59	0.59
2	0.78	0.26	0.78	2	0.81	0.27	0.86
3	0.66	0.22	1.00	3	0.43	0.14	1.00
Panel B. G-3							
β1							
1	2.24	0.75	0.75	1	2.24	0.75	0.75
2	0.67	0.22	0.97	2	0.47	0.16	0.90
3	0.09	0.03	1.00	3	0.29	0.10	1.00
β2							
1	2.11	0.70	0.70	1	1.89	0.63	0.63
2	0.71	0.24	0.94	2	1.01	0.34	0.97
3	0.18	0.06	1.00	3	0.10	0.03	1.00
β3							
1	2.02	0.67	0.67	1	1.84	0.61	0.61
2	0.82	0.27	0.95	2	1.07	0.36	0.97
3	0.15	0.05	1.00	3	0.09	0.03	1.00

Source: Authors' calculations.

Table 5. Correlation of the Yield Factors Between CEC-3 and International Factors [3/]

	CZE	HUN	POL	CZE	HUN	POL	CZE	HUN	POL
	2000-08			2000-04			2004-08		
					β1				
EUR 1/	0.83 (***)	0.11	0.67(***)	-0.12	-.09	0.11	0.89 (***)	0.45 (***)	0.79 (***)
Global Factor 2/	0.80 (***)	0.21(**)	0.73(***)	0.06	0.23 (*)	0.36 (***)	0.82 (***)	0.44 (***)	0.86 (***)
					β2				
EUR	0.71 (***)	-0.38(***)	0.20(**)	0.44 (***)	0.05	0.82 (***)	0.93 (***)	-0.53	-0.44 (***)
Global Factor	0.66 (***)	-0.51(***)	0.07	0.54 (***)	0.17	0.93 (***)	0.97 (***)	-0.61 (***)	-0.54 (***)
					β3				
EUR	0.68(***)	0.11	0.31(***)	0.61 (***)	0.11	0.65 (***)	0.76 (***)	0.16	-0.28 (*)
Global Factor	0.71(***)	-0.01	0.23(**)	0.65 (***)	-0.02	0.68 (***)	0.69 (***)	0.08	-0.38 (***)

Source: Authors' calculations.

1/ Represents yield factors for the Euro area yield curve.

2/ Global Factor represents the first principal component of the yield factors of Germany, Japan and the USA.

3/ (***): p-value of less than 0.01;
(**): p-value of less than 0.05;
(*): p-value less than 0.10

VAR Model

As noted above, spillovers and linkages of the yield curves in the three CEE countries and the euro area can be studied also with a VAR model comprising the three NS parameters $\beta_{i,t}$'s for each of the CEE countries plus the three $\beta_{i,t}$'s of the euro area.[18] Reproducing the VAR specification here for convenience gives rise to the following expressions

$$\tilde{X}_t = \tilde{C}(L)^{-1}\tilde{\mu}_t$$

$$\tilde{X}_t = \left[\beta_{1,t}^{euro}, \beta_{2,t}^{euro}, \beta_{3,t}^{euro}, \beta_{1,t}^{CZE}, \beta_{2,t}^{CZE}, \beta_{3,t}^{CZE}, \beta_{1,t}^{HUN}, \beta_{2,t}^{HUN}, \beta_{3,t}^{HUN}, \beta_{1,t}^{POL}, \beta_{2,t}^{POL}, \beta_{3,t}^{POL}\right]'.$$

As above, the discussion that follows centers on the yield curve's generalized impulse response (GIR) from a model with L=3.

The historical correlations among the reduced-form residuals ($\tilde{\Omega} = E[\tilde{\mu} \times \tilde{\mu}']$) provide insights into the contemporaneous (but unexpected) spillover effects across different countries' yield curves:[19]

- Shocks to the slope of the euro area yield curve (β_2^{euro}) are negatively correlated with shocks to the area level factor (β_1^{euro}), suggesting that an inversion of the yield curve is associated with a fall in long-run inflation expectations in the euro area (Table 6). Also, a flattening of the euro curve (a positive shock to β_2^{euro}) is negatively correlated with the level factor (β_1^i) in the CEE countries, suggesting a positive spillover in the form of lower long-run inflation expectations in the CEE countries. Moreover, these correlations become larger in the second half of the sample, suggesting further nominal interest rate convergence as the euro area yield curve flattened in the post-accession period.

- At the CEE country level, shocks to the curve slopes β_2^i are negatively correlated with the level factors β_1^i, reflecting a fall in long-run inflation expectations in that

[18] The results from this large VAR model have been found to be qualitatively similar to those obtained from a series of smaller two-country models that included the $\beta_{i,t}$'s for two CEE countries at a time, or one CEE country and the euro area.

[19] Besides the correlation in the euro area, the largest correlations have been found to be between curvature and slope in the CEE countries; there are also some strong correlations between the curvature in some countries and the slope in another country.

country (although this correlation has been weaker in Hungary). Across countries, the β_2^i are negatively correlated with the β_1^j such that a tightening in one of the CEE countries lowers the long-run inflation expectations in the others—further strengthening the argument for spillovers.

Table 6. Residual Variance-Correlation Matrix, 2000–08

		Euro			Czech Republic			Hungary			Poland		
Panel A. 2000-08		b_1^{euro}	b_2^{euro}	b_3^{euro}	b_1^{cze}	b_2^{cze}	b_3^{cze}	b_1^{hun}	b_2^{hun}	b_3^{hun}	b_1^{pol}	b_2^{pol}	b_3^{pol}
Euro	b_1^{euro}	0.04											
	b_2^{euro}	-0.93	0.06										
	b_3^{euro}	-0.17	0.16	0.52									
Czech	b_1^{cze}	0.64	-0.59	-0.07	0.12								
Republic	b_2^{cze}	-0.55	0.56	0.19	-0.87	0.12							
	b_3^{cze}	-0.23	0.20	0.49	-0.64	0.63	0.61						
Hungary	b_1^{hun}	0.19	-0.10	-0.06	0.06	-0.08	0.07	0.11					
	b_2^{hun}	0.14	-0.16	0.04	0.33	-0.24	-0.20	-0.41	0.44				
	b_3^{hun}	0.01	0.01	0.24	0.08	-0.02	0.12	-0.30	0.47	1.49			
Poland	b_1^{pol}	0.49	-0.48	-0.05	0.25	-0.21	0.02	0.23	-0.08	0.08	0.14		
	b_2^{pol}	-0.44	0.43	0.30	-0.17	0.12	0.07	0.06	-0.06	-0.13	-0.70	0.20	
	b_3^{pol}	-0.02	0.16	0.08	0.07	0.04	0.13	0.00	0.34	0.36	-0.46	0.18	1.16
Panel B. 2004-08		b_1^{euro}	b_2^{euro}	b_3^{euro}	b_1^{cze}	b_2^{cze}	b_3^{cze}	b_1^{hun}	b_2^{hun}	b_3^{hun}	b_1^{pol}	b_2^{pol}	b_3^{pol}
Euro	b_1^{euro}	0.02											
	b_2^{euro}	-0.95	0.03										
	b_3^{euro}	0.22	-0.38	0.26									
Czech	b_1^{cze}	0.67	-0.63	0.07	0.04								
Republic	b_2^{cze}	-0.38	0.51	-0.21	-0.78	0.04							
	b_3^{cze}	0.02	-0.09	0.56	-0.59	0.55	0.31						
Hungary	b_1^{hun}	0.57	-0.50	-0.14	0.33	-0.25	-0.06	0.08					
	b_2^{hun}	-0.53	0.45	-0.46	-0.36	0.17	-0.15	-0.46	0.09				
	b_3^{hun}	-0.32	0.23	-0.34	-0.28	0.18	0.05	-0.15	0.85	0.56			
Poland	b_1^{pol}	0.57	-0.56	0.30	0.44	-0.26	0.21	0.50	-0.42	0.02	0.05		
	b_2^{pol}	0.02	0.00	-0.12	0.13	-0.03	-0.32	-0.31	0.27	-0.06	-0.70	0.02	
	b_3^{pol}	-0.52	0.59	-0.67	-0.52	0.57	-0.18	-0.30	0.56	0.33	-0.72	0.47	0.29

Source: Authors' calculations.

Impulse Responses

The generalized impulse responses complement the purely static discussion of $\tilde{\Omega}$ above by characterizing the dynamic spillover effects and linkages. In what follows, shocks associated with $\beta_{i,t}^{euro}$'s are presented first, followed by the generalized impulse responses for shocks originating within the region.

Euro Area Shocks

Once again, the generalized impulse responses for the Czech Republic and Poland share several traits, while those of Hungary are quite different (Figures 8a-c). These differences, however, vanish to a large extent in 2004–08. Specifically, we find the following:

- An increase in the level of the euro area yield curve ($\beta_{1,t}^{euro}$) is gradually transmitted to the CEE countries, and results in a steepening of the yield curve in the Czech Republic and Poland on impact (that is, in the first month after the shock). The pattern becomes clearer in the second subsample, when Hungary's curve response is similar to the other two countries. While the effect of these shocks die out in the euro area after three months, these take longer to die out in the CEE countries (suggesting that it takes longer to re-anchor inflation expectations in the latter group of countries).

- A flattening/tilting of the euro area curve *(* $\beta_{2,t}^{euro}$ *)* has the same effect in the CEE countries (again, only in the second subsample in the case of Hungary), and the response appears to be somewhat more persistent (i.e., beyond the three month period) in the later countries.

- $\beta_{3,t}^{euro}$ shocks have small effects on the yield curves in CEE countries because these shocks have only a minor effect on the yield curve in the euro area.

CEE Country Shocks

Below, the discussion of the generalized impulse response of the yield curve is organized by shocks in the $\beta_{i,t}$'s for each country:

- A shock to β_1^{CZE} tilts the yield curve in the Czech Republic and has a similar effect in the other countries (especially in the second half of the sample; see Figure 9a. It is noteworthy that by six months the effect on the long end of the Czech Republic's yield curve vanishes but persists longer in Poland and, to a lesser extent, in Hungary. A shock to β_2^{CZE} inverts the yield curve in the Czech Republic and shifts down the yield curves in Poland and Hungary (second subsample; see Figure 9b), confirming the strong spillover effect across the CEE countries.

- A shock to the level of the curve in Hungary (β_1^{HUN}) shifts up the yield curves in Poland and, to a lesser extent, in the Czech Republic (Figure 10a). A shock to

β_2^{HUN} inverts the yield curve in Hungary, while producing only small responses in the curves of the Czech Republic and Poland—which are shifted down after accession to the EU (Figure 10-b).

- A shock to level of the yield curve in Poland (β_1^{POL}) is transmitted to the other countries—especially in the short run (first three months; Figure 11-a). The same pattern emerges for shocks to the slope (β_2^{POL}), suggesting stronger spillovers from Poland to the other two CEE countries (Figure 11b).

Despite the important role played by global and regional factors in the CEE country's yield curve dynamics (as shown with principal component and simple correlations), they rarely account for more than 40 percent of the variance of yield curve factors—as reflected in the variance decompositions of Table 7 (top panel). This is much less than in Diebold, Li, and Yue (2007) who find that global factors explain more of the variance of yields in the U.S., U.K., Germany, and Japan; however, the two methodologies are not directly comparable. For the Czech Republic and Poland, euro area yield curve shocks explain between 20 percent and 30 percent of the variance of the level and slope of their own curves. In contrast, for Hungary, regional shocks are more important than euro area-originated shocks, although the importance of the euro area shocks has increased following EU accession.

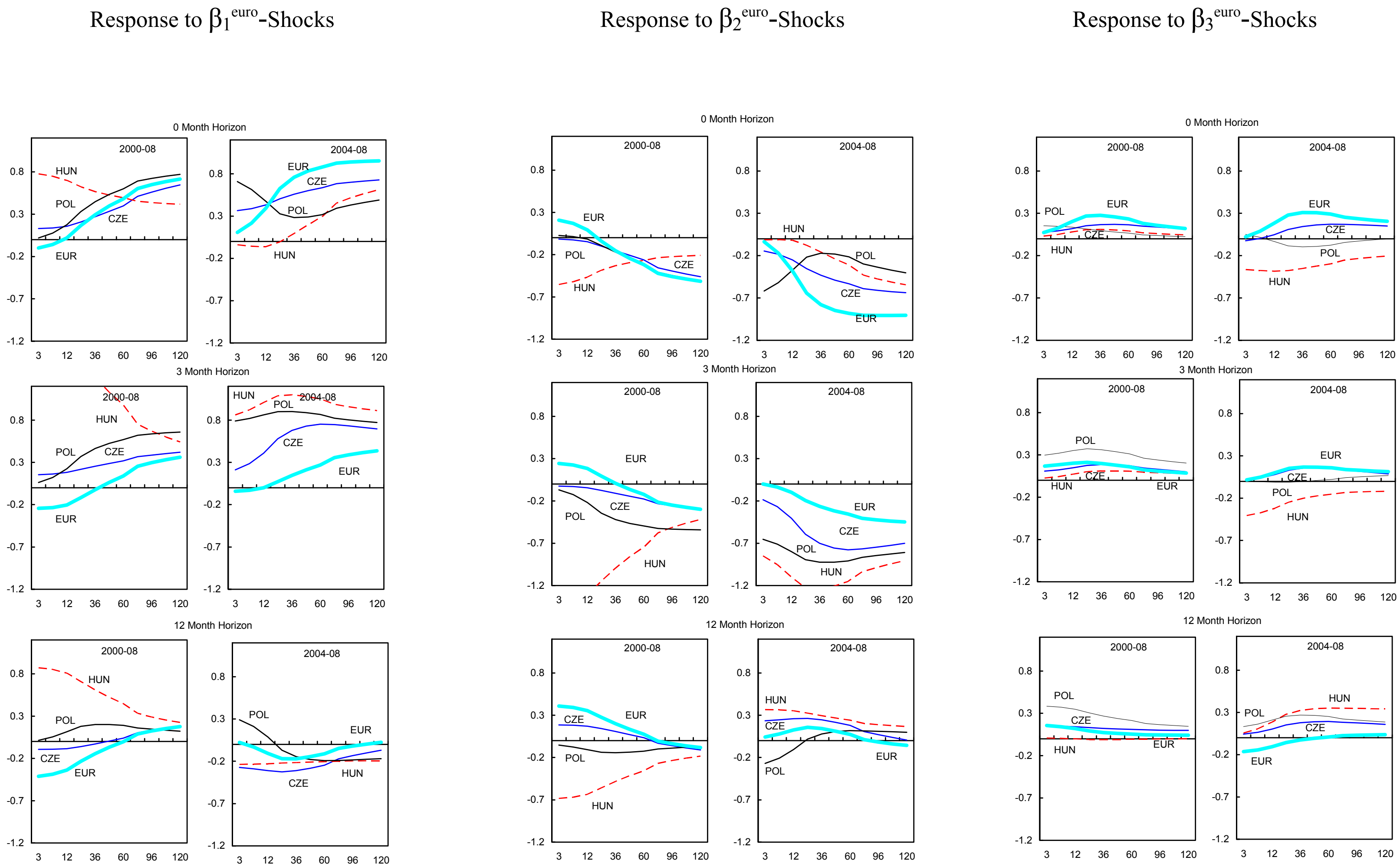
Indeed, after accession (bottom panel of Table 7), the percentage of variance in yield curves explained by regional factors increases markedly in the Czech Republic and Poland, while the importance of euro area factors grows in Hungary. Both results suggest a clear increase in the importance of global and regional factors for the CEE countries, a reflection of the increased integration and/or convergence of policies in the three countries. This is not surprising, given the greater issuance of euro-denominated bonds in these countries over the past five years—among other factors.

Table 7. Variance Decomposition

	Maturity	Euro	Region	Country	Euro	Region	Country	Euro	Region	Country
			β_1			β_2			β_3	
					Panel A: 2000-08					
Czech	0	24	7	69	22	4	73	15	4	81
Republic	3	21	9	70	19	7	74	29	15	56
	12	23	20	58	20	26	54	26	33	41
	60	26	34	39	30	30	40	28	43	29
Hungary	0	4	5	91	3	19	79	4	11	85
	3	10	16	74	14	34	52	12	30	58
	12	7	33	60	26	41	33	15	37	48
	60	9	38	53	28	45	27	17	39	44
Poland	0	20	7	72	23	4	74	2	17	81
	3	28	19	53	21	11	68	12	29	59
	12	28	35	37	21	14	65	14	38	48
	60	26	35	39	16	24	60	13	39	47
					Panel B: 2004-08					
Czech	0	24	22	54	16	18	66	15	10	76
Republic	3	27	34	39	23	28	49	29	42	29
	12	27	44	29	27	35	38	34	46	21
	60	26	55	18	30	41	30	24	64	12
Hungary	0	24	25	51	21	22	57	12	10	78
	3	31	29	41	27	20	53	24	28	49
	12	20	33	47	34	17	49	26	33	41
	60	23	28	50	25	31	44	23	30	46
Poland	0	21	21	58	1	15	85	27	29	44
	3	34	27	39	9	36	55	25	44	31
	12	28	40	32	26	58	16	28	50	22
	60	26	54	20	26	56	18	30	52	18

Source: Authors' calculations.

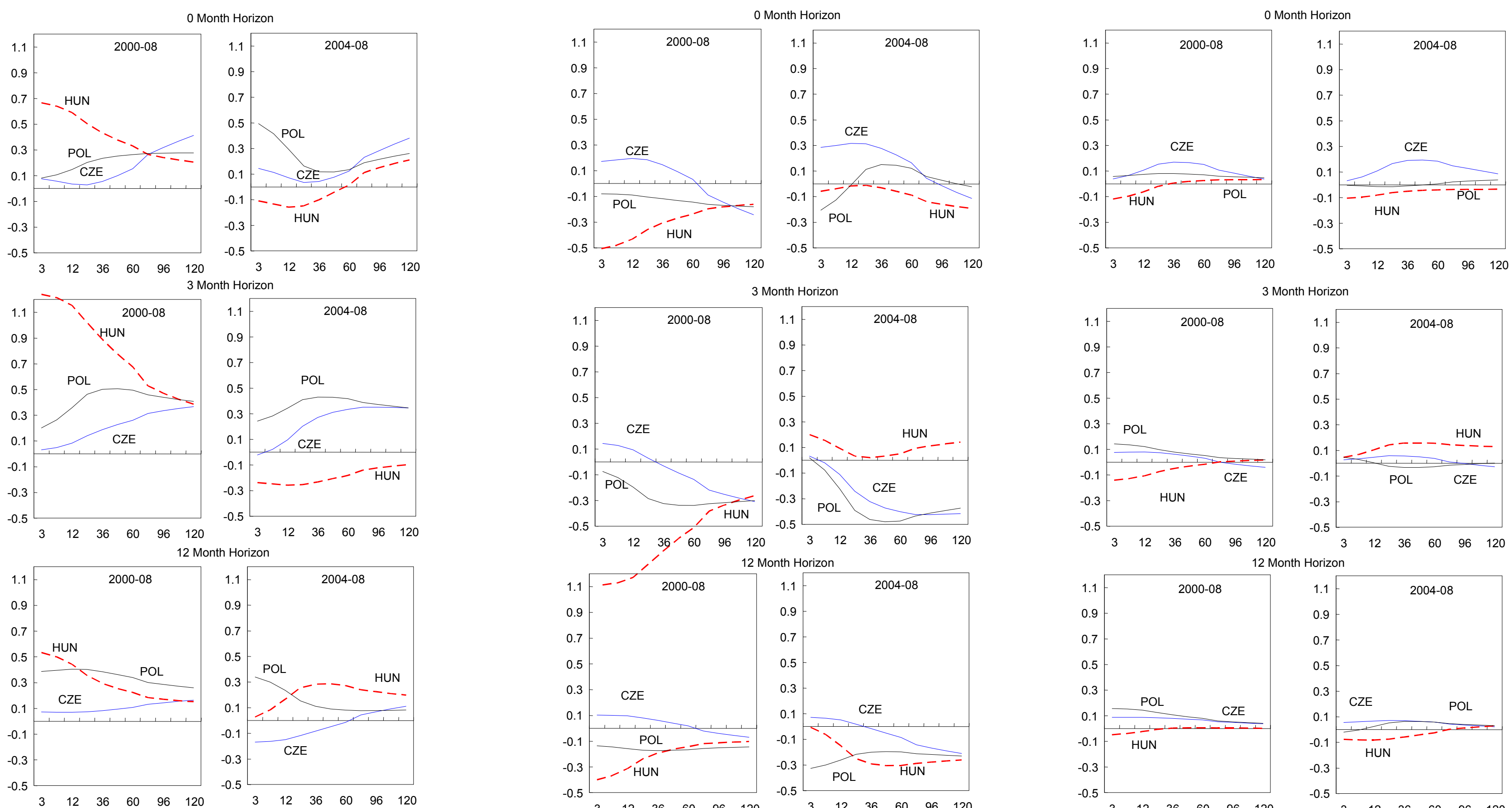
Figure 9. Impulse Response of Yield Curves to β_i^{euro}-Shocks



Source: Authors' calculations.

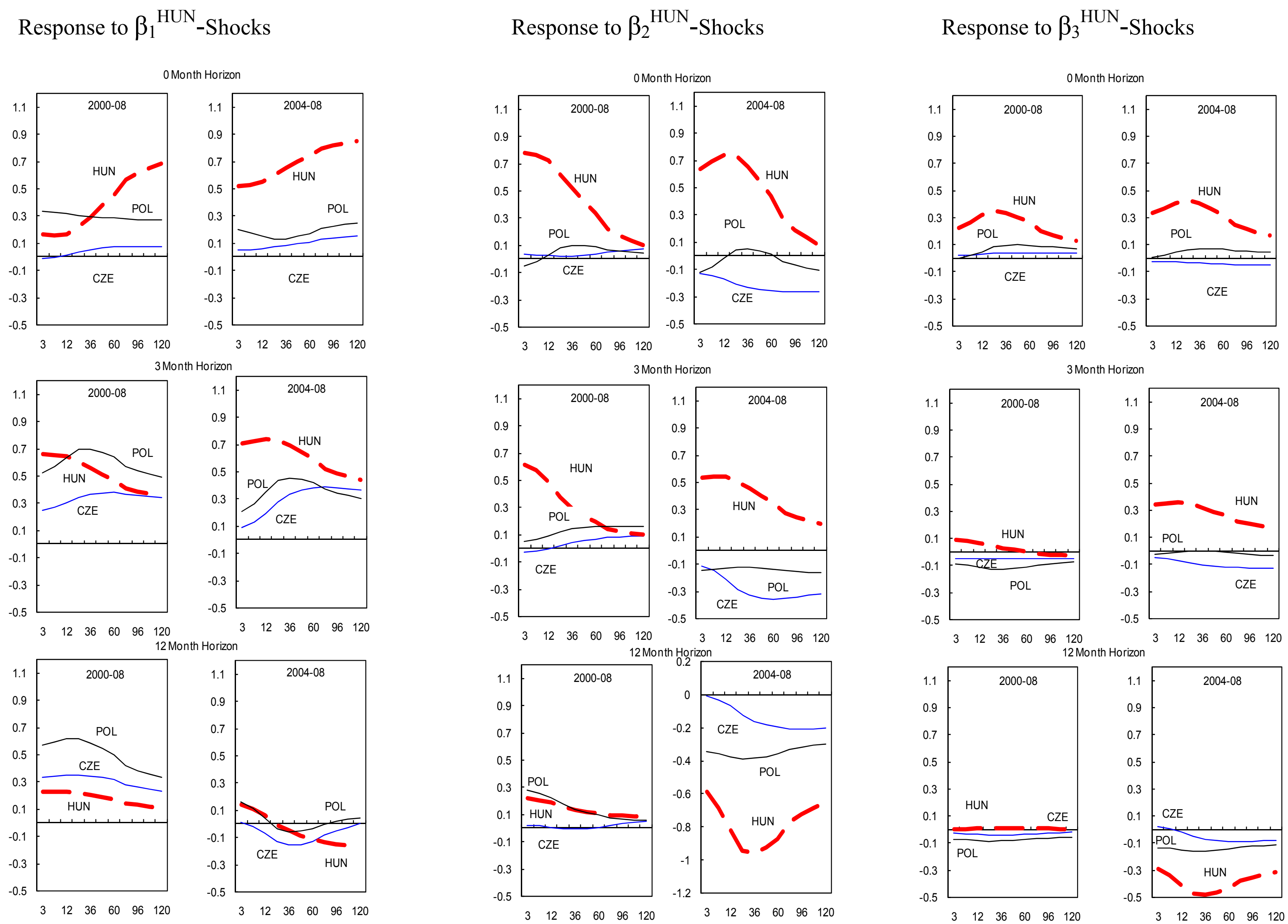
Figure 10. Impulse Response of Yield Curves to β_i^{CZE}-Shocks

Response to β_1^{CZE}-Shocks | Response to β_2^{CZE}-Shocks | Response to β_3^{CZE}-Shocks



Source: Authors' calculations.

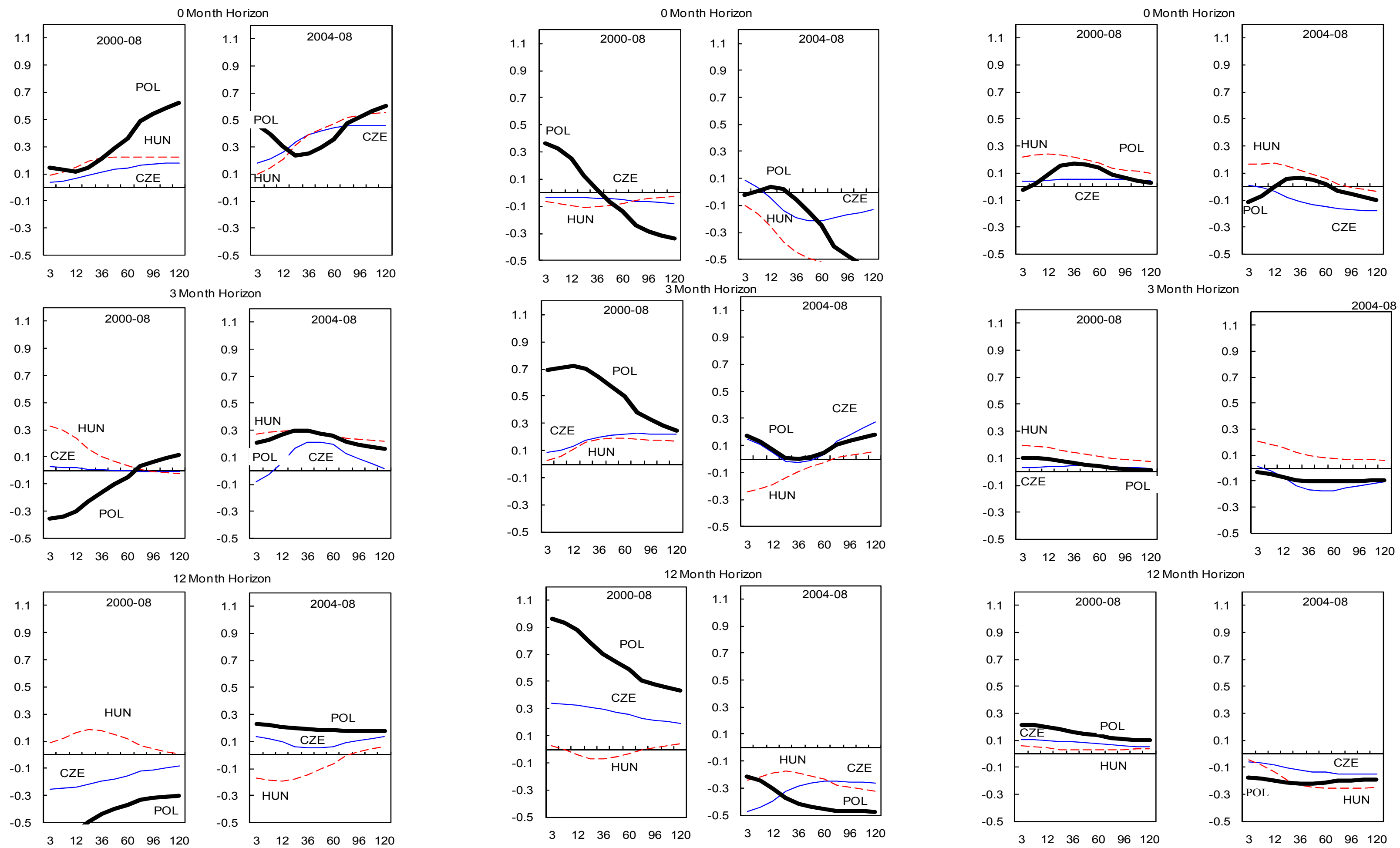
Figure 11. Impulse Response of Yield Curves to β_i^{HUN}-Shocks



Source: Authors' calculations.

Figure 12. Impulse Response of Yield Curves to β_i^{POL}-Shocks

Response to β_1^{POL}-Shocks | Response to β_2^{POL}-Shocks | Response to β_3^{POL}-Shocks



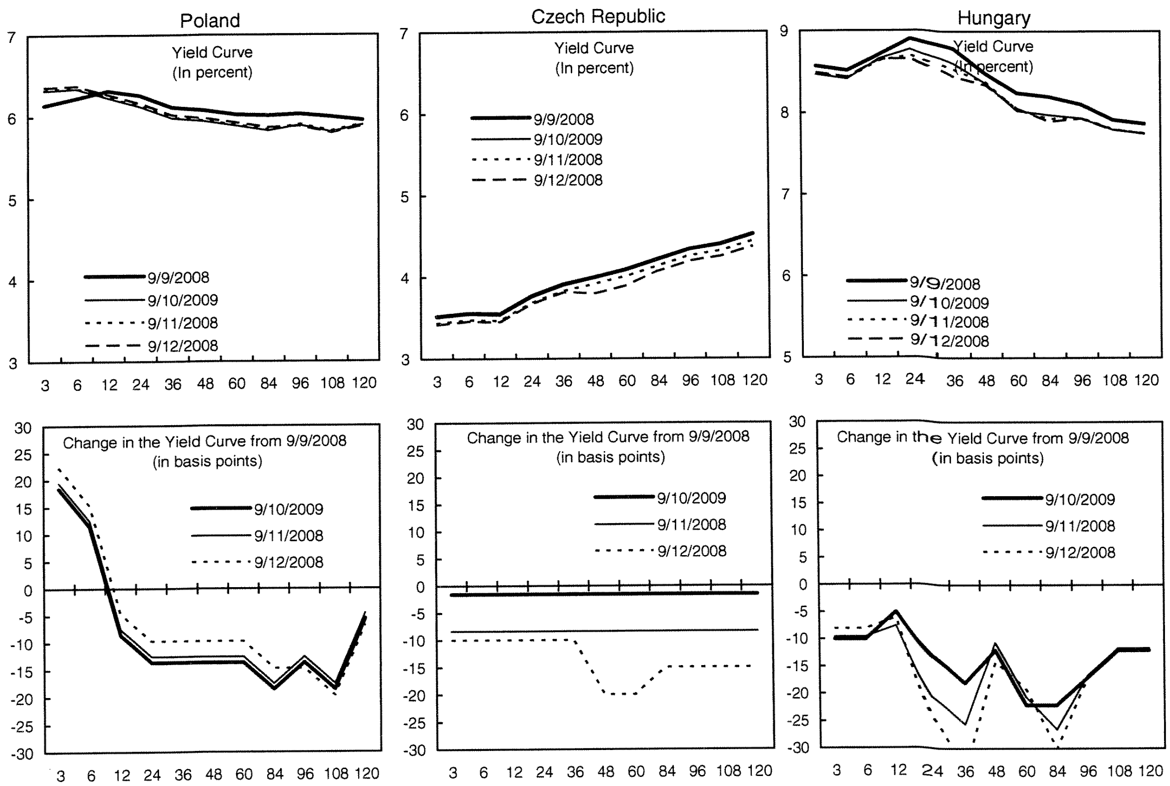
Source: Authors' calculations.

V. Event Study: Poland's Euro Adoption Announcement

To illustrate how the above empirical framework can be used to understand the potential macro implications, including potential spillover effect, of financial market developments, we study the impact of Poland's announcement to adopt the euro just before the intensification of the financial crisis following the closure of Lehman Brothers in mid-September, 2008. This is done in two-stages. First, the shocks to the NS factors implied in the move of yield curves three days after the event are extracted. Second, the estimated shocks to the NS factors in the estimated VARs are used to infer the potential implications for macro variables and potential spillover effects.

Poland's euro adoption announcement. On Wednesday September 10, 2008 Prime Minister Tusk surprised markets by announcing the goal of adopting the euro in 2011. Shortly thereafter, Finance Minister Rostowski was quoted as saying the "Prime Minister adequately outlined the government's aspirations...this is how his statement should be understood." And deputy Finance Minister Zajdel-Kurowska added "Entering ERM2 in 2009 is realistic... if we proceed immediately." There was no official NBP comment then, but MPC member Wojtyna noted that "If the PM confirms the 2011 date... then this factor will gain importance in the next MPC decision." Subsequently, the goal was clarified as obtaining the Commission's approval to adopt the euro in 2011, even though euro notes would not be in circulation until January 1, 2012. In the event, Poland's yield curve steepened on September 10 consistent with expectations of a tighter monetary policy and lower long-run inflation (Figure 12); in the Czech Republic and Hungary the yield curves shifted down, in the former country with a one-day delay.

Figure 13. Central and Eastern European Yield Curves and their Movements
September 9–12, 2008



Source: Bloomberg.

To back out the NS β_i shocks implicit in the observed movements in the yield curves immediately following the event, we take advantage of the simplicity of the NS representation. The yield curve can be expressed as a general function of the NS $\beta_i^{(j)}$ and a vector x of all other variables in the empirical models discussed in this paper. In this regard, consider a general specification yield curve and its total derivative:

$$y_i(\tau) = f\left(\beta_1^{(i)}, \beta_2^{(i)}, \beta_3^{(i)}, x^{(i)}\right),$$

$$dy_i(\tau)\,\big|_{dx^{(i)}=0} = \frac{\partial f(.)}{\partial \beta_1^{(i)}} d\beta_1^{(i)} + \frac{\partial f(.)}{\partial \beta_2^{(i)}} d\beta_2^{(i)} + \frac{\partial f(.)}{\partial \beta_3^{(i)}} d\beta_3^{(i)},$$

where the partial derivatives correspond to the impulse response of the yield curve, $gir_0^{\beta_1^{(j)}}$, and $d\beta_1^{(j)}$ correspond to the change in or shock to the NS $\beta_i^{(j)}$.[20]

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[20] For details, see Hamilton (1994) page 323.

In discrete time, this decomposition suggests that movements in the yield curve can be examined using the following regression equation

$$\Delta^{h-day}\left[y_i(\tau)\right] = \mu^{\beta_1^{(i)}} \times gir_0^{\beta_1^{(i)}}(\tau) + \mu^{\beta_2^{(i)}} \times gir_0^{\beta_2^{(i)}}(\tau) + \mu^{\beta_3^{(i)}} \times gir_0^{\beta_3^{(i)}}(\tau) + \nu_i(\tau)\,,$$

where $\Delta^{h-day}\left[y_i(\tau)\right]$ corresponds to *h*-day change in the yield curve in country i , with *h*=1, 2, and 3, and τ denotes the 11 maturities discussed above. $gir_0^{\beta_1^{(j)}}$ represents the generalized impulse (impact) responses of country *j*'s yield curve to $\beta_i^{(j)}$ shocks (depicted in the first row, columns 2004–08 in Figures 9–11). $\nu_i(\tau)$ denotes error terms with standard properties.

OLS estimation has been used to estimate this equation (Appendix V, Table A1). First the equation is estimated using data for only Poland—the episode-originating country—(11 observations, the first three columns) and then estimated with data for all three countries (panel data, 33 observations, the last three columns). These regressions build on the results from the cross-border spillover VAR model (Section IV.C).[21]

In the second stage, armed with the estimated Poland shocks, the macro-financial models (Section IV.B) are used to obtain the expected macroeconomic effects. For Poland, these shocks are used directly to obtain the macro impulses responses. For the macro spillover effects on Czech Republic and Hungary, these shocks were first translated into the corresponding country shocks using the cross-border spillover VAR model, which in turn were used in the macro-financial models of the Czech Republic and Hungary.

NS β_i Shocks

Results from the first stage suggest that the observed yield curve movements have been associated with multiple NS β_i shocks. Specifically, results suggest that markets reacted to Poland's euro adoption announcement anticipating a tightening in monetary policy, with some effect on long run inflation; however, interest rates would be lowered after a couple of years, presumable to achieve nominal convergence. The yield curve movement in Poland can be adequately understood by its own shocks. Specifically, limiting the sample to Poland, the estimated coefficients (Table A1 column 4) point to a small (statistically insignificant) reduction in β_1 (long-run inflation expectations), a 40 to 50 bps increase in β_2 (the short end of the yield curve or monetary tightening), and decline of almost 200 bps in β_3 (the curvature). These shocks explained about 80 percent of the movement in Poland's yield curve following the announcement. Extending the sample to all countries rendered

[21] The discussion focuses on the impact effect to isolate the euro adoption announcement from subsequent global shocks starting with the collapse of Lehman Brothers over the weekend of September 13–14.

qualitatively similar results (column 8) even though the coefficient estimates were smaller (around 20 basis points); the model explains far less of the movement in the yield curves of the three countries.

Potential Macro Effects

The shock scenario has been built from the NS β_i shock estimates underlying the events in Poland (Table 8). This scenario broadly summarizes the estimated shocks, which in turn are fed into the macro-financial models (Section IV.B).

Table 8. Shock Scenario for Poland's Euro Adoption Announcement
(in basis points)

	μ^{β}_1	μ^{β}_2	μ^{β}_3
Poland shocks	-10	50	-150

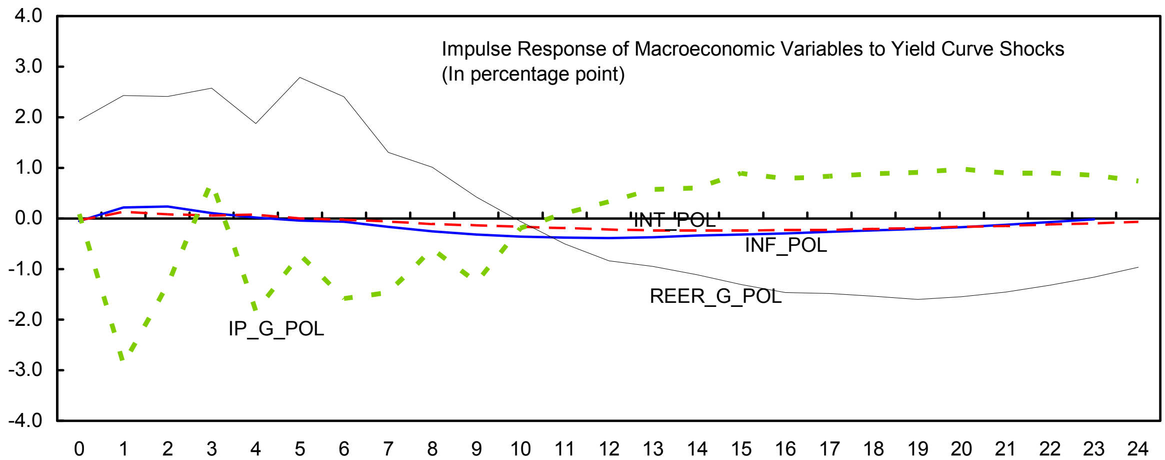
Source: Authors' calculations.

The macroeconomic impact of these shocks to the NS β_i's in Poland is shown next (Figure 13).[22] The real exchange rate appreciates over the next ten months, while the level of activity contracts. Both effects are reversed over the course of the following year. While there is also a small increase in inflation, policy interest rates do not increase much: the model suggests that the increase in (short term) market rates makes policy changes unnecessary. Of course, the resulting macro effects are only illustrative, as Lehman's bankruptcy the week after Poland's announcements would have to be modeled as an additional shock that would change all of the model's predictions.[23]

[22] The figure shows only the impact of shocks to the level and the slope of the curve: the large increase in the curvature is mostly related to portfolio reallocations of fixed income traders to reposition themselves for the new path of interest rates. The large change in curvature induces a high degree of volatility in the macro variables that seems unwarranted over the medium run. Diebold, Li, and, Yue (2007) also consider only shocks to level and slope in the context of a model with cross-country spillovers.

[23] Still, to the extent that Lehman's bankruptcy was unanticipated, the identification of the euro adoption shock and the consequent macro effects remain valid.

Figure 14. Poland: Macroeconomic Effects of Yield Curve Shocks in Response to Poland's Euro Adoption Announcement

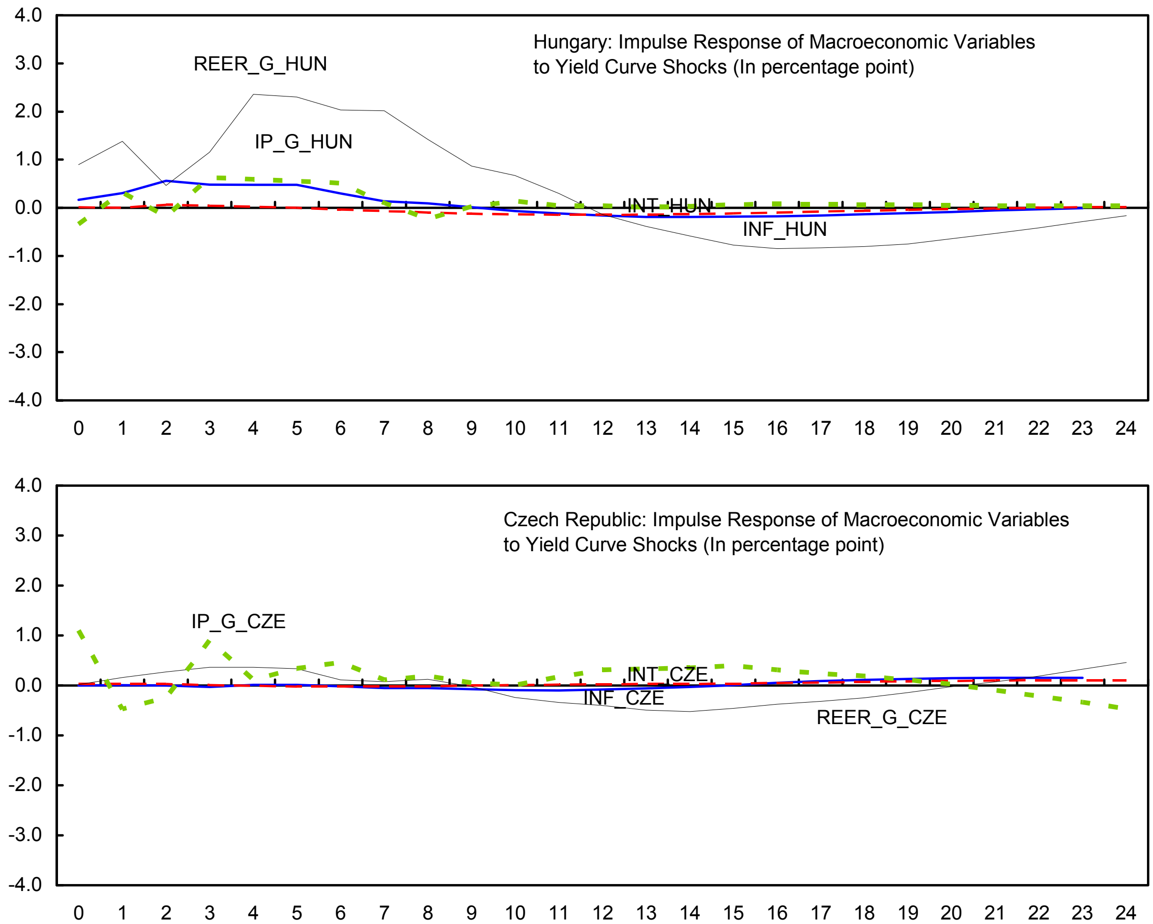


Source: Staff calculations.

Spillover Effects from Poland's Euro Adoption Announcement

Figure 12 shows that Poland's announcement also affected the yield curves in the Czech Republic and Hungary. As discussed in Section IV.C, shocks to Poland's NS β_i's are transmitted to neighboring countries' NS β_i's. The same analysis can then be applied to study regional spillover effects from Poland's announcement. The results are shown in Figure 14 where it can be seen that Hungary experiences a similar real exchange rate appreciation and inflation, but in contrast to Poland, the level of activity improves marginally. Spillovers to the Czech Republic, if any, are much smaller.

Figure 15. Czech Republic and Hungary Spillover Effects of Yield Curve Shocks on Response to Poland's Euro Adoption Announcement



Source: Staff calculations.

VI. Conclusions

Macro-financial linkages in the CEE countries are to some extent similar to those in the mature markets. In particular, increases in policy rates and in inflation shift up the whole curve, as in the mature markets. While policy rate shocks lead to an inversion of the curve over time, the inflation shocks have very little persistence in the CEE countries. This low persistence likely reflects the strong convergence trends in the CEE countries as well as the high persistence of inflation in the euro area (reinforced by the feedback effects of the output gap; see Lemke (2007).

As in the mature markets, responses of macro variables to yield curve shocks are less important than the impact of macro variables on the yield curves. For instance, the well-documented (negative) impact of a yield curve inversion on economic activity is also present in the CEE countries—albeit in a weaker fashion. Also, the increase in short-term interest rates in Hungary is associated with an exchange rate appreciation and much less of a decline in long-term inflation expectations than in the other two CEE countries—a result of Hungary's exchange rate band.

There is an important degree of co-movement between the yield curves of the euro area and those of the CEE countries, especially in the period after these countries' accession to the EU. Still, while shocks to the euro area yield curve explain 20-30 percent of the variation of the Czech Republic and Poland yield curves, the share is somewhat smaller for Hungary—though it has increased after 2004. This study has found that a tightening shock from the ECB shifts upward the curves of the three countries, and that the shifts are highly persistent. Also, a flattening of the euro curve is associated with lower inflation expectations—not just in the euro area, but also in the CEE countries. In other words, shifts in the euro area yield curve are transmitted both to interest rates and inflation expectations in the CEE countries—and the transmission is stronger post 2004.

The three CEE countries as a region have become more integrated with the euro area and the U.S., especially in the convergence of the level of their yield curves. The improved convergence, or greater bond market integration, after 2004 reflects mostly a convergence of policy frameworks in Hungary. In the post-accession period, the correlation between the slopes of the Hungary and Poland curves, on the one hand, and the euro slope on the other, have become negative, suggesting less synchronization of the respective business cycles and/or policy responses. Still, the analysis of Poland's euro adoption announcement suggests that shocks to the yield curve in one country may result in spillover effects to other countries in the region.

Appendix I. Main Developments and Monetary Policy Decisions in the CEE Countries[24]

After a protracted recession in 1997-99, the Czech Republic resumed growth in 2000. Inflation was at around 4 percent, and the yield curve was slightly upward sloping at the 5 and 6 percent level. Strong capital inflows and an appreciating koruna led the CNB to cut (two-week) repo policy rates in 2002 and 2003, to just 2 percent—leaving money market rates sometimes below the ECB policy rate. An increase in inflation expectations steepened the curve in 2004, and after a 50 bps tightening in the summer of 2004 the CNB reassessed the inflation outlook and cut rates to 1 3/4 percent in April 2005. As inflationary pressures brought headline inflation above the 3 percent target, monetary policy was gradually tightened in 2007 and 2008. A steepening of the yield curve in 2006 was then followed by a shift upwards and a mild flattening—with a current spread of around 100 bps between long and short rates.

The NB Hungary reduced policy rates steadily from 18 percent in 1998 to 8 percent in 2002. The forint yield curve shifted downwards significantly during this period, but remained inverted—a result of Hungary's strong convergence prospects. In the first half of 2003, policy rates were cut further to 7 percent, to fend-off an (appreciation) speculative attack against the strong end of the exchange rate band. During this period, the yield curve "normalized" briefly, but it became inverted again in the second-half of 2003. Increasing fiscal and external imbalances forced the NBH to increase policy rates by 600 bps (to 12.5 percent) between June and November 2003. The yield curve remained inverted during 2004, but flattened to a much lower level of around 7 percent in 2005 and 2006. In response to increasing inflationary pressures, policy rates were increased again in the second half of 2007 and early 2008, leading again to an inverted yield curve.

In Poland, after a substantial monetary policy tightening in the aftermath of the Russian crisis, the MPC reduced policy rates by around 1400 bps between early 2001 and mid-2003—when the yield curve normalized and flattened at around 5.5 percent. A new MPC increased rates in 2004 to prevent one-off inflation pressures from becoming entrenched in expectations, and then cut rates by 250 bps (to 4 percent) between March 2005 and March 2006. Following two years of strong economic growth, policy interest rates were increased seven times since April 2007 to 5 ¾ percent, with the yield curve currently displaying a small hump in the short-end—reflecting expectations of further near-term tightening.

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[24] As noted above, the data cut-off date was end-May, 2008.

Appendix II. The Generalized Vector Autoregression

The generalized vector autoregressive (GVAR) analysis of Koop, Pesaran, and Potter (1996) postulated the following vector autoregressive (VAR) model,

$$X_t = \mathrm{C}(L)^{-1}\mu_t$$

that assumes that the shocks μ are distributed multivariate normal, N(0,Ω), and thus X_t is also distributed multivariate normal, $N\left(0, \mathrm{C}(L)^{-1} \times \Omega \times \mathrm{C}(L)^{-1}{}'\right)$. Moreover, it does not impose restrictions on the estimated covariance matrix. Rather than orthogonalizing μ, GIR considers the conditional expectation of X_t given a specific shock to μ_t. Specifically, the expected (average) effect of a shock on X_t from a shock to the ith component μ_{it} can be obtained as the conditional expectation of X_t:

$$\begin{aligned} GIR(X_t, \mu_{i,t} = v) &= E\left[X_t / \mu_{i,t} = v, \Omega\right] \\ &= C(L)^{-1} \times E\left[\mu_t / \mu_{i,t} = v, \Omega\right], \end{aligned}$$

which, given the properties of the multivariate normal distribution, equals:[25]

$$GIR(X_t, \mu_{i,t} = v) = C(L)^{-1} \times \Omega_i \times \sigma_{i,i}^{-1} \times v,$$

where Ω_i and $\sigma_{i,i}$ are the i^{th} column of Ω and the variance of the i^{th} variable in X_t. Setting $v=\sigma_{i,i}^{1/2}$ is standard in GIR analysis as this corresponds to a one standard error shock. In this case, the GIR for a shock to the i^{th} variable can be readily obtained from Cholesky analysis by placing the ith variable first in the ordering. This is because in this case

$$\Omega_i \times \sigma_{i,i}^{-1} \times v = \left[\sigma_{i,i}^{1/2}, \quad \sigma_{2,i} \times \sigma_{i,i}^{-1/2}, \quad \sigma_{3,i} \times \sigma_{i,i}^{-1/2}, \cdots, \quad \sigma_{8,i} \times \sigma_{i,i}^{-1/2}\right]',$$

where right hand side of the equation can be recognized as the first column of the Cholesky factoring.[26]

[25] For details see Dhrymes, 1978, pp. 362–67.

[26] For details of the Cholesky factorization see Hamilton, 1994, pp. 91–2.

Appendix III. Estimating the Nelson Siegel Betas

A. Estimating the Nelson-Siegel Representation

The estimation strategy employed here follows that of Diebold and Li (2006) and Nelson and Siegel (1987) and, for a given λ_0, recovers time series for β_{1t}, β_{2t}, and β_{3t} for the 101 months spun by the sample period (January, 2000 through May, 2008). For each month, ordinary least squares (OLS) estimation has provided coefficient estimates using data for 11 maturities, namely, three, six, 12, 24, 36, 48, 60, 84, 96, 108, and 120 months. The estimates reported below correspond to those obtained with a value of λ_0—determined by a grid search to seven decimal places—that minimizes the sum of squared residuals for the full sample period.

It is reassuring that the NS approximation provides a reasonably good fit of the yield curve in the countries of interest (Table A1). Specifically, for the Czech Republic the mean residual ranges from an over pricing of about 2½ basis points (at 84 months) to an under pricing of less than 2 basis points (at 24 months). A similar range of mean over- and under-pricing emerges in Hungary, but the largest errors correspond to longer maturities (120 months and 36 months). The range (and size) of pricing errors in Poland, however, is larger with over pricing reaching about 8 basis points (at 120 months) and under pricing close to 6 basis points (at 84 months). In any case, these errors are comparable to those reported by DL for US bonds. Before discussing the stylized facts of the estimated time series of coefficients, it is interesting to note that most of the largest model errors occurred prior to EU accession, that is, before 2004; in Poland the largest residuals were concentrated before 2002 (Figure A1). Regarding volatility of pricing errors, there is scant evidence of a pattern but errors have been more volatile in Poland.

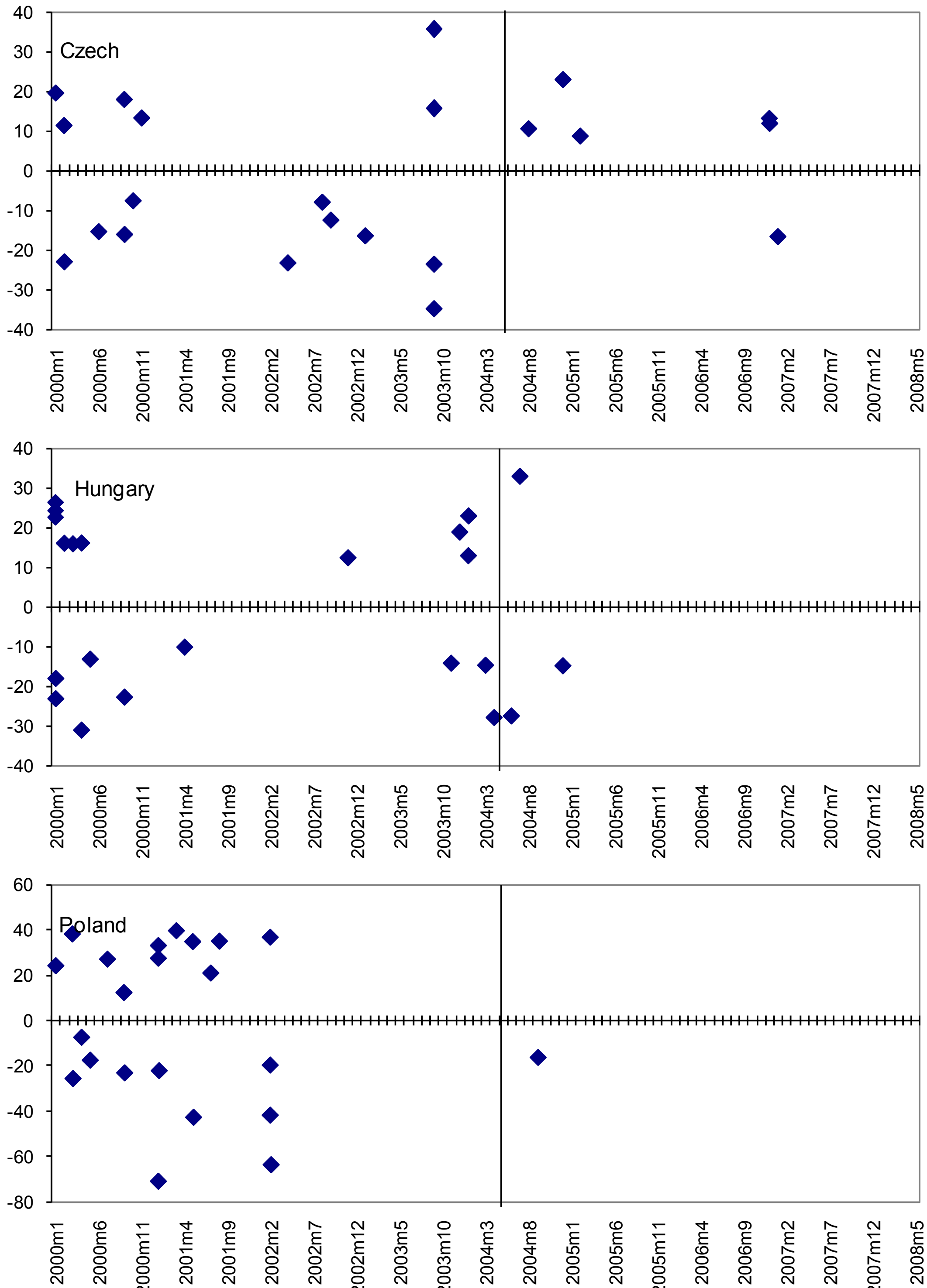
Table A1. Descriptive Statistics of the Yield Curve Residuals

(In basis points, unless otherwise specified)

Maturity (In months)	Mean	Standard deviation
(Czech Republic)		
3	-1.2	6.9
6	0.3	3.5
12	0.4	9.7
24	1.8	7.0
36	1.0	5.4
48	-0.5	6.0
60	-2.2	8.2
84	-2.6	8.5
96	0.9	4.3
108	1.6	4.3
120	0.5	6.0
(Hungary)		
3	1.3	9.0
6	-0.8	7.1
12	-1.5	9.7
24	0.1	5.8
36	2.3	6.1
48	-1.2	7.2
60	-1.4	8.2
84	1.0	6.0
96	1.7	4.3
108	0.2	4.8
120	-1.9	6.0
(Poland)		
3	3.5	11.2
6	-0.2	7.8
12	-3.8	16.2
24	-5.1	11.9
36	0.7	6.5
48	5.4	9.9
60	-0.2	14.9
84	5.9	11.2
96	2.8	6.5
108	-1.1	6.2
120	-7.9	12.8

Source: Authors' calculations using data from 2000:01 through 2008:05.

Figure F1. CEEC's: Maximum and Minimum Yield Curve Residuals, 2000–2008[1/]
(Basis Points)



Source: Authors' calculations using data from 2000m1 through 2008m5.

Figure F2. Euro Area and CECs: Level Factor (β_{1t}) and Long-Term Yields

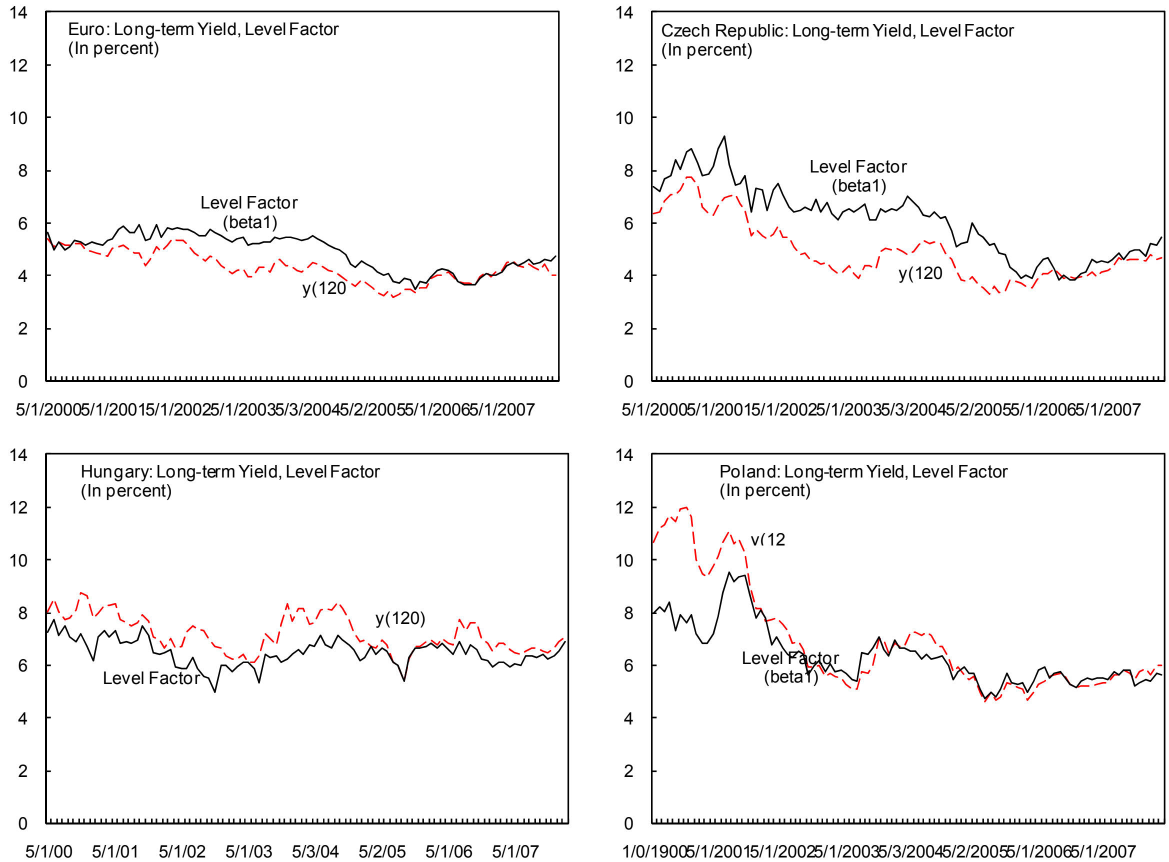


Figure F3. Euro Area and CECs: Slope Factor (β_{2t}) and Slope of the Yield Curve

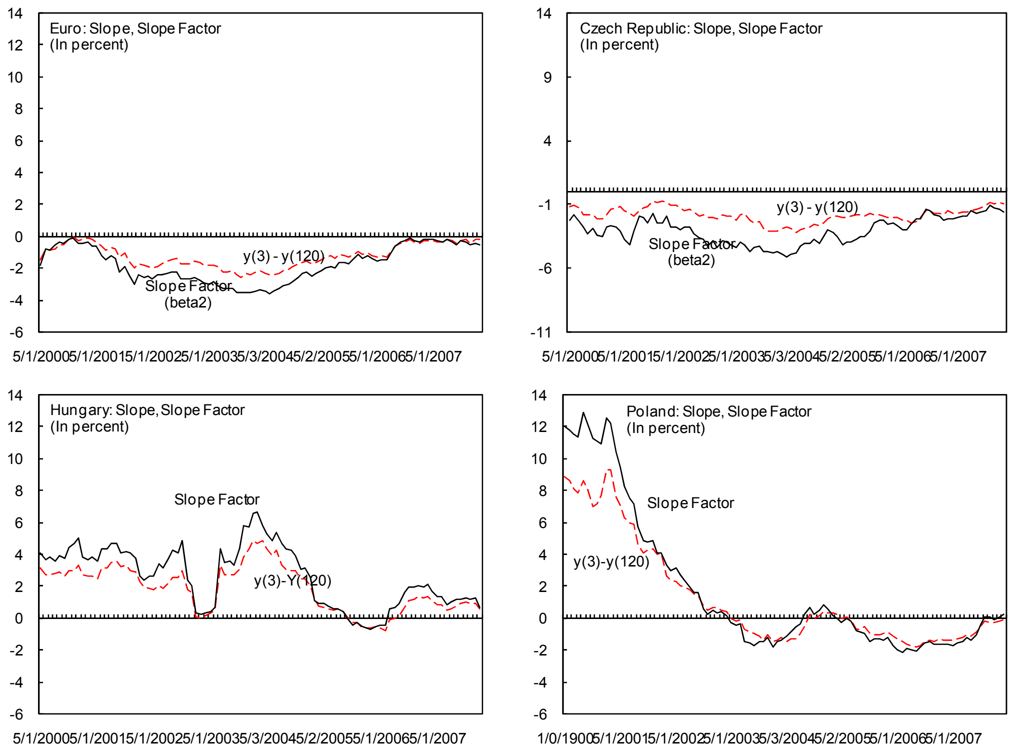
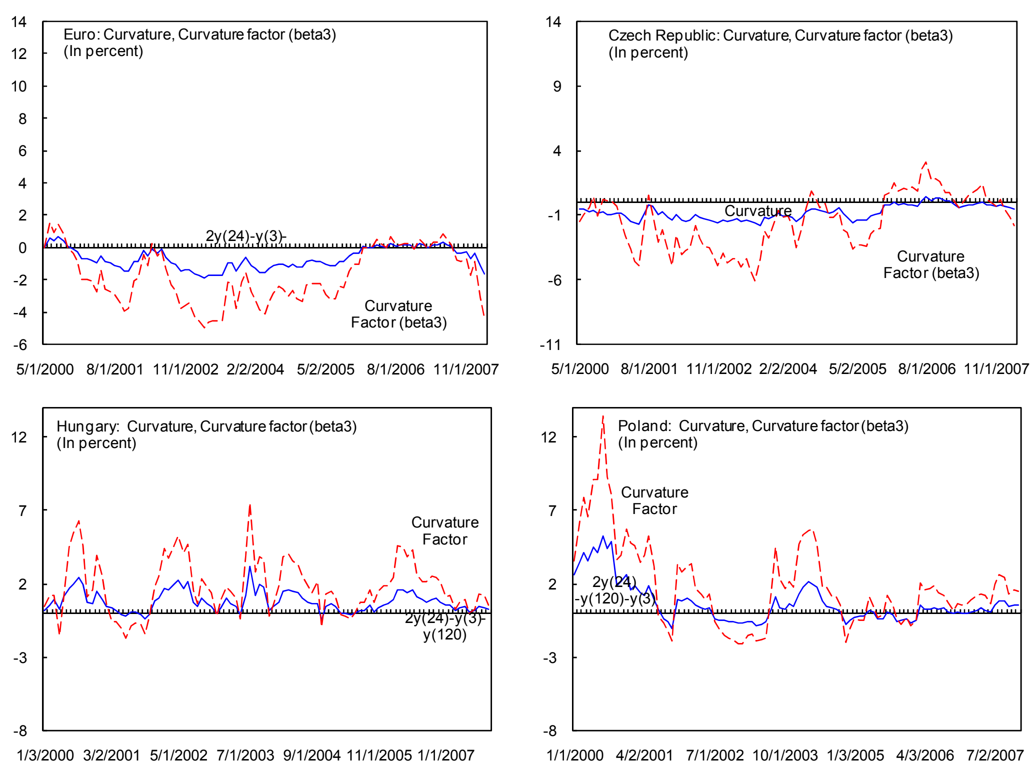


Figure F4. Correlation between Curvature Factor (β_{3t}) and Curvature of Yield Curve



B. Impulse Response Functions

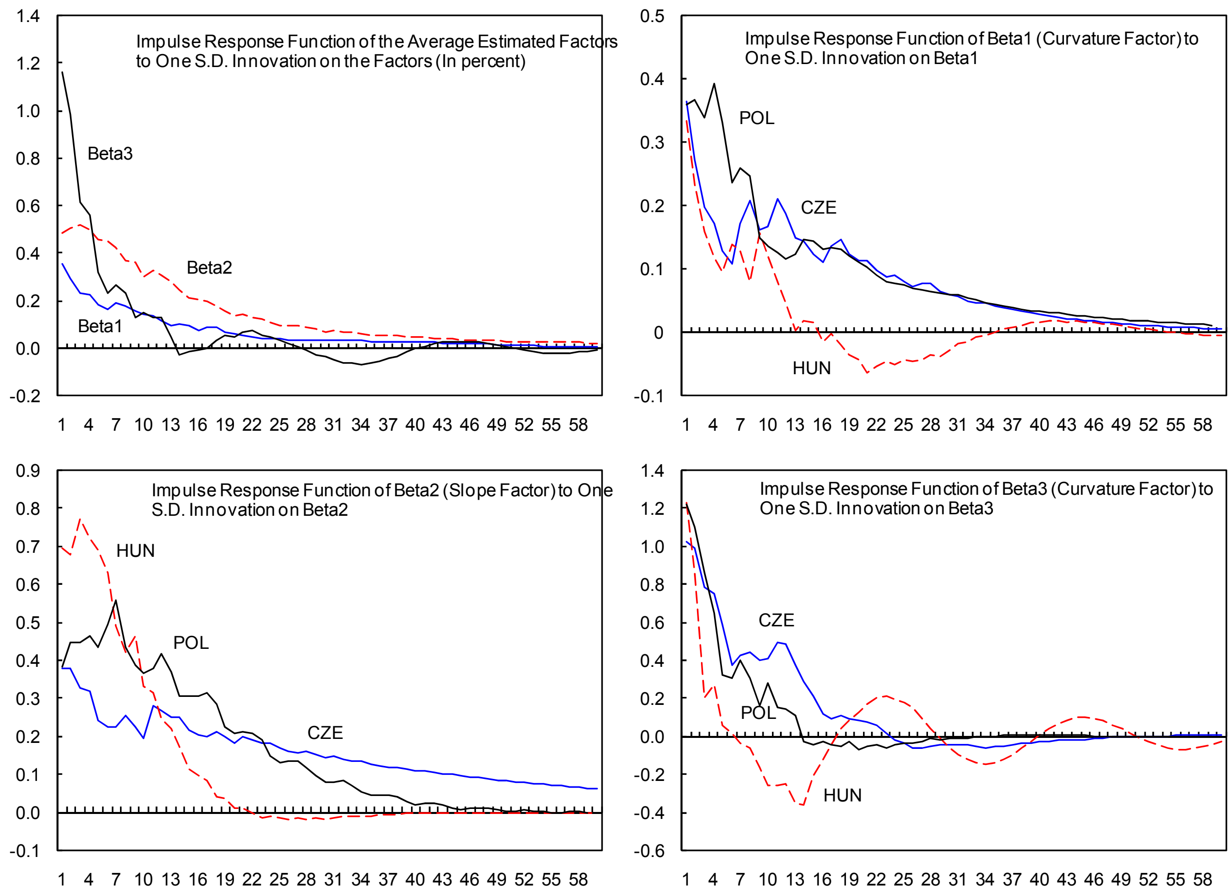
Standard (univariate) time-series analysis have been used to uncover a few characteristics of the estimated parameters of the NS curves. Specifically, autoregressive models (of order 12) were fitted to the estimated $\beta_{i,t}$'s and used to obtain the corresponding impulse response functions (Figure 4). The results suggest that:

- Although shocks to β_{1t} (level) are roughly about 40 basis points in these countries, the persistence of the response of the yield curve lasts longer in Poland where, following the initial jump, the level remains virtually unchanged for about a quarter. In other words, the impact of unanticipated shocks to inflation expectations is more persistent in Poland than in other countries in this study.

- The size and persistence of shocks to β_{2t} (slope) differ. Shocks are larger in Hungary—about 70 basis points compared to 40 basis points in the other two countries—but the effect dies out quicker. As before, following the initial jump in Poland the slope remains virtually unchanged for a number of months (it literally rises further after a few months) before it dissipates. Still, the effect persist the longest in the Czech Republic.

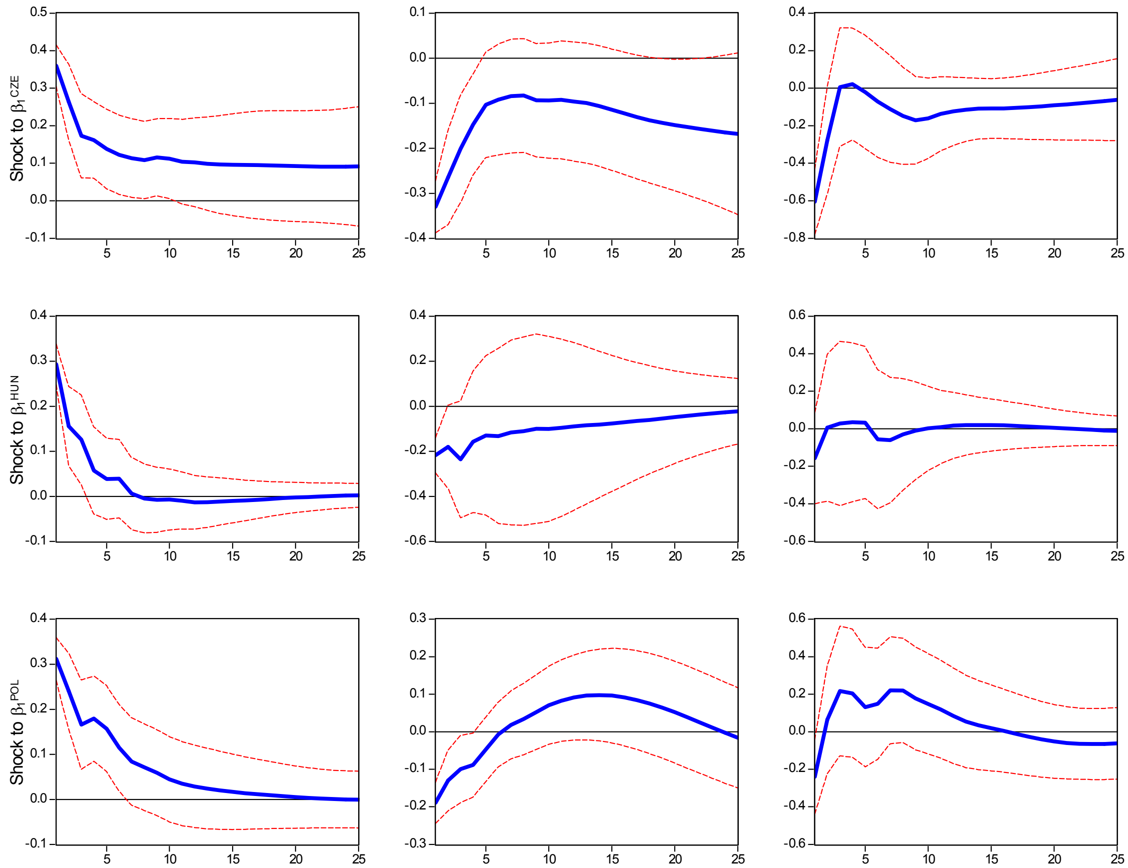
- The size of shocks to β_{3t} (curvature) are larger than those to other $\beta_{i,t}$'s—ranging between 100 and 120 basis points—but their effects dissipate faster.

Figure F5. CEEC3: Impulse Response Functions of the Yield Factors, 2000–08



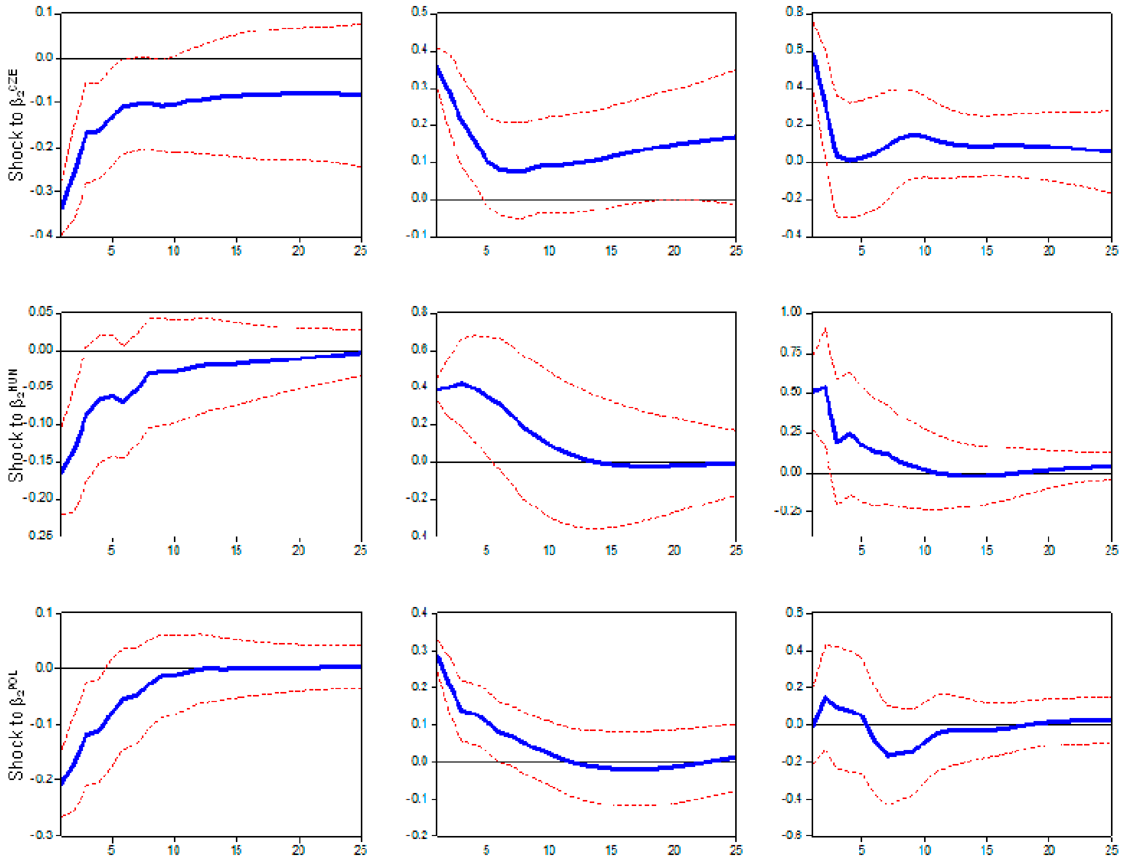
Appendix IV. Impulse Responses for the Nelson Siegel Betas

Figure F6. Response of Nelson Siegel Beta Factors to Shocks to β_{1t} 2000–08
(in percentage points)



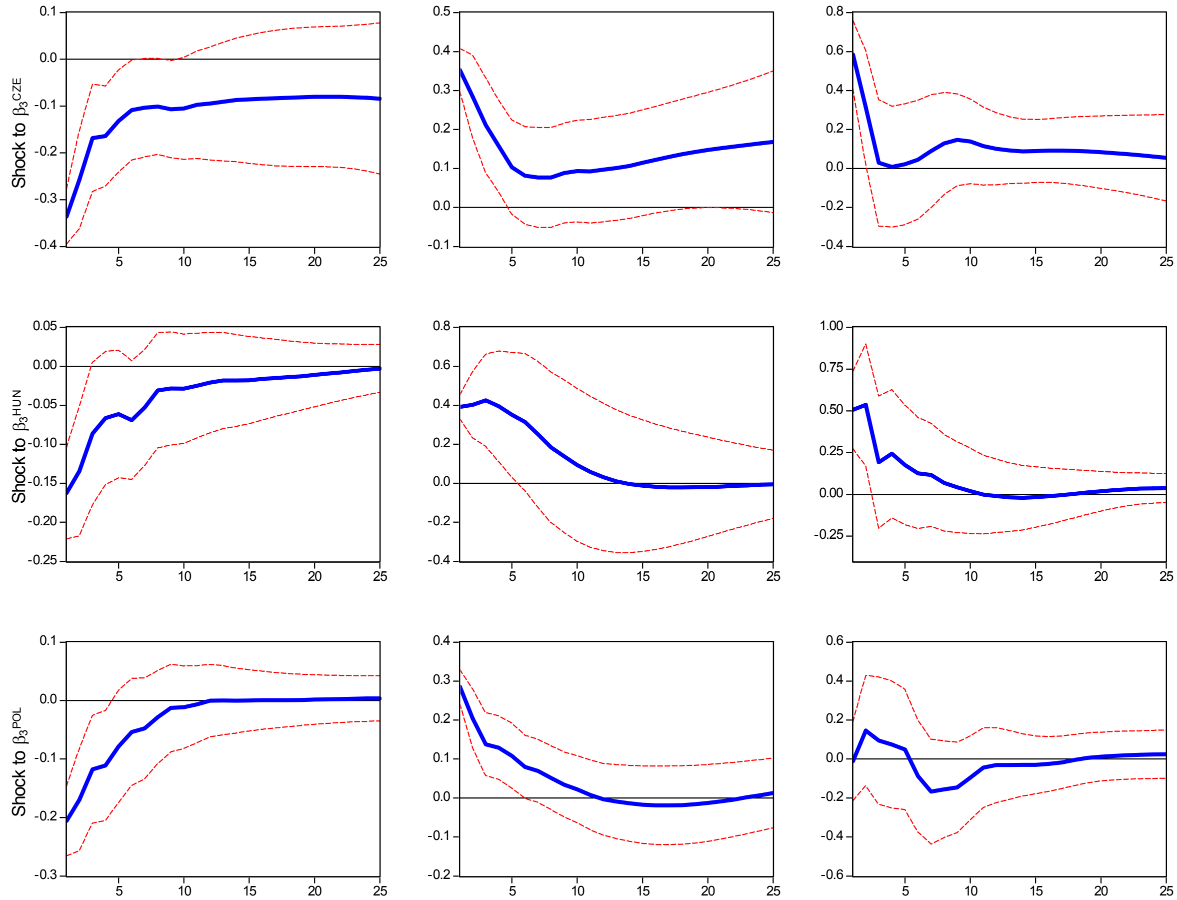
Source: Authors' calculations.
Based on three country VAR models with $x = [i^*, \beta_1, \beta_2, \beta_3, IP, REER, \pi, i^p]$ and L = 3.

Figure F7. Response of Nelson Siegel Beta Factors to Shocks to β_{2t} 2000–08
(in percentage points)



Source: Authors' calculations.
Based on three country VAR models with $x = [r^*, \beta_1, \beta_2, \beta_3, IP, REER, \pi, \pi^*]$ and L = 3.

Figure F8. Response of Nelson Siegel Beta Factors to Shocks to β_{3t} 2000–08
(in percentage points)



Source: Authors' calculations.
Based on three country VAR models with x = [i*, β_1, β_2, β_3, IP, REER, π, i^p] and L = 3.

Appendix V. Estimating Shock for Historical Episodes

Table A2. Shock's Underlying Poland's Euro Adoption Annoucement

Regressors	Coefficient estimates							
	Poland sample				Stacked sample			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	(One-day change in the yield curve, September 10)							
$\mu^{\beta}{}_{1}{}^{POL}$	-0.17			-0.10	-0.19			-0.09
	-1.89			-1.36	-4.70			-1.31
$\mu^{\beta}{}_{2}{}^{POL}$		0.29		0.38		0.28		0.20
		2.57		3.70		7.06		2.47
$\mu^{\beta}{}_{3}{}^{POL}$			-0.51	-1.86			-0.32	-0.47
			-0.76	-5.66			-1.64	-3.62
Standard error of the estimate	0.13	0.12	0.15	0.06	0.09	0.08	0.12	0.06
R squared	--	0.11	--	0.82	--	0.32	--	0.54
R bar	--	0.11	--	0.78	--	0.32	--	0.51
Number of observations	11	11	11	11	33	33	33	33
	(Two-day change in the yield curve, September 11)							
$\mu^{\beta}{}_{1}{}^{POL}$	-0.15			-0.07	-0.24			-0.19
	-1.64			-0.92	-5.89			-2.29
$\mu^{\beta}{}_{2}{}^{POL}$		0.26		0.39		0.32		0.12
		2.40		3.80		6.93		1.25
$\mu^{\beta}{}_{3}{}^{POL}$			-0.57	-1.81			-0.16	-0.43
			-0.89	-5.52			-0.73	-2.72
Standard error of the estimate	0.13	0.11	0.14	0.06	0.09	0.09	0.13	0.08
R squared	--	0.14	--	0.82	--	0.07	--	0.27
R bar	--	0.14	--	0.78	--	0.07	--	0.22
Number of observations	11	11	11	11	33	33	33	33
	(Three-day change in the yield curve, September 12)							
$\mu^{\beta}{}_{1}{}^{POL}$	-0.14			0.03	-0.29			-0.18
	-1.45			0.42	-6.28			-1.71
$\mu^{\beta}{}_{2}{}^{POL}$		0.29		0.50		0.37		0.18
		2.73		4.69		6.78		1.42
$\mu^{\beta}{}_{3}{}^{POL}$			-0.43	-1.59			0.05	-0.22
			-0.66	-4.58			0.20	-1.06
Standard error of the estimate	0.13	0.11	0.14	0.06	0.11	0.10	0.16	0.10
R squared	0.00	0.31	-0.16	0.83	-0.02	0.06	-1.28	0.15
R bar	0.00	0.31	-0.16	0.78	-0.02	0.06	-1.28	0.09
Number of observations	11	11	11	11	33	33	33	33

Source: Authors' calculations based on data from September 10 through September 12, 2009.

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