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Measuring the Financial Markets' Perception of EMU Enlargement:

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Martin Cincibuch and Martina Horníková*

Abstract

Market views on EMU enlargement are measured by a new indicator based on the short-term dynamics of forward spreads. Conceptually, this indicator stems from the notion of uncertainty averse agents and equilibrium indeterminacy. The method was applied on data from central European countries, including the Czech Republic, Hungary, Poland and Slovakia. Comparing our results with financial market opinion surveys, the results of the proposed method seems to be in accordance with market expectations.

Nontechnical Summary

This paper proposes a novel and theoretically motivated method for the measurement of market beliefs (probability) of future euro area enlargement. It is a relevant topic for the central bank, because the reaction of forward-looking market participants to current monetary policy may depend strongly on agents' assessment of future EMU membership. Indeed, there is an alternative way available for determining market expectations, i.e. to rely on survey evidence. However, responses in the surveys do not necessarily reflect respondents' beliefs, and it is important to have a method based on market data which is compatible with market participants' expectations.

This paper differs from other so-called EMU calculators in its attempt to use the dynamics of interest rate spreads, while the typical calculator is based on *the level* of the spreads. Our approach might be particularly useful when interest rate forward spreads are very narrow, which makes level-based methods difficult to apply.

The theoretical background to the paper is the theory of uncertainty aversion[1] (as opposed to risk aversion) originally studied by Bewley and recently developed by Rigotti and Shannon. We use this theory to support the intuitive notion that short-term noisy fluctuations in forward spreads do not necessarily reflect changes in fundamentals and beliefs. In standard models (with rational, forward-looking, risk-averse agents, which maximise their expected utility and in which the financial markets are frictionless), it is hard to achieve such decoupling. In them, any change in the market price reflects information about fundamentals. On the contrary, fundamentals and beliefs in our model only constrain the process of day-to-day changes in forward spreads and may only lead to error correction behaviour. This stems from the fact that the current framework of ambiguity aversion allows more than one market equilibrium supported by one set of fundamentals. In fact, a continuum of equilibrium prices and quantities is possible.

In the applied part we rely on a simple version of our model that takes the form of linear error correction. We investigate Central and Eastern European countries, such as the Czech Republic, Hungary, Poland and Slovakia. Reassuringly, our empirical results correspond well with the findings of the Reuters EMU Poll survey for the Czech Republic, Slovakia, Hungary and Poland. Specifically, as the table shows, the expected entry years indicated by our estimation procedure as of May 2006 correspond well with the latest expected dates from the EMU Poll:

	Estimated date	Latest date from the survey
Czech Republic	2014	2015
Slovakia	2011	2011
Hungary	> 2015	2016
Poland	> 2015	2015

Therefore, using this evidence, one may conclude that the surveys may serve as a reliable proxy for true market expectations.

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[1] Sometimes the term ambiguity aversion is used.

1. Introduction

In this paper, we propose a new approach to the measurement of how markets perceive the prospects of future euro area enlargement. For a monetary authority this is a relevant analysis, because financial market participants are forward looking. Therefore, their reactions to changes in official short-term rates depend on their beliefs about the future time path of interest rates. Future interest rates depend on entry into the monetary union. Thus, to have a good forecast of how the markets will react to current monetary policy, one needs to measure what they think about the EMU prospects of the national economy.

Even before the EMU was launched in 1999 considerable interest was focused on methods of extracting market views of the project, and various methods were designed to infer the probability of a particular country becoming a member of the EMU.

In the EMU calculators, which are typically[2] based on the term structure of interest rates, the EMU entry of a given country is treated as a random event and observed or implied interest rate forwards are used to estimate its probability. In particular, forward spreads are viewed as a weighted average of zero, stemming from the union scenario being realised, and some non-zero value of the non-EMU scenario.

In our approach we attempt to recover additional information by analysing the short-term dynamics of forward spreads. To justify this we appeal to the robust equilibrium indeterminacy arising in the ambiguity aversion model of Rigotti and Shannon (2005). We argue that this approach might be particularly useful when forward spreads are narrow.

In general, there are two ways to assess market perceptions. They can be inferred from prices of market instruments or, alternatively, survey evidence can be relied upon. It is of significance to cross-check these two information sources, because the beliefs expressed in the surveys are not necessarily incentive compatible and may differ from the beliefs or assumptions that investors act on. Therefore, we compare the results of our market-data-based method with the results of the Reuters opinion survey.

The relationship between forwards, the probability of EMU membership at time τ and the conditional expected interest rate differential is usually written as

$$f_{t,\tau,T} - f^*_{t,\tau,T} = \left(1 - \pi^{EMU}_{t,\tau}\right) E_t \left(r_{\tau,T} - r^*_{\tau,T} | nonEMU\right), \tag{1.1}$$

where $r_{\tau,T}$ and $r^*_{\tau,T}$ denote respectively national and foreign (euro area) interest rates[3] as of time τ and with maturity T. Further, $f_{t,\tau,T}$ and $f^*_{t,\tau,T}$ are domestic and foreign interest rate forwards as of time t, with horizon τ and maturity T. Indeed, all the probability distributions involved in relationship (1.1) are risk-neutral ones. The literature mostly ignores this fact, although Bates (1999) argues that this neglect is harmless.

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[2] Bates (1999) surveys the methods and categorises them between those based on currency option contracts (e.g., Butler and Cooper, 1997; Aguilar and Hördahl, 1998) and methods utilising European forward interest rates (e.g., JPMorgan, 1997; Favero, Giavazzi, Iacone and Tabellini, 2000; Angeloni and Violi, 1997; Lund, 1999). A lack of data hinders use of the option based approaches. At the time of the analysis the earliest new EMU entrants could have been expected in several years' time, but the maturity of the interbank currency option contracts for the analysed countries did not extend much over one year. Also, it is too early to apply the time series exchange rate models. Indeed, Aguilar and Hördahl (1998) show that GARCH volatility estimates fell to very low levels only around two years before the EMU was launched.

[3] Before the EMU was launched the asterisk usually denoted Germany, but in the current context it denotes European rates.

There are several complications regarding equation (1.1). First, forward rates for long horizons on the left-hand side might not be traded and then implied forwards need to be estimated. Nevertheless, estimating them from yield or swap curves is quite straightforward and reliable under the no-arbitrage assumption. In the applied part we rely on implied forwards estimated from interest rate swaps via the bootstrapping method.

Second, a much more difficult problem is how to determine the expected future interest rate spread conditional on non-EMU membership at time τ. This is the major aspect in which the term-structure based calculators differ.

Third, as Bates (1999) notes, the EMU calculators are most robust when national and foreign (euro area) interest rates would differ substantially in the case of the country not joining. In other words, formula (1.1) can form a basis for estimating $\pi_{t,\tau}^{EMU}$ only if the expected future spread $E_t\left(r_{\tau,T} - r_{\tau,T}^* | nonEMU\right)$ is large enough in absolute terms. Otherwise, forecast errors and other potential biases would make the EMU and non-EMU cases hard to distinguish and $\pi_{t,\tau}^{EMU}$ would not be identifiable. Therefore, the early research focused mainly on Italy and several other countries with a history of substantial interest rate differentials that could be extrapolated into the future as non-EMU interest rate paths[4]. However, low inflation has prevailed in Central and Eastern European (CEE) countries and many of these potential euro area entrants have independent central banks pursuing inflation targets close to the ECB target. One can assume that the low inflation environment would be sustained and that interest rate spreads would remain low regardless of whether these countries join the euro area. The Czech Republic is a good example of a country for which the EMU and non-EMU scenarios for interest rates could be too close for making reliable assessments about the EMU probabilities using forward rate levels only.

Finally, in our view there is the crucial issue of how to interpret daily fluctuations in forward spreads. Both explanatory factors of the current forward spread on the right-hand side of equation (1.1) characterise medium to long-term expectations, and the fundamental information that could affect them arrives at a relatively low frequency. In other words, it does not seem reasonable that either $\pi_{t,\tau}^{EMU}$ or $E_t\left(r_{\tau,T} - r_{\tau,T}^* | nonEMU\right)$ can vary in such a way as to explain the short-term fluctuations of $f_{t,\tau,T} - f_{t,\tau,T}^*$. However, the short-term dynamics might contain useful information, especially when the conditional non-EMU interest rate spread is low and the probability of entry is high.

These issues are addressed in section 2, where we develop a framework involving Bewley preferences and equilibrium indeterminacy. In section 3 the estimation of a linearised version of this model is estimated for forward spreads of several CEE countries. These empirical results are in section 4 pit against market surveys and discusses official euro strategies, because these are natural reference points for methods aimed at measuring market views. Section 5 concludes. The Appendix discusses the estimation of forward rates.

2. Utilising the Short-term Dynamics of Forward Spreads

Relationship (1.1) is an abstraction which neglects short-term influences that may cause day-to-day fluctuations in the observed forward spreads. We can write a formal decomposition of the

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[4] For instance, at the beginning of 1996, three years before the EMU was launched, the long horizon spreads for Italy were close to four percentage points.

actual market spread (ACT) between the fundamental part (FEQ) and the short-term noise.

$$\left(f_{t,\tau} - f^*_{t,\tau}\right)^{ACT} = \left(f_{t,\tau,T} - f^*_{t,\tau,T}\right)^{FEQ} + \epsilon_{t,\tau}, \tag{2.2}$$

However, we need to clarify which factors might be behind the noise. In fact, it is difficult to explain short-term fluctuations with relatively stable fundamentals (i.e. probability $\pi^{EMU}_{t,\tau}$ and the factors behind the expected future interest rate spread) and to keep the simple paradigm of frictionless markets dominated by rational agents who maximise their expected utility. The modern financial markets always clear and an equilibrium situation prevails. Therefore, one needs a model that allows multiple equilibria supported by one set of fundamentals. One possible option is the general equilibrium model with uncertainty averse agents developed by Rigotti and Shannon (2005). This model features financial markets that are characterised by robust indeterminacy in equilibrium prices and allocations for any specification for initial endowments.

2.1 Equilibrium Indeterminacy

The distinction between risk and uncertainty is a promising way of addressing various financial market puzzles. In the Knightian sense (Knight, 1921), agents face risk when they know the probability distribution of an event. But the event is uncertain when the distribution itself is unknown. Since the expected utility is not readily defined in this framework, there has to be an alternative way of modelling the preferences of agents.

Theories of uncertainty aversion have become increasingly studied recently. Backus, Routledge and Zin (2004) provide an overview of ways of modelling exotic preferences, and a very incomplete list of the more recent research on the subject and its application includes Mukerji and Tallon (2004a,b), Barillas, Hansen and Sargent (2007) and Hansen and Sargent (2007).

In our application, we appeal to Rigotti and Shannon (2005), who, inspired by Bewley (2002), consider a general equilibrium model in which agents' beliefs may be characterised by *multiple priors* and in which an agent prefers one consumption bundle to another one if it has a larger expected utility for *all* priors that the agent considers to be reasonable.

For example, agents may estimate the future interest rate differential using some econometric model, perhaps in a way similar to JPMorgan (1997) or Favero et al. (2000), but they cannot be certain that their model is correct. But even if they believe in the model, they cannot estimate its parameters exactly, because they have only a limited number of observations. They can estimate the mean and variance of the future interest rate spread only with an error. In effect, they obtain not a single distribution of future interest rates, but rather a set of distributions, perhaps characterised by confidence intervals of means and variances. Or viewed from a Bayesian perspective, they arrive at a joint distribution of parameters and variables.

Agents who maximise their expected utility would use the estimated distribution of these parameters and calculate the posterior forecast, but uncertainty averse agents make a different choice. They are indifferent with regard to the outcomes associated with any of the admissible distributions. In other words, an agent who is maximising his expected utility cares even about a marginal difference between the market and his theoretical price, while uncertainty averse agents do not mind unless the difference is significant.

2.2 Ambiguity Aversion in the Context of EMU Calculators

In this section we show that ambiguity aversion may explain the existence of an interval of forward spreads that are all consistent with one set of fundamentals. First, some further notation needs to be introduced. All the contracts we are dealing with are of the same maturity. Therefore, for simplicity, the subscript T indicating their maturity can be dropped. Furthermore, let $x_\tau = r_\tau - r_\tau^*$ denote the future spread as of date τ. And let y_τ be a Bernoulli variable for which $y_\tau = 1$ if the country is an EMU member at time τ and $y_\tau = 0$ if it is not.

Since the money markets merge under EMU membership, rational agents must believe that $x_\tau = 0$ in such case. However in the non-EMU cases, x_τ could be a random variable described at time t by some probability distribution function.

Agents face uncertainty and so their beliefs about x_τ may be characterised by more than one distribution. Denote by $\Gamma_{t,\tau}$ a collection of distributions that agents consider reasonable. Formally, the distributions of the future interest rate spread conditional on EMU membership at time τ as perceived by market participants at time t can be written as

$$
\begin{aligned}
p_t(x_\tau = 0 \mid y_\tau = 1) &= 1 & (2.3) \\
p_t(x_\tau \mid y_\tau = 0) &\in \Gamma_{t,\tau} & (2.4)
\end{aligned}
$$

Agents are also uncertain about the marginal distribution of y_τ. This uncertainty about the Bernoulli distribution is described by some subset $\Pi_{t,\tau}$ of the unit interval so that $p_t(y_\tau = 1) \in \Pi_{t,\tau}$.

The joint distribution of the EMU-membership indicator y_τ and the spread x_τ can be factorised as $p_t(x_\tau, y_\tau) = p_t(x_\tau | y_\tau) \, p_t(y_\tau)$ and therefore the set of reasonable joint distributions $H_{t,\tau}$ is

$$
H_{t,\tau} = \{p_{t,\tau} \; : \; p_{t,\tau} = p_t(x_\tau | y_\tau) \, p_t(y_\tau), \text{ where } p_t(x_\tau | y_\tau) \text{ satisfies (2.3) to (2.4) and } p_t(y_\tau = 1) \in \Pi_{t,\tau}\} \tag{2.5}
$$

If the distribution of future spreads and EMU membership is exogenous for market participants, it is possible to use the indeterminacy result of Rigotti and Shannon (2005). They show that any equilibrium with some set of beliefs is also an equilibrium with a larger set of beliefs. Furthermore, if there is a unique equilibrium only with risk (i.e. with no uncertainty) then all equilibria under uncertainty converge to that risk equilibrium as uncertainty shrinks.

Let us further assume that there exists an equilibrium only with risk for every distribution in $H_{t,\tau}$ and denote by $\widetilde{H}_{t,\tau}$ the set of these risk neutral counterparts. Any $\tilde{p}_{t,\tau} \in \widetilde{H}_{t,\tau}$ can be factorised as

$$
\tilde{p}_{t,\tau} \equiv \tilde{p}_t(x_\tau, y_\tau) = \tilde{p}_t(x_\tau | y_\tau) \tilde{p}_t(y_\tau). \tag{2.6}
$$

The definitions of $\widetilde{\Gamma}_{t,\tau}$ and $\widetilde{\Pi}_{t,\tau}$ follow:

$$
\begin{aligned}
\widetilde{\Gamma}_{t,\tau} &= \{\tilde{g}_{t,\tau} \; : \; \tilde{g}_t(x_\tau) = \tilde{p}_t(x_\tau | y_\tau = 0) \text{ for any } \tilde{p}_{t,\tau} \in \widetilde{H}_{t,\tau}\}, & (2.7) \\
\widetilde{\Pi}_{t,\tau} &= \{\pi_{t,\tau}^{EMU} \; : \; \pi_{t,\tau}^{EMU} = \tilde{p}_t(y_\tau = 1) \text{ for any } \tilde{p}_{t,\tau} \in \widetilde{H}_{t,\tau}\}. & (2.8)
\end{aligned}
$$

Finally, for a particular $\tilde{p}_{t,\tau}$ rewrite relationship (1.1) as

$$f_{t,\tau} - f_{t,\tau}^* = E^{\tilde{p}_{t,\tau}}(x_\tau) = E^{\tilde{p}_{t,\tau}}\left(E^{\tilde{p}_{t,\tau}}(x_\tau | y_\tau)\right) = \tilde{p}_{t,\tau}(y_\tau = 0) E^{\tilde{g}_{t,\tau}}(x_\tau) = \left(1 - \pi_{t,\tau}^{EMU}\right) E^{\tilde{g}_{t,\tau}}(x_\tau) \tag{2.9}$$

We can see that there may exist an interval of forward rate differentials consistent with agents' beliefs. For example, assume that $\widetilde{\Pi}_{t,\tau} = \langle \pi_{t,\tau}^L, \pi_{t,\tau}^H \rangle$ and also that $\gamma_{t,\tau}^L = \inf_{\tilde{g}_{t,\tau} \in \widetilde{\Gamma}_{t,\tau}} E^{\tilde{g}_{t,\tau}}(x_\tau)$ and $\gamma_{t,\tau}^H = \sup_{\tilde{g}_{t,\tau} \in \widetilde{\Gamma}_{t,\tau}} E^{\tilde{g}_{t,\tau}}(x_\tau)$ and that the union of all admissible $E^{\tilde{g}_{t,\tau}}(x_\tau)$ is also an interval. Then it follows from (2.9) that the equilibrium forward rate differential falls in the interval between $B_{t,\tau}^L = (1 - \pi_{t,\tau}^H)\gamma_{t,\tau}^L$ and $B_{t,\tau}^H = (1 - \pi_{t,\tau}^L)\gamma_{t,\tau}^H$.

The size of this band depends on the level of the EMU-membership probability and on the magnitude of the uncertainties involved. For example, assume that the estimated expected future spread is 200 basis points (b.p.) and that the 95% confidence interval of this estimate is 150 b.p. wide. Therefore, we might put $\gamma_{t,\tau}^L = 125\,b.p.$ and $\gamma_{t,\tau}^H = 275\,b.p.$ Unless agents are 100% sure that the country will be a member state at a given date τ, there could be an interval of equilibrium forward spreads. For example, if the uncertainty is $\widetilde{\Pi}_{t,\tau} = \langle 0.9, 1 \rangle$ then any forward spread in $\langle 0, 35 \rangle$ b.p. can represent equilibrium. However, consider the same uncertainty about EMU membership but for a lower probability level; for example let $\widetilde{\Pi}_{t,\tau} = \langle 0.4, 0.5 \rangle$. In this case the band of equilibrium spreads would be much wider, at $\langle 62.5, 165 \rangle$ b.p.

2.3 Band of Inaction with Fully Optimising Speculators

Here we argue that the band of inaction may arise even when the most relevant agents maximise expected utility. Assume that there are two types of market participants, namely hedgers and speculators.

Hedgers enter the market to unload their idiosyncratic interest rate position. These might be, for example, banks with a mismatch between their assets and liabilities or corporations that trade swaps to exploit their comparative advantage in the respective market segments. We assume that they do not have any predictive ability regarding the future interest rate path. In terms of the Rigotti and Shannon (2005) model, one may view them as being uncertainty averse, and considering *any* distribution of future spreads reasonable.

On the other hand, we assume that speculators are expected utility maximisers. At the same time, they gather information and have a good forecasting capability and, therefore, they are those who actually determine the market price and make it informative. We assume that they have no natural position in the future interest rate spread and thus they do not have any hedging needs.

Consider the speculator's problem. If his theoretical future value is higher than the market forward value, then he might consider opening a long forward position. However, he would enter the deal only if its expected return is high enough to compensate for the risk involved. If we abstract from correlations of such investment with returns of other contracts, then its expected return per unit of standard error should be positive. On the other hand, we assume that there is an upper bound for the Sharpe ratio of any available contract. Any deal with a higher Sharpe ratio would be just too good and would be quickly arbitraged away. Let $\lambda_t^{\max}$ denote this upper limit and write this condition as

$$E(return)/\sqrt{var(return)} \le \lambda_t^{\max}. \tag{2.10}$$

In the context of standard asset pricing models $\lambda_t^{\max}$ could be identified with the Sharpe ratio of the market portfolio.

Let $\hat{p}_{t,\tau}$ and $\hat{g}_{t,\tau}$ denote the forecast distributions that characterise the beliefs of the speculator as of time t. Denote also $\hat{\pi}_{t,\tau} = \hat{p}_{t,\tau}(y_\tau = 1)$ and $\Sigma_{t,\tau}^2 = var^{\hat{g}_{t,\tau}}(x_\tau)$. Then his expectations of the future interest rate spread and its variance read:

$$E^{\hat{p}_{t,\tau}}\left(x_\tau\right) = E^{\hat{p}_{t,\tau}}\left(E^{\hat{p}_{t,\tau}}\left(x_\tau|y_\tau\right)\right) = \left(1 - \hat{\pi}_\tau\right) E^{\hat{g}_{t,\tau}}\left(x_\tau\right)$$

$$var^{\hat{p}_{t,\tau}}\left(x_\tau\right) = E^{\hat{p}_{t,\tau}}\left(var^{\hat{p}_{t,\tau}}\left(x_\tau|y_\tau\right)\right) + var^{\hat{p}_{t,\tau}}\left(E^{\hat{p}_{t,\tau}}\left(x_\tau|y_\tau\right)\right)$$

$$= \left(1 - \hat{\pi}_\tau\right)\Sigma_{t,\tau}^2 + \hat{\pi}_{t,\tau}(1 - \hat{\pi}_\tau) E^{\hat{g}_{t,\tau}}\left(x_\tau\right)^2 . \tag{2.11}$$

The expected profit from a long position in the forward spread is therefore

$$E^{\hat{p}_{t,\tau}}\left(x_\tau - (f_{t,\tau} - f_{t,\tau}^*)\right) = \left(1 - \hat{\pi}_{t,\tau}\right) E^{\hat{g}_{t,\tau}}\left(x_\tau\right) - (f_{t,\tau} - f_{t,\tau}^*), \tag{2.12}$$

and the variance of this profit is

$$var^{\hat{p}_{t,\tau}}\left(x_\tau - (f_{t,\tau} - f_{t,\tau}^*)\right) = \left(1 - \hat{\pi}_{t,\tau}\right)\{\Sigma_{t,\tau}^2 + \hat{\pi}_\tau E^{\hat{g}_{t,\tau}}\left(x_\tau\right)^2\}. \tag{2.13}$$

It follows from (2.10) that

$$\left|\left(1 - \hat{\pi}_{t,\tau}\right) E^{\hat{g}_{t,\tau}}\left(x_\tau\right) - (f_{t,\tau} - f_{t,\tau}^*)\right| \leq \lambda_t^{\max}\sqrt{\left(1 - \hat{\pi}_{t,\tau}\right)\{\Sigma_{t,\tau}^2 + \hat{\pi}_{t,\tau} E^{\hat{g}_{t,\tau}}\left(x_\tau\right)^2\}}. \tag{2.14}$$

Inequality (2.14) shows us that there is an interval of admissible forward spreads consistent with a single set of fundamentals reflected in speculators' beliefs $\hat{\pi}_{t,\tau}$, $E^{\hat{g}_{t,\tau}}\left(x_\tau\right)$ and $\Sigma_{t,\tau}$. The right-hand side of inequality (2.14) shows the maximum deviation of the forward spread from the theoretical value which is not arbitraged away. It also defines a band of inaction; any forward spread within the band may represent some market equilibrium. Thus, within the band the forward spread may fluctuate erratically, responding to the immediate supply and demand conditions. But forward-looking agents (i.e. speculators) would prevent the price from moving beyond the band's boundary.

2.4 Model Predictions

Inequality (2.14) implies that when markets are *not* 100% sure that the country will be a member of the monetary union, i.e. when $\hat{\pi}_{t,\tau} < 1$, then the width of this band is positive.

Furthermore, the less precise is the market estimate of the non-EMU conditional spread (i.e. the greater is $\Sigma_{t,\tau}^2$), the wider is the band. Also, the greater is the estimated non-EMU spread $E^{\hat{g}_{t,\tau}}\left(x_\tau\right)$, the wider is the band. For wider bands it could be more difficult to detect any error correction. Therefore, it could be more difficult to find any error correction for countries with a history of volatile and high interest rate spreads. These may be countries that are less integrated into European trade or countries with a history of high inflation.

There is some ambiguity as regards the dependence on the horizon τ. There may be two factors pulling in opposite directions. The percieved probability of entry $\hat{\pi}_{t,\tau}$ is – at least in the baseline scenario – non-decreasing as a function of horizon. This would lead to a narrowing band. On the contrary, the forecasting error $\Sigma_{t,\tau}^2$ increases with the horizon making the inactivity band

wider. The actual balance of the two effects depends on features of individual countries. The first effect could be weak for countries that are likely to join soon because for them $\hat{\pi}_{t,\tau}$ is quite high already for short horizons. The second effect is likely to be stronger for countries that are for some reason more difficult to analyse.

Moreover, for given beliefs regarding the non-EMU spread, the band width also depends on the perceived probability $\hat{\pi}_{t,\tau}$. In particular, for $\Sigma_{t,\tau}^2 > E^{\hat{g}_{t,\tau}}(x_\tau)^2$ the band monotonically widens with declining $\hat{\pi}_{t,\tau}$, otherwise there is a point $\hat{\pi}_{t,\tau}^* = (E^{\hat{g}_{t,\tau}}(x_\tau)^2 - \Sigma_{t,\tau}^2)/2E^{\hat{g}_{t,\tau}}(x_\tau)^2 \in (0,1)$ that maximises the size of the region where prices are not informative.

This dependence is important for the reliability of EMU calculators, which may be relatively high when $\hat{\pi}_{t,\tau} \to 1$ but declines quickly when EMU membership becomes less likely.

2.5 Noise Distribution

The goal of this section is to provide a link between relationship (2.2) and the models of sections 2.1 and 2.3: The reality is more complicated than simple two-period models. Agents are heterogeneous, and there will be trading between times t and τ. Therefore, one can hardly expect any sharp breaks at the edges of the indifference interval. It might be more natural to view them as fuzzy reflecting barriers which push the market back with an intensity negatively dependent on the distance between the market price and the barrier.

The reflecting barriers can be modelled so that the forward spread walks randomly when it appears far enough from any of the barriers, but is pushed strongly back if it approaches a barrier or even moves beyond it. Therefore, when the band is wide enough one might note almost no error correction, but for a narrow band the error correction could be very strong.

The nonlinear error-correction behaviour of $f_{t,\tau}$ and $f_{t,\tau}^*$, which depends on the spread's distance from the barriers $B_{t-1,\tau}^L$ and $B_{t-1,\tau}^H$, may take the form

$$\Delta f_{t,\tau} = -k\left(B_{t-1,\tau}^L - (f_{t-1,\tau} - f_{t-1,\tau}^*)\right) + k\left(f_{t-1,\tau} - f_{t-1,\tau}^* - B_{t-1,\tau}^H\right) + \varepsilon_{t,\tau} \quad (2.15)$$

$$\Delta f_{t,\tau}^* = k^*\left(B_{t-1,\tau}^L - (f_{t-1,\tau,T} - f_{t-1,\tau}^*)\right) - k^*\left(f_{t-1,\tau} - f_{t-1,\tau}^* - B_{t-1,\tau}^H\right) + \varepsilon_{t,\tau}^*, \quad (2.16)$$

where $k(x) = ae^{-\frac{b}{a}x}$ for $x \geq 0$ and $k(x) = -x + a$ for $x < 0$. Here parameter b controls how thick the barrier is, while a determines how strongly it pushes back. Moreover, parameter a must depend on the width of the band. While it needs to accommodate narrow bands, it must also allow for almost independent drift within a wide band. A convenient form might be $a = A \tanh\left(\frac{B_{t-1,\tau}^H - B_{t-1,\tau}^L}{A}\right)$, where parameter $A > 0$ would represent the maximum push for the very wide bands as $\lim \tanh(x) = 1$ for $x \to \infty$.

Finally, function k^* is defined similarly, except that a, b and A are replaced by a^*, b^* and A^*. In this specification, when far enough from any of the barriers the variable can walk almost randomly with a very small drift towards the middle point of the band.

In the empirical part we use a linear approximation of equations (2.15) and (2.16). However, the nonlinear specification is useful for interpretation of the estimated linear error correction coefficients. They represent the average reaction, which is low for bands that are wide relative to the size of shocks and high for narrow bands. As discussed above, the width of the bands is related to the uncertainty and also to the level of perceived EMU probability.

The linear approximation of (2.15) and (2.16) can be written as

$$\Delta f_{t,\tau} = -\alpha_L \left(f_{t-1,\tau} - f^*_{t-1,\tau} - B^L_{t-1,\tau} \right) - \alpha_H \left(f_{t-1,\tau} - f^*_{t-1,\tau} - B^H_{t-1,\tau} \right) + \varepsilon_{t,\tau} \quad (2.17)$$

$$\Delta f^*_{t,\tau} = \alpha^*_L \left(f_{t-1,\tau} - f^*_{t-1,\tau} - B^L_{t-1,\tau} \right) + \alpha^*_H \left(f_{t-1,\tau} - f^*_{t-1,\tau} - B^H_{t-1,\tau} \right) + \varepsilon^*_{t,\tau}. \quad (2.18)$$

Assuming for simplicity that $\alpha_L = \alpha_H = 2\alpha$ and $\alpha^*_L = \alpha^*_H = 2\alpha^*$, we may rewrite the error correction as

$$\Delta f_{t,\tau} = -\alpha \left(f_{t-1,\tau} - f^*_{t-1,\tau} - \beta_{t,\tau} \right) + \varepsilon_{t,\tau} \quad (2.19)$$

$$\Delta f^*_{t,\tau} = \alpha^* \left(f_{t-1,\tau} - f^*_{t-1,\tau} - \beta_{t,\tau} \right) + \varepsilon^*_{t,\tau} \quad (2.20)$$

We may interpret $\beta_{t,\tau}$ as an approximation of $\left(f_{t,\tau} - f^*_{t,\tau} \right)^{FEQ}$. If it moves only slowly with t or if it is constant, then (2.19) and (2.20) can represent an error correction mechanism for the cointegrating relationship (2.2).

3.   Estimation and Empirical Results

If the time series of $f_{t,\tau}$ and $f^*_{t,\tau}$ are not stationary, then (2.2) defines a cointegrating relationship between them. If the non-stationarity of the estimated forward spread can be rejected and the series do not diverge, then we may estimate the error-correction model

$$\Delta f_{t,\tau} = \alpha_1 + \alpha_C \left(f_{t-1,\tau} - f^*_{t-1,\tau} - \beta_C \right) + \Psi(L) \Delta f_{t,\tau} + \Phi(L) \Delta f^*_{t,\tau} + \eta_t \quad (3.21)$$

$$\Delta f^*_{t,\tau} = \alpha^*_1 + \alpha^*_C \left(f_{t-1,\tau} - f^*_{t-1,\tau} - \beta_C \right) + \Psi^*(L) \Delta f_{t,\tau} + \Phi^*(L) \Delta f^*_{t,\tau} + \eta^*_t. \quad (3.22)$$

Using this model we analysed the dynamics of forward differentials against euro rates for several European countries. The countries were selected to have distinct prospects of becoming EMU members and also to have different economic characteristics. In particular, the Czech Republic, Slovakia, Hungary and Poland were analysed. This selection was dictated mainly by data availability. We also considered useful to do the same analysis for some other countries with no or a negligible chance to adopt euro any time soon. Indeed, benchmark results for these countries might help to assess overall usefulness of the method. Therefore, we also analysed Denmark (narrow ERM II band, opt-out clause), Sweden (wide ERM II band, no opt-out clause) and the UK (outside the ERM II opt-out clause).

Because we replace $\beta_{t,\tau}$ in (2.19) and (2.20) by a constant β_C, there is a trade off as regards the sample length. The time series should be as long as possible to capture the dynamic properties, but on the other hand the market assessment of the country's prospects of joining the union may evolve over time, so in this respect a shorter sample would be more desirable. As a compromise we choose to estimate the above specification for the four-month period running up to May 2006, when our sample finishes.

For the sample period we estimated the average error correction coefficients for each country and horizon and tested the residuals for stationarity. To do so we employed the augmented Engle-Granger (AEG) test for residual-based cointegration with a constant and time trend. We used the critical values devised by Davidson and MacKinnon (1993).

Section 2.4 suggests that several regularities should be observed. First, from the discussion of equation (2.14) it follows that the higher is the perceived probability of EMU entry, the tighter should be the no-arbitrage band. For a given disturbance size, a higher perceived probability should lead to higher linear error correction coefficients. Therefore, some prospective EMU entrants might exhibit higher coefficients as the forward horizon increases. This more or less holds for the Czech republic, for which it is unclear when it intends to enter, but which has a stable and transparent monetary policy and therefore the perceived distribution of non-EMU future spreads might not depend strongly on the horizon.

On the other hand, Slovakia has articulated its willingnes to become an EMU member quickly. Thus, the error-correction coefficients do not necessarily need to be increasing function of the horizon, they may even decline for long horizons as a result of increasing $\Sigma_{t,\tau}$.

Conversely, for countries with no EMU prospects one should observe no horizon dependence. Moreover, for countries with independent monetary policy, such as the UK or Switzerland, one might even expect no cointegration. On the other hand, for countries with low π_τ^{EMU} but closely linked to the EMU, for example by a fixed exchange rate regime (e.g. Denmark), cointegration should be detected due to presumably low $\Sigma_{t,\tau}$.

Also, higher uncertainty regarding the conditional non-EMU spread, and also a higher expected value thereof, should lead to a wider band and consequently to lower coefficient estimates. Therefore, countries with a history of relatively high and volatile spreads could have weaker error correction. Thus, we might expect Poland and Hungary to exhibit lower coefficients than, for example, the Czech Republic.

Furthermore, since a small country's rate is more likely to be attracted to the euro rate than vice versa, it seems reasonable to expect that most of the adjustment would happen through changes in $f_{t,\tau}$ rather than $f_{t,\tau}^*$. Therefore, we might expect $\alpha_C \leq 0$ and $\alpha_C^* \approx 0$.

3.1   Data and Estimation of Implied Forwards

Forward contracts are traded for some maturities and horizons, but the implied forwards most often have to be estimated. Estimation is possible using government bond yields or interest rate swap rates. This issue is mostly technical and, compared to the other potential difficulties, is relatively easy to tackle. But it may gain in importance when the absolute difference between forward rates is low relative to the potential errors introduced by the estimation methods.

While Favero et al. (2000) estimate instantaneous forward rates from government bond yields using the specification of Svensson (1994), Lund (1999) derives instantaneous forwards from the zero-coupon curve estimated using the bootstrap method with linear interpolation from interest rate swaps. Others, like JPMorgan (1997) or Angeloni and Violi (1997), directly used forward rates with finite maturity (five- and one-year maturity respectively), also derived from interest rate swap rates. As described in Favero et al. (2000), since the forward rate with horizon τ and maturity T, i.e. $f_{t,\tau}$, is the average of instantaneous forward rates over the period between τ and $\tau + T$, the estimated probability in this case is rather the average of the 'instantaneous' probabilities over the period weighted by the interest rate differentials.

For estimating forward rates, we prefer to use benchmark interest rate swaps rather than government bond yields, because they are standardised and have a favourable structure, which allows for derivation of precise zero coupon curves. This is important because we have to deal with

relatively narrow forward spreads and therefore we tried to avoid any interpolation and ad hoc specification. Therefore, we did not follow Favero et al. (2000) or Lund (1999) in estimating instantaneous forwards from the Nelson-Siegel specification, but rather adapted the approach of Angeloni and Violi (1997).

We estimated one-year forwards directly from benchmark interest rate swaps, quoted in annual maturities. The daily data are available from Bloomberg. First, to extract the term structure of interest rates (the zero coupon curve) we used the bootstrapping procedure (see section 5), which relies only on an assumption of liquid and well arbitraged markets. Then we calculated the implied synthetic one-year forward rates for different horizons.

In general, the data on the benchmark IRS curves are of very good quality, but some large outliers may occur. We checked the data very carefully and cleaned these obvious data errors.

The solid lines on Figure 5.1 graph the IRS yield curves for several countries in comparison with the euro benchmark curve, plotted as a dashed line. Figure 5.2 shows the dynamics of Czech forward rates in relation to euro rates. Similar graphs for other countries are shown in Figures 5.3 to 5.8.

3.2   Czech Republic, Slovakia, Hungary and Poland

The estimates for the *Czech Republic* listed in Table 5.1 reflect the predictions of the model well. Moreover, they are also consistent with surveys of finance professionals in the Reuters poll reported in Table 4.1. This is encouraging, as the method is intended to gauge market views.

The implied forward spreads for the Czech Republic can be considered stationary for all horizons over two years. Moreover, the speeds of adjustment parameters α_{CZK} are all significantly negative and intuitively sized. They increase with the forward horizon and become almost unity for 8-year horizon forwards. This fits well with the Reuters poll. At the time the survey was conducted, all the respondents thought that the Czech Republic would be an EMU member by 2015 at the latest. It is also reassuring that the euro adjustment parameter α_{CZK}^* is insignificant.

The negative value of the equilibrium spread β_{CZK} should not be surprising given the main characteristics of the Czech economy. These include an inflation target of the Czech National Bank that is quite close to that of the euro area, and real appreciation, which apparently stems from the converging economy. These two factors imply trend appreciation of the Czech koruna and, through the interest parity condition, low domestic interest rates.

The results suggest that for short horizons the markets focus on domestic macroeconomic indicators and the inflation forecast of the Czech National Bank, while the long end is likely to be driven by EMU pricing. The current trading practice seems to be in line with such an understanding, as has been confirmed in informal discussions with fixed income dealers.

This interpretation, if true, would have important consequences for the monetary policy of the Czech National Bank. It would mean that even now it can steer only a part of the forward curve, and the closer the unification date gets, the less power its monetary policy will have.

The results for *Slovakia*, reported in Table 5.2, indicate that forward rates can be considered cointegrated with European ones from horizons over five years, which refers to the year 2011.

The estimated coefficients are negative and, from that horizon onwards, also quite high in absolute terms. This corresponds well with the Reuters poll, in which all respondents expected Slovakia to join the EMU in 2011 or before. Also similarly to the Czech case, the euro adjustment parameters are very small and insignificant.

However, the results are not as clear cut as in the Czech case. One might argue that if the interest-rate swap market signals a likely date for EMU enlargement, say τ^*, then after τ^*, α_C should remain as negative as at τ^* and the fundamental spread β_C should remain as close to zero as at τ^*. However, this seems not to be the case with Slovakia.

Therefore, one may be tempted to test, for example, whether the estimated parameter α_{SKK} for the horizon $9 * 10$ is significantly lower in absolute terms than the ones for the shorter horizons. However, one would perhaps be expecting too much from the method. Recall that the linear model (3.21) to (3.22) was inspired by the non-linear error-correction relationships (2.15) and (2.16). Therefore, one should not be surprised by some irregularities. Indeed, the linear method was chosen for simplicity and as a first approximation. Ideally, one should estimate the properly chosen non-linear specification directly.

Cross-country comparisons might be problematic, but weaker error correction could be noticed for Slovakia when compared to the Czech case for horizons beyond the maximum entry (EMU Poll) date. One may speculate that the higher historical level and higher volatility of Slovak interest rates lie behind this result.

The estimation results for *Hungary* in Table 5.3 reveal a much weaker tendency of Hungarian rates to revert towards European rates than in the Czech and Slovak cases. There was an indication of cointegration from the six-year horizon onwards, but the error correction coefficients are low in absolute terms. However, the Reuters poll maximum is 2016, which is even beyond the scope of our empirical analysis.

Polish rates, reported in table 5.4, also exhibit a quite weak tendency to revert to euro rates for longer horizons. The maximum of the EMU poll for Poland is 2015, i.e. on the nine-year forward. Poland's historically high and volatile interest rates are consistent with this result. And note here also that the European parameters α^*_{HF} and α^*_{PLZ} are again virtually zero.

3.3   Other Countries

To obtain a better feel about the new methodology it might be useful to ponder the results for other countries that have distinctly different characteristics, namely Denmark, Sweden and the UK. Denmark participates in the tight ERM II regime, but has negotiated an opt-out clause. Sweden represents an EU country without an opt-out clause but outside of the ERM II, and the UK is both outside the ERM II and opting out.

After a referendum that rejected EMU membership in 2000, *Denmark* participates in the ERM II with a very narrow fluctuation band for its currency. No new referendum is planned and there is no chance of Denmark participating in the EMU in the foreseeable future. The effectively pegged exchange rate with free trade and capital flows means that the Danish central bank has to mimic the monetary policy of the ECB. Under such circumstances one might naturally expect Danish forwards not to diverge far from European ones. And indeed, there is cointegration between these pairs for all horizons, as Table 5.5 shows. However the speed of adjustment α_{DK} does not reveal any dependence on horizon, which is intuitive.

Swedish forward rates are cointegrated with European ones and tend to be affected by the differential over the euro for almost any horizon. On the other hand, the European coefficient α^*_{SK} again came out small and in most cases insignificant.

The United Kingdom's future membership in the euro area remains highly uncertain. The pound floats freely and is a currency of global importance. Under such circumstances, one should expect no easily detectable short-term relationship between the dynamics of British and European forward rates. As demonstrated in Table 5.7, the forward rates tend not to be cointegrated, and for cases where stationarity of the residuals cannot be rejected the error correction coefficients α_{BP} are insignificant.

4. Official Strategies and Survey Interpretations of EMU Enlargement

Since the late 1990s, Reuters newswire services have been conducting market surveys on several different topics concerning EU and EMU enlargement. During the 1990s, such market polls were conducted monthly for the old EU Member States, with the results being released and often compared to several published EMU calculators, such as the JPMorgan EMU calculators (published in the Financial Times). Currently, these polls are conducted biannually on the non-euro area EU Member States that joined the EU in May 2004 (Cyprus, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland and Slovakia) and in January 2007 (Bulgaria and Romania) as well as for several candidate countries (Croatia and Turkey). Reuters surveys around 30 strategists and political analysts across Europe for their views on dates of joining the monetary union and the exchange rate mechanism as one of the preconditions prior to EMU entry. The respondents also provide their expectations about exchange rate parities of national currencies vis-à-vis the euro.

The polls provide a genuinely helpful insight into market perceptions about the timing of euro adoption and about other related issues. However, the results have to be interpreted with some caution, because there are big outliers among them and also because some of the answers are not internally consistent. In particular, some responses to questions about ERM II entry and euro adoption do not reflect the Maastricht requirement that a country should remain in the exchange rate mechanism for at least two years before it is allowed to adopt the euro.

Table 4.1 presents a summary of the latest results on the question of the expected timing of EMU accession for the Czech Republic, Hungary, Poland and Slovakia.

Table 4.1: Expected EMU Entry Dates According Reuters Poll of 38 Professional Respondents in May 2006. (In what year do you expect the following countries to enter the EMU, i.e. formally adopt the euro?)

	Median	Mean	Mode	Latest	Earliest
Czech Rep.	2010	2010	2010	2015	2009
Slovakia	2009	2009	2009	2011	2008
Hungary	2010	2011	2010	2016	2010
Poland	2012	2012	2012	2015	2010

One might expect the Reuters polls to coincide approximately with the official strategies adopted by these countries[5]. In mid-2006, the national euro adoption plans in all the sample countries except for Slovakia were being postponed, mainly due to loose fiscal policy and a worsening fiscal stance. Indeed, the Slovak ministry of finance reported that Slovakia's preparations for entering the Eurozone were proceeding according to the government's plan to adopt the single currency in 2009.

As regards Poland, in May 2006 its former government set 1 January 2012 as the target date for euro introduction. Later on, both the National Bank of Poland and the new government declared an intention to join the euro area as soon as possible, but only after the budget is close to balance. This is expected to delay ERM II entry until 2011 and euro entry until 2013 or 2014. However, opinion polls indicate that most Poles would like the euro to be the Polish currency.

At the time of the survey, Hungary was planning to adopt the euro as its official currency on 1 January 2010, but that date has since been abandoned because of an excessively high budget deficit. Currently, there is no clear target date, but a euro adoption plan is scheduled to be prepared in mid-2008.

Turning to the Czech Republic, its original plan was to enter the ERM II in 2008 or 2009, which was later postponed to 2010. However, the current government has officially dropped any target date, saying the Czech Republic will clearly not meet the economic criteria, mainly due to the large general government deficit. Currently, 2013 is considered the earliest changeover date, although the recently communicated new euro-adoption strategy does not specify any date for euro adoption, preferring to wait for future fiscal developments and the future impact of fiscal reforms.

Table 4.2: Expected Euro Adoption Dates, as of May 2006.

	Estimation	Survey (Latest)	Survey (Mode)	Official strategy
Czech Rep.	2014	2015	2010	*date dropped*
Slovakia	2011	2011	2009	2009
Hungary	> 2015	2016	2010	2010
Poland	> 2015	2015	2012	2012

For a general overview of our model estimates, surveys and official dates see Table 4.2. It shows that the latest date revealed by the poll seems to be the most relevant indicator for comparison with the empirical method. On the other hand, the modes of the survey responses correspond very well with the countries' official strategies.

5.   Conclusions and Suggestions for Future Research

This paper deals with the measurement of market beliefs about the entry of several of the non-euro area EU countries into the euro area. A novel market-based indicator complements the traditional so-called EMU calculators, which use the current level of forward interest rate spreads.

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[5] The national euro-strategies referred to in this section date back to mid-2006, to be in line and comparable with the results of the Reuters polls of the same date as well as with our estimations.

By contrast, the new indicator makes use of their short-term dynamics. Therefore, it is less dependent on the difficult estimation of the hypothetical non-EMU paths for interest rates, which are essential for other EMU calculators. Thus, this method might also be suitable for countries for which the non-EMU scenario is very close to the EMU rates.

The theoretical background to the method is the theory of uncertainty proposed by Bewley (2002) and recently refined by Rigotti and Shannon (2005). Specifically, the most important feature of the theory is that it allows a continuum of equilibrium prices and allocations. This equilibrium indeterminacy justifies modelling short-term noise in the forward spreads and leads to the notion of the constrained inaction band and, in general, to non-linear error correction behaviour.

The empirical investigation that illustrates the theory makes use of a simplified linear version of the error-correction model. It was applied on financial market data in Central European countries, including the Czech Republic, Slovakia, Hungary and Poland. The results were compared with the Reuters EMU Poll survey as well as with the plans presented by national authorities. It turns out that the latest date revealed by the poll seems to be the most relevant indicator for the comparison. It is reassuring that our method and the survey results give a consistent message.

However, the theory of Bewley (2002) and Rigotti and Shannon (2005) is only a static model and therefore the link to the empirical method is rather suggestive. For explicit treatment of the subject, one would need a fully specified dynamic forward-looking model. Also, the empirical model is a linear approximation of the non-linear error correction supported by theoretical considerations, and therefore a change in specification in this direction may lead to further improvement of the results. However, these issues are beyond the scope of the current paper and are left for future research.

Furthermore, the paper presents a partial equilibrium analysis and treats beliefs about future interest rates as exogenous. Ideally, beliefs about equilibrating price variables should be derived from the model parameters via equilibrium analysis. Nevertheless, this also exceeds the narrow subject of EMU calculators and is also left for possible future investigation.

Tables

Table 5.1: Estimated Error Correction Coefficients for the Czech Republic

	Horizon								
	1*2	2*3	3*4	4*5	5*6	6*7	7*8	8*9	9*10
AEG τ-test, (Crit. values for 1%, 5% and 10% are -3.90, -3.34 and -3.04.)	-2.423	-3.499	-5.531	-5.683	-6.534	-6.966	-8.703	-7.506	-8.393
α_{CZK} (p-value)	-0.153 (0.002)	-0.215 (0.000)	-0.304 (0.000)	-0.340 (0.001)	-0.461 (0.000)	-0.587 (0.000)	-0.765 (0.000)	-0.953 (0.000)	-0.919 (0.000)
α^*_{CZK} (p-value)	-0.027 (0.478)	-0.039 (0.396)	0.040 (0.425)	-0.007 (0.899)	-0.014 (0.806)	-0.013 (0.787)	-0.070 (0.198)	0.005 (0.932)	0.070 (0.269)
β_{CZK} s.e.	-0.440 [0.013]	-0.232 [0.011]	-0.141 [0.012]	-0.107 [0.012]	-0.080 [0.010]	-0.083 [0.011]	-0.069 [0.011]	-0.035 [0.011]	-0.008 [0.013]
LM test CZK	0.913	0.304	0.966	0.573	0.058	0.381	0.376	0.791	0.549
LM test EU	0.583	0.881	0.277	0.525	0.795	0.368	0.244	0.196	0.956

Table 5.2: Estimated Error Correction Coefficients for Slovakia

	Horizon								
	1*2	2*3	3*4	4*5	5*6	6*7	7*8	8*9	9*10
AEG τ-test, (Crit. values for 1%, 5% and 10% are -3.90, -3.34 and -3.04.)	-2.104	-2.227	-3.399	-2.588	-7.309	-8.205	-9.192	-8.931	-7.142
α_{SKK} (p-value)	-0.033 (0.277)	-0.080 (0.123)	-0.282 (0.003)	-0.218 (0.005)	-0.720 (0.000)	-0.757 (0.000)	-0.814 (0.000)	-0.764 (0.000)	-0.453 (0.000)
α^*_{SKK} (p-value)	0.011 (0.525)	0.036 (0.248)	0.050 (0.308)	0.010 (0.871)	0.046 (0.317)	-0.026 (0.465)	-0.041 (0.465)	-0.041 (0.481)	0.040 (0.425)
β_{SKK} s.e.	0.454 [0.026]	0.294 [0.015]	0.197 [0.012]	0.116 [0.009]	0.038 [0.012]	0.029 [0.016]	-0.027 [0.011]	-0.078 [0.012]	-0.160 [0.013]
LM test SKK	0.535	0.409	0.265	0.985	0.987	0.196	0.249	0.162	0.052
LM test EU	0.198	0.372	0.827	0.371	0.664	0.624	0.628	0.613	0.561

Table 5.3: Estimated Error Correction Coefficients for Hungary

	Horizon								
	1*2	2*3	3*4	4*5	5*6	6*7	7*8	8*9	9*10
AEG τ-test, (Crit. values for 1%, 5% and 10% are -3.90, -3.34 and -3.04.)	-2.724	-1.581	-1.755	-1.928	-2.387	-5.797	-3.524	-4.762	-3.513
α_{HUF} (p-value)	-0.211 (0.004)	-0.046 (0.390)	-0.027 (0.575)	-0.046 (0.316)	-0.072 (0.260)	-0.122 (0.125)	-0.231 (0.014)	-0.275 (0.005)	-0.267 (0.009)
α^*_{HUF} (p-value)	0.025 (0.341)	-0.007 (0.798)	0.004 (0.890)	-0.004 (0.845)	-0.004 (0.834)	-0.000 (0.989)	-0.017 (0.463)	0.001 (0.971)	0.012 (0.675)
β_{HUF} s.e.	3.259 [0.019]	3.214 [0.020]	3.118 [0.022]	2.975 [0.024]	2.772 [0.024]	2.547 [0.021]	2.246 [0.021]	1.980 [0.021]	1.748 [0.022]
LM test HUF	0.632	0.290	0.609	0.356	0.872	0.388	0.517	0.430	0.821
LM test EU	0.673	0.277	0.717	0.066	0.314	0.364	0.413	0.703	0.227

Table 5.4: Estimated Error Correction Coefficients for Poland

	Horizon								
	1*2	2*3	3*4	4*5	5*6	6*7	7*8	8*9	9*10
AEG τ-test, (Crit. values for 1%, 5% and 10% are -3.90, -3.34 and -3.04.)	-2.397	-1.969	-2.597	-3.555	-4.018	-3.631	-3.236	-2.563	-5.869
α_{PLZ} (p-value)	-0.063 (0.037)	-0.047 (0.087)	-0.057 (0.107)	-0.070 (0.061)	-0.111 (0.058)	-0.106 (0.034)	-0.175 (0.017)	-0.146 (0.020)	-0.169 (0.017)
α^*_{PLZ} (p-value)	-0.010 (0.454)	-0.009 (0.580)	0.005 (0.756)	0.005 (0.765)	0.007 (0.755)	-0.000 (0.992)	-0.011 (0.604)	0.002 (0.932)	0.008 (0.768)
β_{PLZ} s.e.	0.951 [0.034]	1.153 [0.030]	1.245 [0.030]	1.201 [0.029]	1.038 [0.021]	0.932 [0.020]	0.804 [0.021]	0.713 [0.020]	0.622 [0.022]
LM test PLZ	0.080	0.206	0.675	0.749	0.968	0.410	0.775	0.845	0.472
LM test EU	0.107	0.744	0.617	0.106	0.171	0.499	0.397	0.372	0.417

Table 5.5: Estimated Error Correction Coefficients for Denmark

	Horizon								
	1*2	2*3	3*4	4*5	5*6	6*7	7*8	8*9	9*10
AEG τ-test, (Crit. values for 1%, 5% and 10% are -3.90, -3.34 and -3.04.)	-2.520	-4.236	-6.569	-5.657	-5.003	-5.831	-6.313	-6.045	-7.711
α_{DKK} (p-value)	0.037 (0.605)	0.063 (0.535)	0.011 (0.925)	-0.078 (0.571)	-0.012 (0.911)	-0.153 (0.270)	-0.069 (0.652)	-0.144 (0.243)	-0.025 (0.862)
α^*_{DKK} (p-value)	0.077 (0.279)	0.161 (0.099)	0.139 (0.206)	0.113 (0.425)	0.087 (0.402)	0.105 (0.371)	0.159 (0.206)	0.141 (0.265)	0.284 (0.057)
β_{DKK} s.e.	0.117 [0.006]	0.119 [0.005]	0.117 [0.005]	0.118 [0.004]	0.116 [0.004]	0.123 [0.004]	0.120 [0.004]	0.122 [0.005]	0.121 [0.005]
LM test DKK	0.652	0.073	0.194	0.041	0.146	0.525	0.472	0.137	0.207
LM test EU	0.758	0.083	0.183	0.124	0.106	0.457	0.501	0.419	0.514

Table 5.6: Estimated Error Correction Coefficients for Sweden

	Horizon								
	1*2	2*3	3*4	4*5	5*6	6*7	7*8	8*9	9*10
AEG τ-test, (Crit. values for 1%, 5% and 10% are -3.90, -3.34 and -3.04.)	-2.849	-3.925	-4.660	-4.845	-5.637	-6.449	-7.274	-5.583	-6.398
α_{SEK} (p-value)	-0.115 (0.009)	-0.121 (0.011)	-0.183 (0.006)	-0.252 (0.002)	-0.238 (0.027)	-0.581 (0.000)	-0.658 (0.000)	-0.506 (0.000)	-0.500 (0.000)
α^*_{SEK} (p-value)	-0.011 (0.772)	-0.030 (0.498)	-0.010 (0.862)	-0.023 (0.750)	0.136 (0.139)	0.026 (0.758)	0.086 (0.345)	0.078 (0.383)	0.012 (0.886)
β_{SEK} s.e.	-0.115 [0.012]	0.172 [0.012]	0.221 [0.010]	0.194 [0.007]	0.150 [0.006]	0.099 [0.006]	0.015 [0.006]	-0.052 [0.007]	-0.122 [0.008]
LM test SEK	0.440	0.392	0.367	0.543	0.989	0.981	0.493	0.970	0.350
LM test EU	0.914	0.660	0.108	0.157	0.558	0.315	0.695	0.933	0.282

Table 5.7: Estimated Error Correction Coefficients for the United Kingdom

	Horizon								
	1*2	2*3	3*4	4*5	5*6	6*7	7*8	8*9	9*10
AEG τ-test, (Crit. values for 1%, 5% and 10% are -3.90, -3.34 and -3.04.)	-4.597	-4.517	-2.445	-2.116	-2.145	-3.647	-3.451	-3.522	-3.192
α_{GBP} (p-value)	-0.016 (0.629)	-0.046 (0.238)	-0.068 (0.100)	-0.058 (0.087)	-0.101 (0.019)	-0.145 (0.003)	-0.138 (0.003)	-0.186 (0.002)	-0.159 (0.002)
α^*_{GBP} (p-value)	0.022 (0.540)	0.007 (0.836)	0.002 (0.955)	-0.004 (0.915)	-0.019 (0.597)	-0.034 (0.398)	-0.033 (0.395)	-0.032 (0.491)	-0.040 (0.429)
β_{GBP} s.e.	1.203 [0.012]	1.134 [0.013]	0.959 [0.013]	0.810 [0.013]	0.687 [0.012]	0.546 [0.011]	0.382 [0.012]	0.232 [0.012]	0.086 [0.012]
LM test GBP	0.318	0.565	0.298	0.210	0.808	0.948	0.808	0.134	0.141
LM test EU	0.363	0.760	0.712	0.467	0.404	0.348	0.654	0.866	0.578

Figures

Figure 5.1: Swap curves as of 25 Apr 2006. (Solid line - the local currency IRS, dashed line the euro IRS. Maturity in years.)
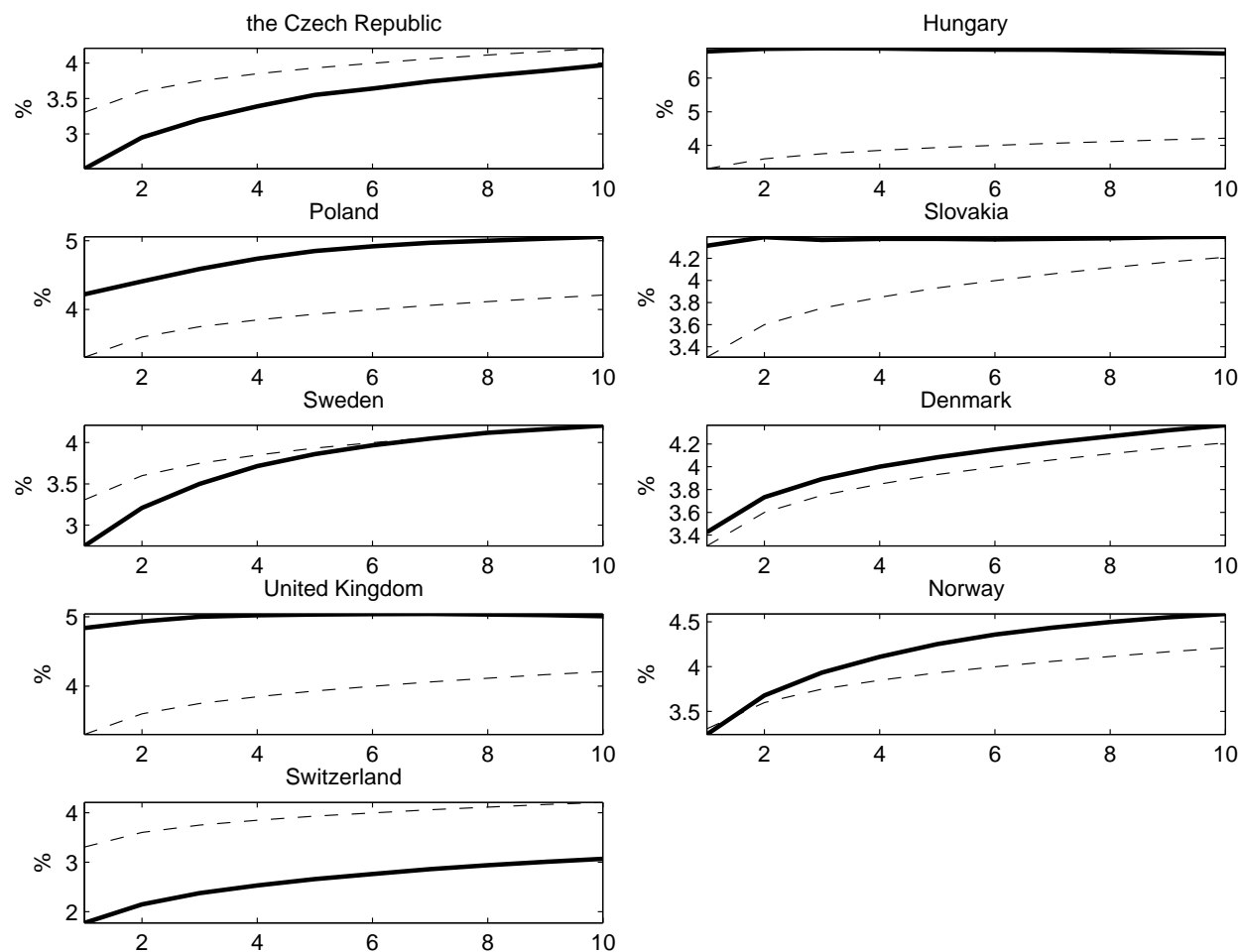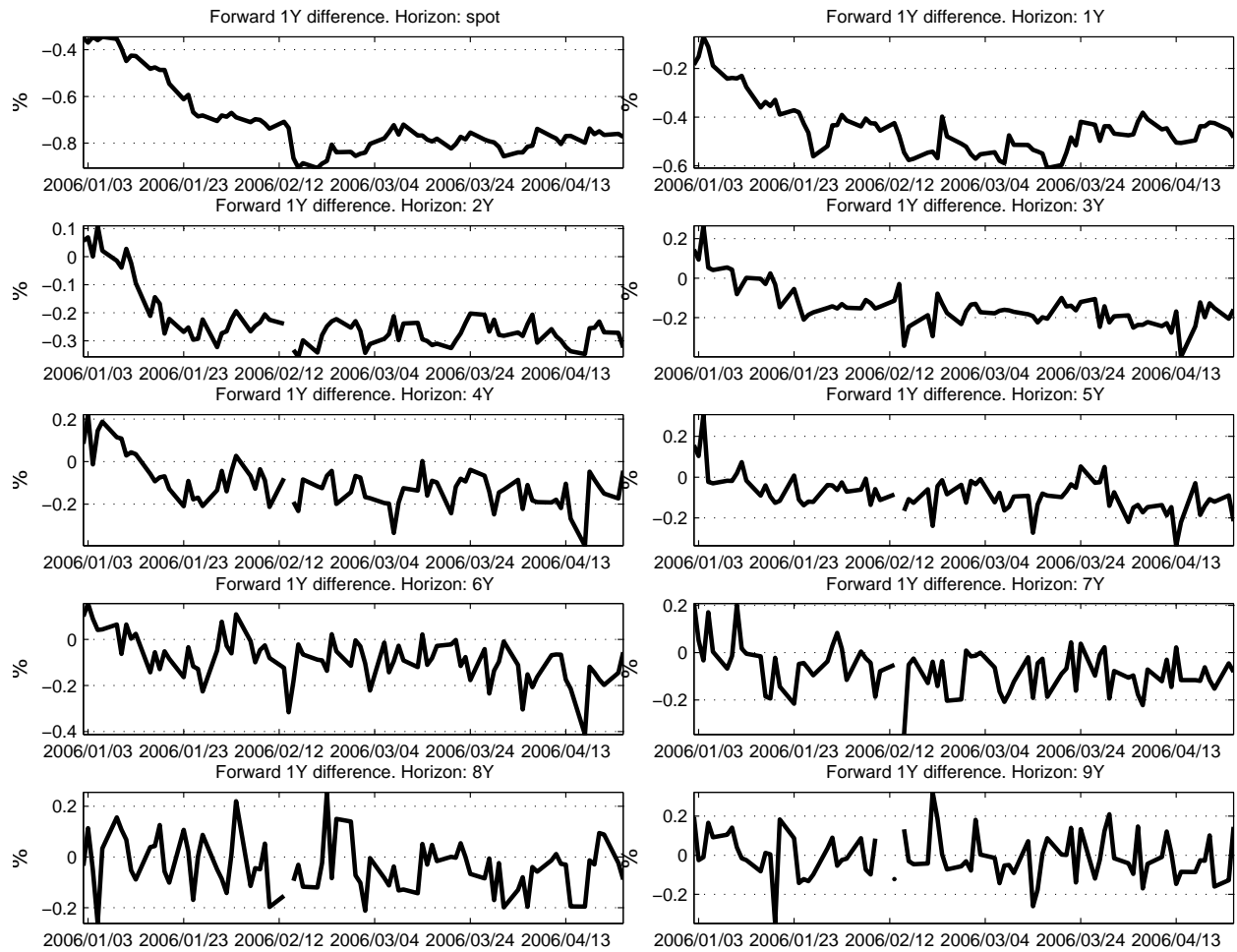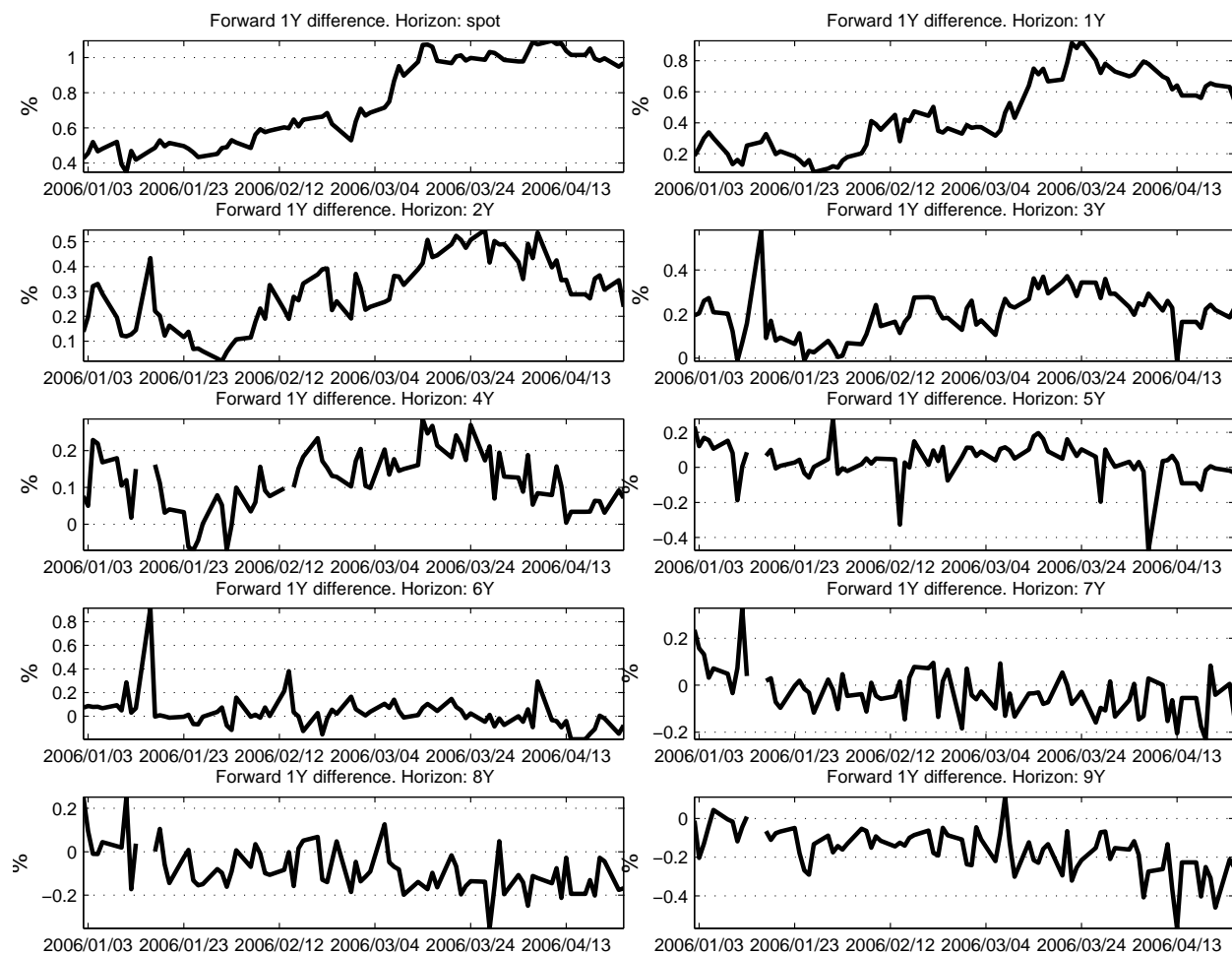
Figure 5.2: EMU and Czech Republic forward spreads

Figure 5.3: EMU and Slovakia forward spreads
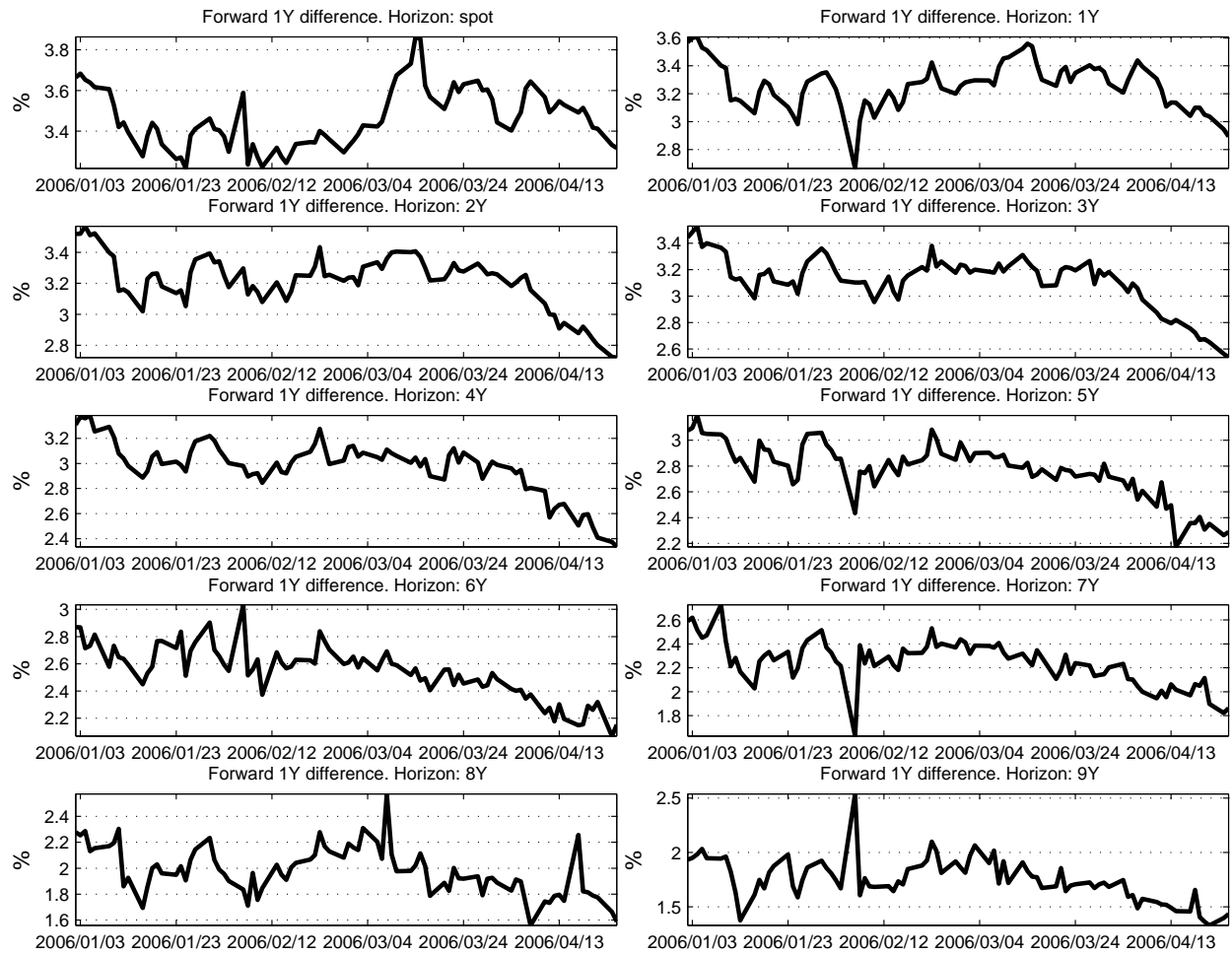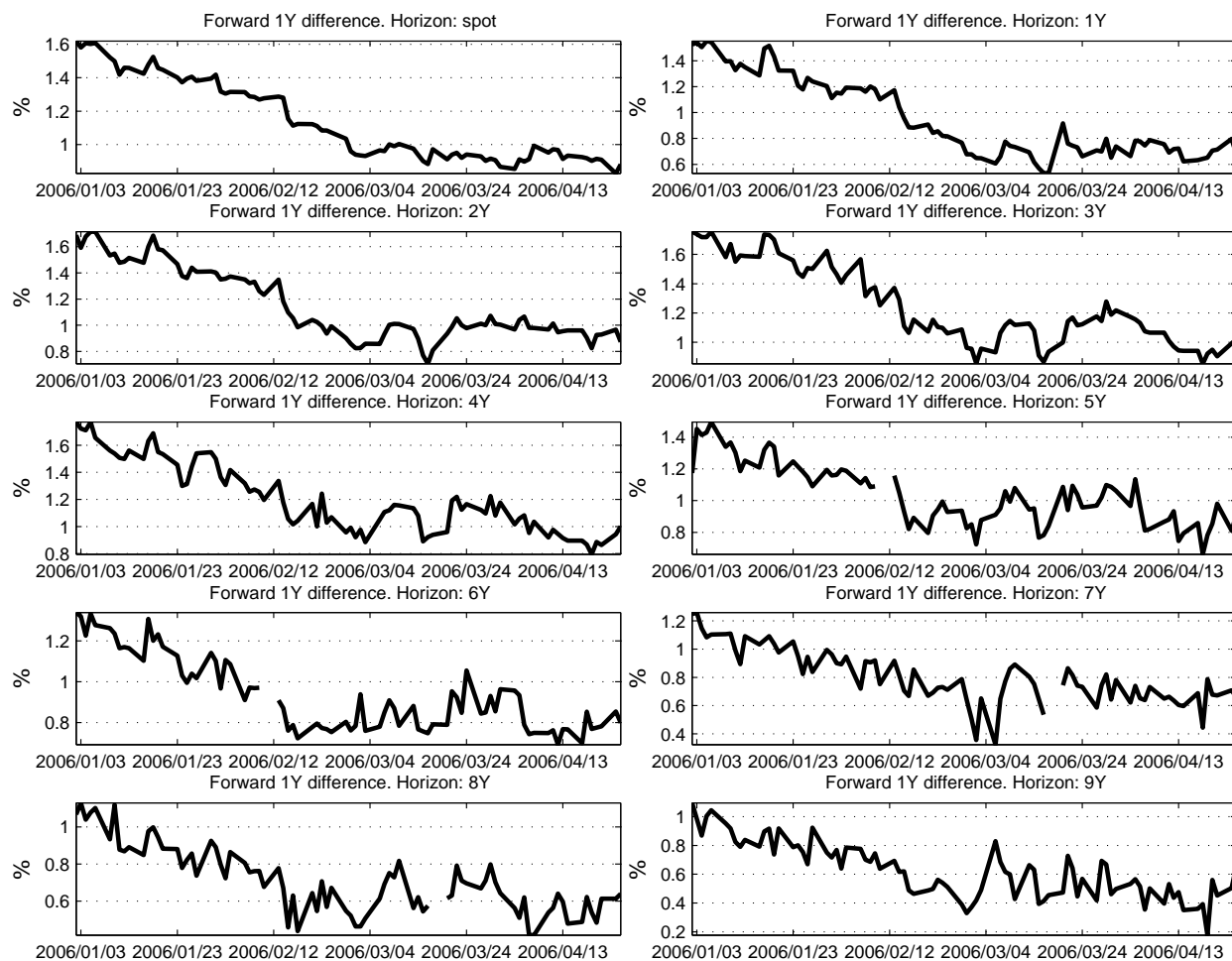
Figure 5.4: EMU and Hungary forward spreads

Figure 5.5: EMU and Poland forward spreads

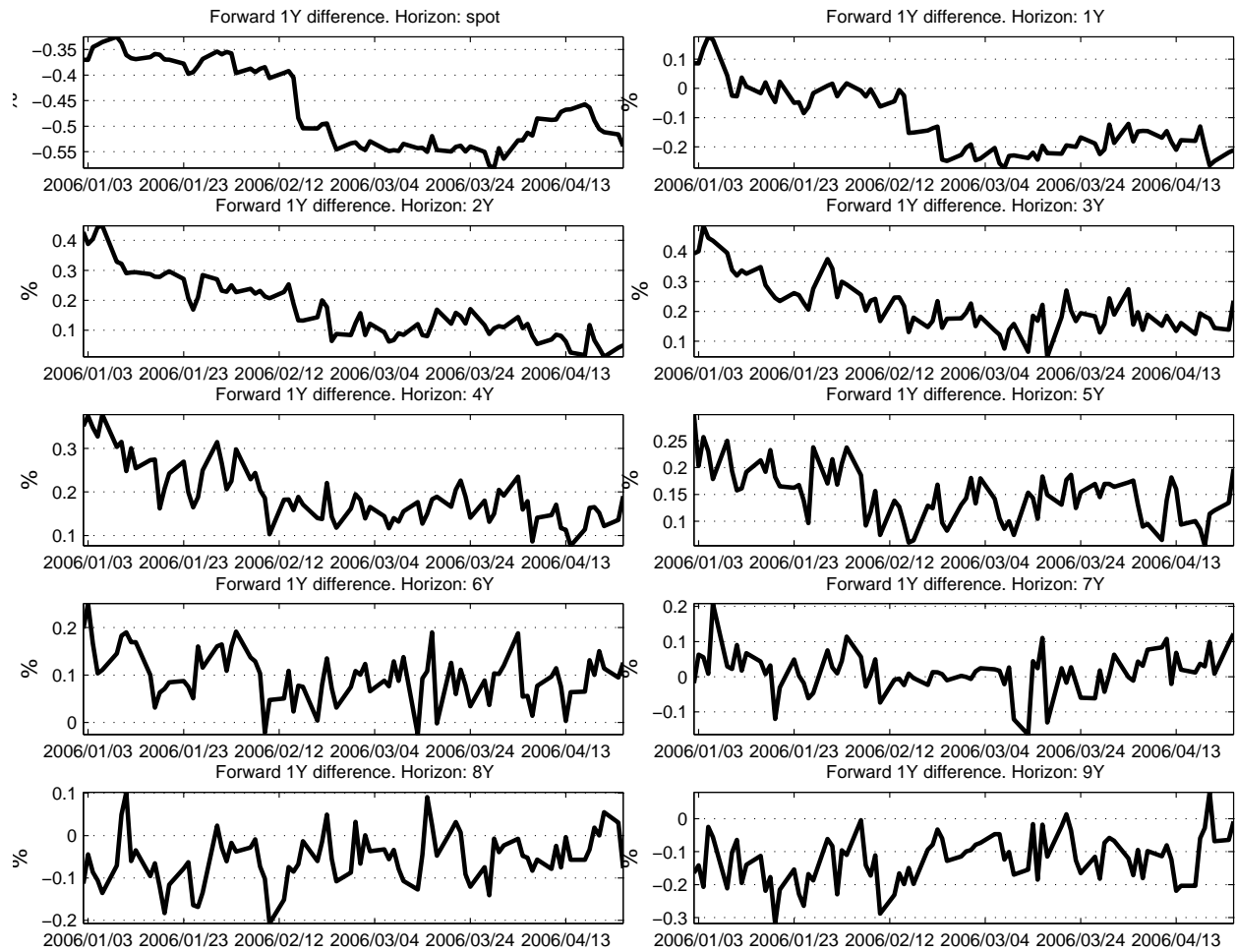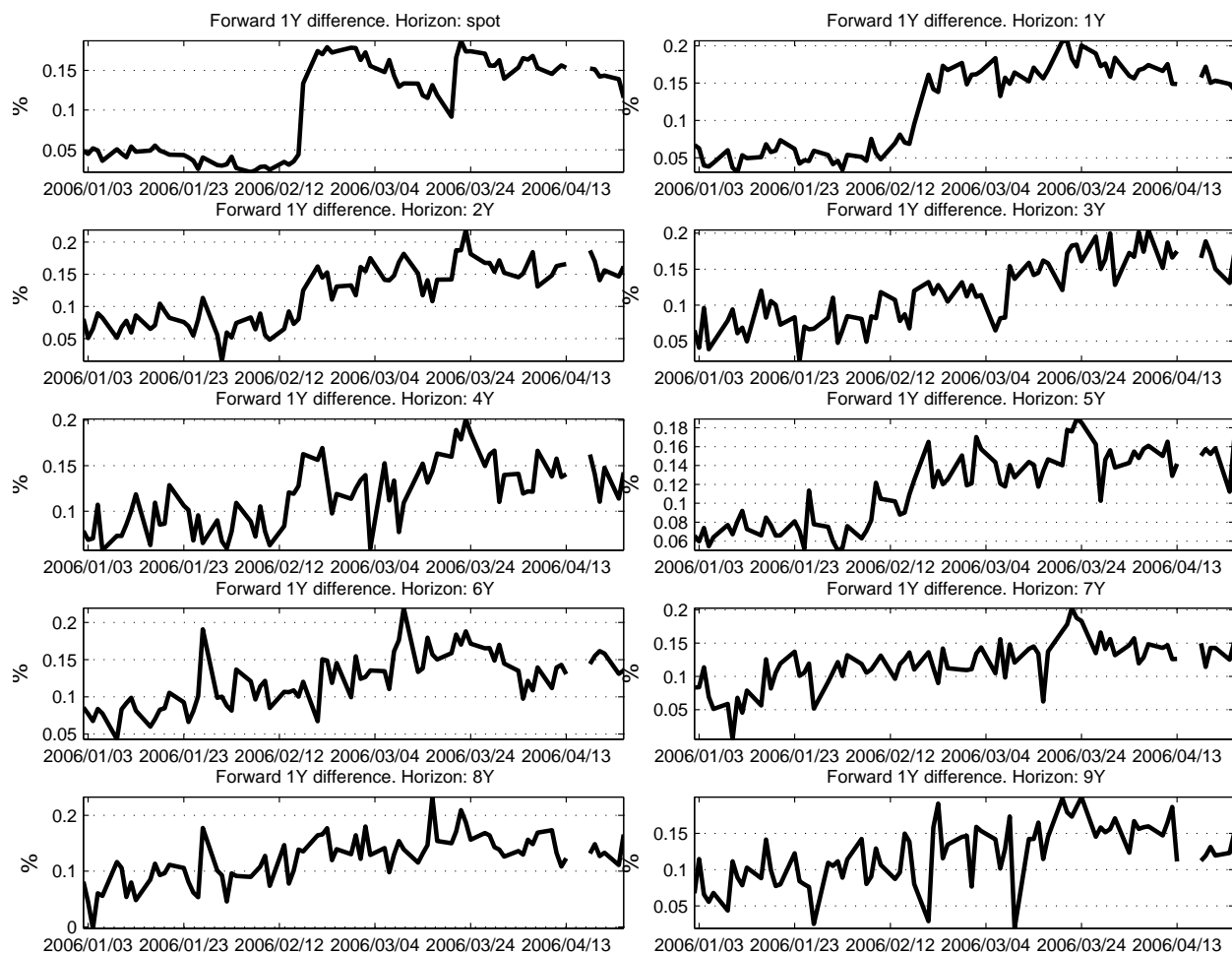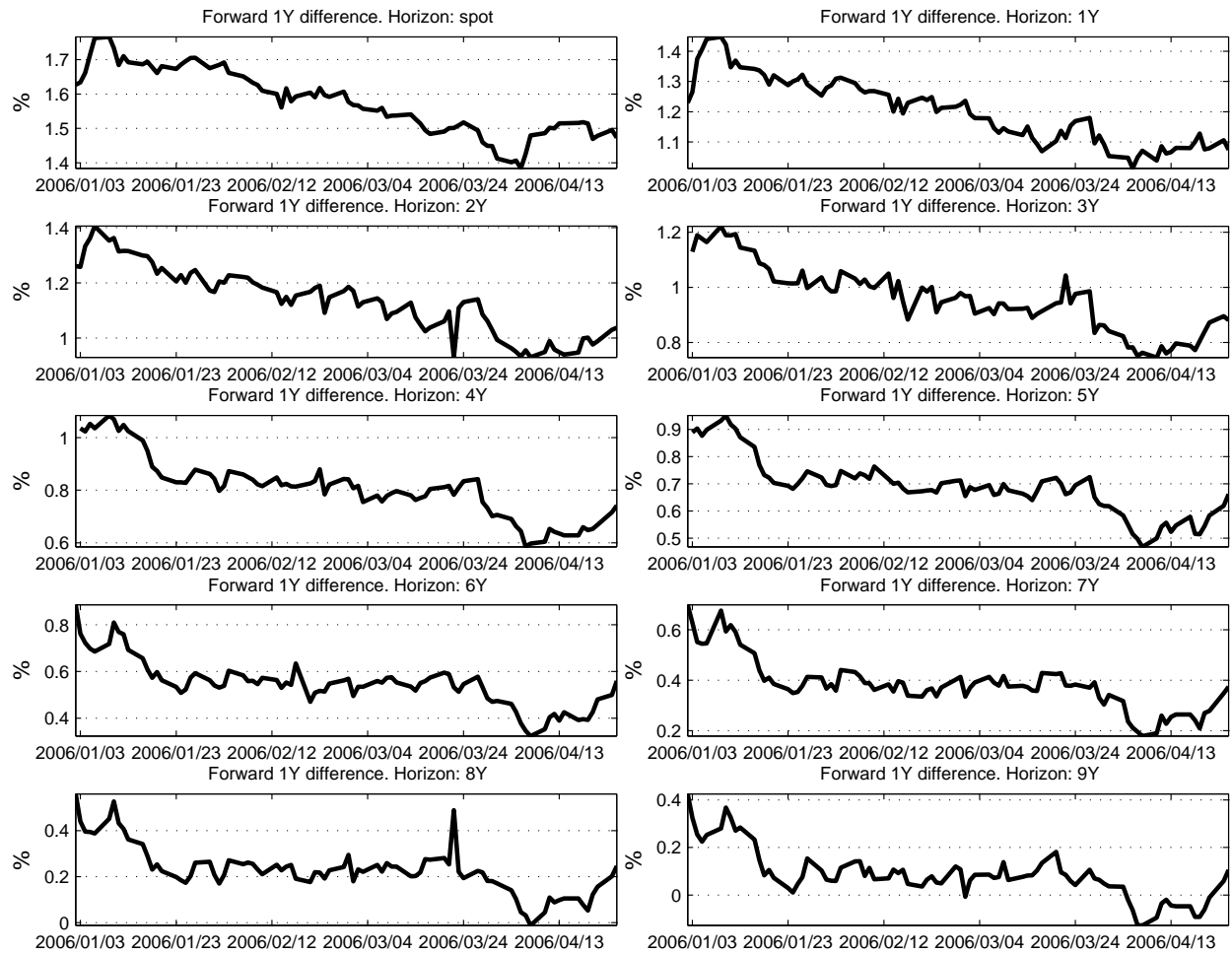## Figure 5.6: EMU and Sweden forward spreads

Figure 5.7: EMU and Denmark forward spreads

Figure 5.8: EMU and United Kingdom forward spreads

Appendix: Bootstrapping procedure

In order to find out the forward interest rate implied by the swap curve it is convenient first to derive the term structure of interest rates (the zero coupon yield curve). As shown below this is possible without any approximation for only some types of swaps, but fortunately, in practice, the suitable swaps are often used. Let $B_{t,M}$ be the price of the discount bond with maturity M applicable at time t. Further, let $I_t^M(m,v)$ denote the interest rate fixed for the floating swap with maturity M as of trade date t, which is based on the floating rate with maturity m. Let the fixed leg of the swap be settled v times a year and let all interest rates be expressed in terms of annual compounding. For simplicity we abstract from any credit risk.

Then the present value of the cash flow for the fixed leg of the swap $I_t^M(m,v)$ on the unit notional amount is given by

$$PV_{fixed} = \left(\left[1 + I_t^M(m,v) \right]^{\frac{1}{v}} - 1 \right) \sum_{k=1}^{vM} B_{t,\frac{k}{v}}. \tag{5.23}$$

Further, let $f_{t,\tau}$ be the forward rate as of the trade date t at horizon τ and maturity T based on the term structure of risk free bonds. Using this term structure of forward rates, swap sellers[6] can hedge their exposure to interest rate risk. The present value of this hedged floating leg's cash flow of the interest rate swap is then

$$PV_{floating} = \sum_{j=0}^{\frac{M}{m}} \left[(1 + f_{t,jm,m})^m - 1 \right] B_{t,(j+1)m} \tag{5.24}$$

The non-existence of arbitrage opportunities further dictates the relationship between discount rates and forward rates obviously expressed as

$$\frac{B_{t,jm}}{B_{t,(j+1)m}} = (1 + f_{t,jm,m})^m, \tag{5.25}$$

so when it is substituted in (5.24) for $(1 + f_{t,jm,m})^m$ one may write

$$PV_{floating} = \sum_{j=0}^{\frac{M}{m}} B_{t,jm} - B_{t,(j+1)m}$$

and since most of the terms in this series cancel it is possible to conclude that

$$PV_{floating} = 1 - B_{t,M} \tag{5.26}$$

Since both the fixed leg and the hedged floating leg represent streams of certain payments, the no-arbitrage condition on the swap rate is that the present values of both payment streams are equal, therefore

$$\left(\left[1 + I_t^M(m,v) \right]^{\frac{1}{v}} - 1 \right) \sum_{k=1}^{vM} B_{t,\frac{k}{v}} = 1 - B_{t,M}. \tag{5.27}$$

---

[6] Swap sellers receive fixed rate and pay floating rate payments.

This formula relates prices of discount bonds and interest rate swap rates. Further, if a sufficient number of interest rate swaps is traded, then it is possible to use this formula to calculate prices of discount bonds recursively. After some straightforward algebraic manipulations it follows from (5.27) that

$$B_{t,\frac{1}{v}} = \frac{1}{\left[1 + I_t^{\frac{1}{v}}(m,v)\right]^{\frac{1}{v}}} \tag{5.28a}$$

$$B_{t,M} = \frac{1 + \sum_{k=1}^{vM-1} B_{t,\frac{k}{v}}}{\left[1 + I_t^M(m,v)\right]^{\frac{1}{v}}} - \sum_{k=1}^{vM-1} B_{t,\frac{k}{v}}, \text{ for } M > \frac{1}{v} \tag{5.28b}$$

(5.28) shows that if prices of discount bonds with maturities $\frac{1}{v}, \frac{2}{v}, \frac{3}{v}, ..., \frac{vM-1}{v}$ are known, then knowledge of the swap rate $I_t^M(m,v)$ enables determination of the discount factor $B_{t,M}$.

If these discount bond prices are known, then, similarly to (5.25), one may obtain the implied forward rates of maturity $\frac{1}{v}$ as

$$f_{t,\frac{k-1}{v},\frac{1}{v}} = \left(\frac{B_{t,\frac{k-1}{v}}}{B_{t,\frac{k}{v}}}\right)^{\frac{1}{v}} - 1. \tag{5.29}$$

The possibility of performing this procedure hinges on the condition that there are enough points on the swap curve in relation to the settlement frequency of the fixed part of the swap contracts. In particular, there must be vT equally spaced swap rates to allow determination of the vT discount factors. On the contrary, there is no such condition on the maturity of the underlying floating rate. Fortunately, and perhaps not surprisingly, swap rates are often quoted for maturities in whole years and with annual settlement, i.e. $v = 1$, which facilitates the empirical analysis.

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