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DISSERTATION

Titel der Dissertation

„Essays on Exchange Rate Regime-Related Risks in CEE Countries"

Verfasser

Diplom-Volkswirt Tomáš Slačík

angestrebter akademischer Grad

Doctor of Philosophy (PhD)

Wien, im März 2008

Does the choice of exchange rate regime matter? Few questions in international economics have sparked as much debate and yielded as little consensus.

-A. R. Ghosh, A. M. Gulde and H. C. Wolf in Exchange Rate Regimes: Choices and Consequences. The MIT Press. Cambridge 2002

Essays on Exchange Rate Regime-Related Risks in CEE Countries

Tomáš Slačík

To my beloved grandma, wherever she is...

Abstract

The exchange rate is one of the key prices in an economy. By directly affecting the behaviour and the decision making of economic agents such as households, entrepreneurs, governments, banks and other financial institutions it has a crucial impact on many important macroeconomic variables. The exchange rate can function as a vital anchor and a stabilising force. However, particularly against the backdrop of increasingly globalised capital flows, each exchange rate regime holds its specific immanent risks. The present dissertation delves in three chapters into two particular exchange rate regime related risk aspects - the peril of a currency crisis and the possibility of an inflation acceleration in the wake of the euro adoption hotly debated in most of the euro candidate countries.

In the first chapter we thus propose exploiting the term structure of relative interest rates to obtain estimates of changes in the timing of a currency crisis as perceived by market participants. Our early warning indicator can be used to evaluate the relative probability of a crisis occurring in one week as compared to a crisis happening after one week but in less than a month. We give empirical evidence that the indicator performs well for two important currency crises in Eastern Europe: the crisis in the Czech Republic in 1997 and the Russian crisis in 1998.

Also the second chapter deals with early warning mechanisms for currency crises. We tackle explicitly the issue of model uncertainty in the framework of binary variable models of currency crises. Using Bayesian model averaging techniques, we assess the robustness of the explanatory variables proposed in the recent literature for both static and dynamic models. Our results indicate that the variables belonging to the set of macroeconomic fundamentals proposed by the literature are very fragile determinants of the occurrence of currency crises. The results improve if the crisis index identifies a crisis period instead of a crisis occurrence. In this setting, the extent of real exchange rate misalignment and financial market indicators appear as robust determinants of crisis periods.

In contrast, using the case study of the Czech Republic in the last chapter we analyse relevant macro and microeconomic forces driving inflation with a particular focus on how these inflation channels are likely to change in the wake of the euro adoption. We employ an autoregressive distributed lag (ARDL) model combined with the Bayesian model averaging and a Bayesian model selection technique based on posterior model inclusion probabilities. As a side-product we estimate the time-

varying natural rate of interest purged from the risk premia which is, to our best knowledge, the first attempt to do so for the Czech Republic. Our results suggest that the costs attributable to the lack of koruna appreciation after euro adoption are likely to be rather low. In contrast, a low inflation environment and a harmonization of the business cycles between the Czech Republic and the euro area are essential for a smooth inflation development after the euro adoption. The fulfilment of the Maastricht inflation criterion should not be enforced, however, by non-standard policy measures. The potential inflationary effect of the changeover cannot be eliminated altogether but it may well be substantially reduced.

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Chapter 1

An "Almost-Too-Late" Warning Mechanism For Currency Crises[1]

1.1   Introduction

The abundance and severity of currency crises and speculative attacks over the past 15 years has spawned a rebirth of interest among researchers, politicians and central bankers in these events. Currency crises tend to be painful and costly for the affected economy and usually afflict the population immediately. Two recent cases illustrate this. The Argentine financial turmoil of 2002, one of the most violent currency crises on record, induced an increase in the poverty rate of more than 50% and drove nearly two out of three Argentines below the poverty line. The Indonesian crisis in 1998 caused a drop in GDP of more than 13% within a year.

Considering the spectrum of distinct exchange rate arrangements, ranging from totally flexible exchange rates to monetary unions and currency boards, there seems to be ample evidence that intermediate regimes such as fixed and crawling pegs or fluctuation bands are most prone to speculative attacks. For this reason, countries have increasingly abandoned intermediate regimes for arrangements at the extremes (see e.g. Fischer, 2001).[2] Yet today an unravelling of current global imbalances threatens even a large economy with a flexible exchange rate regime like the United States with abrupt and significant currency depreciation (see Roubini and Setser, 2005). Admittedly, experts who consider this possibility as remote likely outnumber proponents

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[1]This chapter is based on a joint work with my adviser, Prof. Dr. Jesús Crespo-Cuaresma.

[2]Notwithstanding the fact that *de jure* exchange rate regimes need not coincide with *de facto* ones (see Calvo and Reinhart, 2002).

of the view that the Bretton Woods 2 system of exchange rates is unsustainable (see e.g. Dooley et al, 2004; or Hausmann and Sturzenegger, 2005). However, there are undoubtedly countries whose fundamental macroeconomic and/or political conditions make them vulnerable to a speculative attack. Hungary, for example, which still maintains an intermediate exchange rate arrangement, was running a current account and fiscal deficit in 2005 roughly 2.5 times greater than Argentina before its devastating crisis hit.

Motivated by the damaging potential effects of currency crises, both the theoretical and empirical literature has received renewed attention from researchers. Not only have new generations of currency crisis models been developed in response to unsatisfactory theoretical instruments to describe and explain causes and frequency of crises in the 1990s, economists have also started pondering ways to predict the timing of such events (e.g. both the Mexican crisis in 1994 and the Asian crisis in 1997 essentially caught the international community flat-footed). Unfortunately, the forecasting models proposed over the past few years have generally demonstrated only fair-to-middling predictive accuracy.

Kaminsky, Lizondo and Reinhart (KLR) pioneered quantitative currency crisis early-warning systems (EWS) with their "indicator" model developed in a series of papers (see Kaminsky, Lizondo and Reinhart, 1998). The KLR approach monitors the evolution of several economic variables (indicators) such that when a variable deviates from its "normal" level beyond a threshold value, a signal is said to be issued. The threshold value is chosen as follows. Let S represent the number of periods in which the indicator issued a good signal (a crisis signal which was followed by a crisis, or a non-crisis signal which was not followed by a crisis) and let N denote the number of periods in which the indicator issued a bad signal or "noise". For each indicator, KLR find the "optimal" threshold, defined as that threshold that minimizes the noise-to-signal ratio, N/S. Eventually, a composite indicator is constructed as an average of indicators, weighted by frequency of correct predictions.

Following KLR, academics and economists working in the private sector produced a wide range of currency crisis forecasting models. Unlike the indicator approach, most are variations of logit or probit regressions. It is noteworthy that the vast majority of such models, including the KLR model, use fundamental data such as current account, exchange rate overvaluation or export growth as explanatory variables. The variable choices are predominantly inspired by the three generations of theories of balance-of-payment crises, but they tend to be limited by availability

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of data. The academic models tend to be long-sight approaches with forecasting horizons of up to two years, while their private-sector counterparts usually focus on a brief windows of one to three months (the Deutsche Bank's DB Alarm Clock, for instance, has a horizon of one month, while the prediction horizon of the original KLR contribution is of 24 months).

Berg et al. (2004) not only give a very helpful overview of the abundant literature on EWS, they also address the question of how much, if any, out-of-sample forecasting value derives from EWS. They put particular emphasis on the potential performance of models in real time and reach rather disappointing conclusions. Only one of the long-horizon forecasts under consideration (the KLR model forecast) provides better accuracy compared with pure guesswork and non-model based predictions, while short-horizon private-sector approaches by and large perform poorly. In contrast, Anzuini and Gandolfo (2003), when testing whether the KLR would have forecast the Thai crisis of 1997, conclude the indicator approach has strong ex-post explanatory power but quite limited predictive abilities.

In this paper, we propose exploiting the term structure of relative interest rates to obtain estimates of changes in the timing of a currency crisis as perceived by market participants. To our knowledge, only a handful of researchers have used this approach. The essence of our model is based on the seminal work by Collins (1984), who applied her analysis to the March 1983 devaluation of the French franc relative to the German Deutsche mark. The Collins approach was also used in Anzuini and Gandolfo (2003), who compare the predictive power of the KLR approach and the Collins model. They conclude that the Collins non-structural approach forecasts well but does not explain, while the opposite is true for the structural KLR model. This approach does not rely on the estimation of thresholds (eventually common to a group of countries) based on fundamentals, but instead extracts expectations on the timing of the crisis from country-specific interest rate data. Compared with the KLR and other EWS, this approach has several important advantages. First, it requires no definition of a crisis in terms of percentage devaluation/depreciation, which tends to be rather arbitrary in the literature. Moreover, no pooling of data is necessary to obtain a sample of a usable size. For each country in question, only its own specific data may be used. In addition, as the model uses only very basic data such as interest and exchange rates; it is not heavily limited by data availability. We construct an early warning indicator that can be used to evaluate the relative probability of a crisis occurring in one week as compared to a crisis happening after one week but in less than a month. Subsequently, we provide empirical evidence

that the indicator performs well for the currency crises in the Czech Republic in 1997 and in Russia in 1998.

This chapter is structured as follows. In the next section the theoretical model is developed. In section 1.3 we apply the indicator to the crisis of the koruna and the ruble in 1997 and 1998, respectively. Section 1.4 concludes and sets up paths of further research.

1.2 Uncovered interest rate parity and currency crises

Recent research on the predictive power of markets suggests that markets are capable to aggregate disperse information and that market-based forecasts are usually fairly accurate. Moreover, such forecasts typically outperform alternative forecasting tools, including highly sophisticated forecasting models, polls or expert surveys (see e.g. Wolfers and Zitzewitz, 2004). The basic objective of our analysis is to examine the ability of foreign exchange markets to foresee exceptional exchange rate devaluations in fixed exchange rate arrangements. In substance, the following model is based on the work by Collins (1984) (see also Anzuini and Gandolfo, 2003), designed to study the behavior of speculators prior to the French franc realignment in 1983. Our theoretical setting extends and generalizes the original one in several aspects. On the one hand, we explicitly take into account the potential existence of a time-varying risk premium, assumed constant in the references above. On the other hand, we also allow for the possibility of appreciation expectations in order for the approach to be usable for exchange rates in target zones or other types of intermediate exchange rate regimes.

The aim of our study is to construct an indicator based on basic economic theory (the uncovered interest rate parity, henceforth UIP) to proxy the change in the time structure of the underlying expected probabilities of devaluation implied by the relative term structure of interest rates. Accommodating risk aversion, the uncovered interest rate parity can be modified as

$$\frac{(1 + i_{t,k})}{(1 + i_{t,k}^*)} = \frac{E(e_{t+k}|\Omega_t)}{e_t} + \rho_{t,k} \tag{1.1}$$

where e_t is the spot exchange rate at time t, defined as the price of foreign currency in domestic currency units, $i_{t,k}$ and $i_{t,k}^*$ are, respectively, the domestic and foreign interest rates at time t on deposits with maturity k. $E(e_{t+k}|\Omega_t)$ stands for the expected exchange rate in period $t + k$ given the information available at time t (the

information set Ω_t) and $\rho_{t,k}$ represents a premium for risks not immediately related to the exchange rate movements (e.g. country default risk).[3]

Equation (1.1) states that the relative yield on domestic deposits of a given maturity is equal to the expected exchange rate movement and some well defined country risk premium. Rewriting (1.1), we obtain

$$E(e_{t+k}|\Omega_t) = \left[\frac{(1+i_{t,k})}{(1+i_{t,k}^*)} - \rho_{t,k} \right] e_t. \qquad (1.2)$$

From the perspective of the market agent forming expectations in time t the exchange rate can either remain stable, appreciate or depreciate. Therefore, the exchange rate expected as of t for the period $t+k$ is a weighted average of these scenarios, where the weight assigned to each possible exchange rate movement is the subjectively perceived probability of these events. Formally, this implies that

$$E(e_{t+k}|\Omega_t) = (1 - \pi_{t,k})s_{t,k} + \pi_{t,k}z_{t,k} \qquad (1.3)$$

where $z_{t,k}$ is the expected exchange rate in period $t+k$ in case of devaluation, $z_{t,k} = \gamma_{t,k}e_t$, where $\gamma_{t,k} > 1$, and $s_{t,k}$ is the expected exchange rate conditional on no devaluation (in other words, the exchange rate remains stable or appreciates), $s_{t,k} = \delta_{t,k}e_t$, where $\delta_{t,k} \leq 1$. The subjective probability of devaluation having occurred after k periods is therefore $\pi_{t,k}$. In addition, along the lines of Collins (1984), we shall assume that the rate of depreciation or appreciation does not depend on the temporal horizon, so that $\delta_{t,k} = \delta_t$ and $\gamma_{t,k} = \gamma_t$. It follows that

$$E(e_{t+k}|\Omega_t) = \left[\delta_t + (\gamma_t - \delta_t)\pi_{t,k} \right] e_t, \qquad (1.4)$$

which can be substituted in (1.2) so as to establish the link between relative yields and the subjective devaluation probability:

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[3]In our specification the risk premium is assumed to be unrelated to the exchange rate. In certain cases, it may be reasonable to assume that the exchange rate level has an effect on the risk premium. For instance, if domestic debt is denominated mostly in the currency of the foreign country, one might expect that after a substantial devaluation the probability of default and thus the risk premium in the domestic country would rise. This effect can be easily incorporated in the model by specifying a functional form linking the exchange rate to the risk premium. If the relationship is assumed to be linear in the exchange rate with a slope that is not maturity-dependent, the results presented in this section remain unchanged. For a general functional specification of this relationship, the model could still be applied after calibrating some extra parameters. We do not follow this avenue in the present study, although this generalization is proposed as an interesting future path of research.

$$\alpha_{t,k} = \delta_t + (\gamma_t - \delta_t)\pi_{t,k} + \rho_{t,k}, \tag{1.5}$$

where $\alpha_{t,k} = (1+i_{t,k})/(1+i_{t,k}^*)$. The essential term in equation (1.5) is the perceived probability of a devaluation between time t and $t+k$, $\pi_{t,k}$. Anticipating the empirical application of the method, we will restrict ourselves to devaluations occurring within each of the time intervals corresponding to the maturities of the available time deposits. If there are $J-1$ different maturities of deposits ordered from the shortest to the longest, there are J possible states of the world at time t. An exceptional devaluation might occur before the time implied by the shortest maturity available, between maturities of deposits k and $k+1$ or, finally, there might be a devaluation after the longest deposit matures. Defining as $\eta_{t,j}$ the probability of a devaluation happening between period $t+j-1$ and $t+j$,

$$\pi_{t,k} = \sum_{i=1}^{k} \eta_{t,i}, \tag{1.6}$$

which implies that

$$\frac{\alpha_{t,k} - (\delta_t + \rho_{t,k})}{(\gamma_t - \delta_t)} = \sum_{i=1}^{k} \eta_{t,j}. \tag{1.7}$$

For the sake of illustration let us suppose that there are two maturities $k = 7$ and $k = 30$ days (this will be the case in the empirical illustration in the following section). Then, it can be easily shown that

$$\eta_{t,7} = \frac{\alpha_{t,7} - (\delta_t + \rho_{t,7})}{(\gamma_t - \delta_t)}, \tag{1.8}$$

$$\eta_{t,30} = \frac{\alpha_{t,30} - \alpha_{t,7} - (\rho_{t,30} - \rho_{t,7})}{(\gamma_t - \delta_t)}. \tag{1.9}$$

By taking ratios or log-ratios of the expressions above, we can identify changes in the time structure of subjective probabilities of a devaluation implied by the term structure of interest rates. Furthermore, the ratios are independent of the assumed size of the devaluation, γ_t. Assume an ordering of maturities, where the shortest one is normalized to one, the second shortest is two, and so on. In particular, the indicator proposed, $I_{j,s,t}$, aimed at comparing the probabilities of devaluation at horizons j and s ($j < s$), corresponding to two observed maturities, is

$$I_{j,s,t} = \begin{cases} \log\left(\frac{\alpha_{t,j}-\alpha_{t,j-1}-(\rho_{t,j}-\rho_{t,j-1})}{\alpha_{t,s}-\alpha_{t,s-1}-(\rho_{t,s}-\rho_{t,s-1})}\right) & \text{for } j,s > 1, \\ \log\left(\frac{\alpha_{t,j}-(\delta_t+\rho_{t,j})}{\alpha_{t,s}-\alpha_{t,s-1}-(\rho_{t,s}-\rho_{t,s-1})}\right) & \text{for } j = 1, s > 1. \end{cases} \tag{1.10}$$

In order to make (1.10) operational, the expected appreciation parameter, δ_t and the respective risk premia for each maturity need to be imputed. If we assume that $\delta_t = 1$ and $\rho_{t,i} = 0$ $\forall i$, (1.10) boils down to the expression put forward in Collins (1984). The problem with this setting is that it can lead to negative probability estimates for empirical applications, and thus log-ratios which are not defined. The inclusion of risk premia and potential appreciation expectations in (1.10) allows us to elaborate corrections of the basic indicator in order to avoid negative probability ratios.

The problem of negative probability ratios is particularly important when dealing with data from Eastern European transition economies. Taking the simple case without risk premium (i.e., setting $\rho_{t,s} = 0$ $\forall s$ above) and $\delta_t = 1$, it can be easily seen that for a relatively flat yield curve in the domestic economy, if $i_{t,7}$ tends to be higher than $i_{t,30}$, negative values can be obtained in the numerator of (1.9). This constellation, caused by a downward-sloping yield curve in the domestic economy, is not unusual in the recent history of Eastern European economies, where sustained disinflationary experiences rendered a term structure of interest rates with lower nominal interest rates in longer maturities.

A simple correction to the simple setting based on future expected inflation can be put forward to link the setting including risk premium to a yield curve which is potentially negatively sloped. In a disinflationary framework, with $E(\Delta p_{t,30}|\Omega_t) < E(\Delta p_{t,7}|\Omega_t)$, where $\Delta p_{t,g}$ is the inflation rate for the period t to $t+g$, we can correct the interest rate with longer maturity (we denote the corrected rate by $i_{t,30}^{corr.}$) by substracting $(E(\Delta p_{t,30}|\Omega_t) - E(\Delta p_{t,7}|\Omega_t))$ from the original rate, so that the corrected numerator of equation (1.9) for the case without risk premium is given by

$$\begin{aligned} \alpha_{t,30}^{corr.} - \alpha_{t,7} &= \frac{1+i_{t,30}^{corr.}}{1+i_{t,30}^*} - \frac{1+i_{t,7}}{1+i_{t,7}^*} = \\ &= \frac{1+i_{t,30} + E(\Delta p_{t,7}|\Omega_t) - E(\Delta p_{t,30}|\Omega_t)}{1+i_{t,30}^*} - \frac{1+i_{t,7}}{1+i_{t,7}^*} = \\ &= \frac{1+i_{t,30}}{1+i_{t,30}^*} - \frac{1+i_{t,7}}{1+i_{t,7}^*} + \frac{E(\Delta p_{t,7}|\Omega_t) - E(\Delta p_{t,30}|\Omega_t)}{1+i_{t,30}^*} = \end{aligned}$$

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$$= \alpha_{t,30} - \alpha_{t,7} + \frac{E(\Delta p_{t,7}|\Omega_t) - E(\Delta p_{t,30}|\Omega_t)}{1 + i^*_{t,30}}.$$

In other words, $\rho_{t,7}$ and $\rho_{t,30}$ in equation (1.9) may be interpreted as, respectively, $\frac{E(\Delta p_{t,7}|\Omega_t)}{1+i^*_{t,30}}$ and $\frac{E(\Delta p_{t,30}|\Omega_t)}{1+i^*_{t,30}}$ if the basic setting is employed and long-maturity interest rates are corrected for expected disinflation. In practice, this correction could be carried out for maturities of 7 and 30 days , for instance, by replacing $E(\Delta p_{t,7}|\Omega_t)$ with the realized inflation level at time t and using a time series model in order to obtain forecasts for $\Delta p_{t,30}$. In our empirical application we report the results of such a correction based on inflation forecasts.

1.3 The warning mechanism in action in the Czech and Russian currency crises

In this section, we apply the indicator put forward above to data from two recent currency crises in Eastern Europe: the Czech Republic crisis in May 1997 and the Russian crisis in 1998.[4] In both cases, we describe the economic framework in which the currency crises took place and present the real-time estimates of our indicator for both economies during the crisis period.

1.3.1 The 1997 crisis in the Czech Republic

In the early 1990s, the Czech Republic introduced a tight peg of the koruna to the Deutsche mark (DM) and the US dollar (USD). The currency basket used was made up of 65% DM and 35% USD from May 1993. The peg had fluctuation bands of $\pm$ 0.5% up to February 1996 and $\pm$ 7.5% from February 1996 until the May 1997 crisis, which materialized with the abandonment of the peg on May 26, 1997 for a managed float regime. The trade balance in the Czech Republic, which had been systematically positive since the break-up of Czechoslovakia, turned negative in 1996, with a corresponding slowing of economic growth. Horváth (1999) interprets the current account deficit in the Czech Republic as a reflection of insufficient private savings, which, coupled with the institutional framework of the Czech banking sector at that time, made the deficit unsustainable. Furthermore, the real exchange rate appreciated persistently and continuously in the period 1992-1997. Although trend appreciation is a common phenomenon in transition economies, which can be (at least partly) explained through the Balassa-Samuelson effect by differential pro-

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[4]The choice of the crises is exclusively based on data availability.

ductivity increases, Begg (1998) and Horváth (1999) argue that the real exchange rate dynamics implied a loss of competitiveness of the Czech economy. The adverse macroeconomic framework, together with an unstable political environment, led to a speculative attack on the koruna and a change in the exchange rate regime in May 1997.
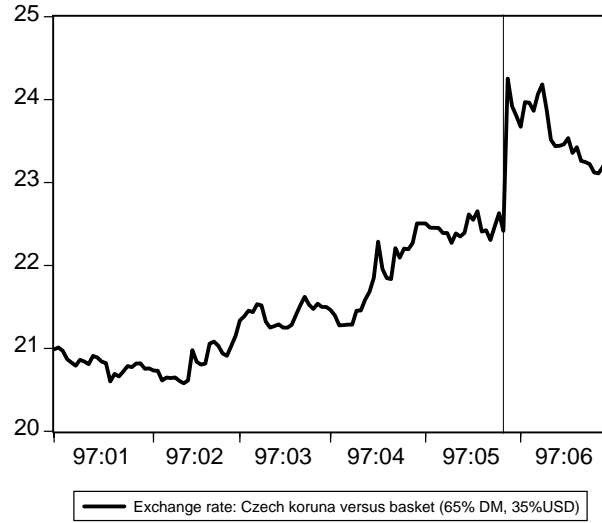
Figure 1.1: Czech Koruna: Exchange rate, January 1997-June 1997

Figure 1.1 shows the daily exchange rate of the Czech koruna against the basket in the period January-June 1997. The vertical line corresponds to the abandonment of the peg. As it can be seen in Figure 1.1, the exchange rate remained inside the ± 7.5% bands during the turmoil preceding the change in the exchange rate regime, and depreciated strongly as soon as the managed float regime was in place. The fact that the monetary authorities were able to keep the koruna inside the fluctuation bands was mainly due to the heavy central bank interventions taking place in the week preceding the breakdown of the peg (see Horváth, 1999).

We calculated the indicator given by (1.10) using two different maturities for the Czech koruna exchange rate against both the DM and the USD for the period ranging from January 1st, 1997 to the abandonment of the peg on May 27th, 1997. We used the daily interbank rates with maturity one week ($i_{t,7}$ and $i^*_{t,7}$) and one month

($i_{t,30}$ and $i^*_{t,30}$) for the Czech Republic and, alternatively, Germany and the US.[5] The main reason behind the choice of these maturities for our indicator is based on exploiting differences in interest rates in the very short run, a time domain which has not been assessed in the literature on currency crises hitherto.[6]

The yield curve implied by the term structure of interbank rates in the Czech Republic is downward-sloping for most of the sample. If we were to obtain an indicator based on the assumptions imposed in Collins (1984) (that is, imposing $\delta_t = 1$ and $\rho_j = 0$ $\forall j$ in (1.10)), the results would imply negative values in the argument of the log-ratio corresponding to $\eta_{t,30}$, since $\alpha_{t,30}$ tends to be systematically smaller than $\alpha_{t,7}$ for the sample at hand. A possible way to overcome this problem would be to redesign the log-ratio of probabilities by adding a constant to the numerator and denominator of the expressions in the log of (1.10) after setting $\rho_{t,i} = 0$ for all i. In our setting, this can be reconciled with the existence of a certain maturity structure in the risk premium, such that, for example, $\rho_{t,j} - \rho_{t,j-1} = \rho_{t,s} - \rho_{t,s-1} = c < 0$ $\forall j > 1, j < s$. For the indicator related to the shortest maturity, appreciation expectations (so that $\delta < 1$) can also lead to the same type of correction. In Figure 1.2 we show the resulting indicator after adding 0.3 to the denominator of the expression for $I_{7,30,t}$ in (1.10), so as to keep the structure of relative changes in the original estimates of $\eta_{t,7}$ and $\eta_{t,30}$ but avoid negative relative probabilities. The results presented in Figure 1.2 correspond to using the US as the foreign economy, but are identical to those using Germany. The results are also qualitatively identical for constants different from 0.3, as long as they avoid negative values in the argument of the log ratio. Furthermore, the results are also similar if the indicator is constructed under the assumption that $k = 7$ is not the shortest maturity.

Changes in the indicator can be interpreted as changes in the perceived probability of a crisis occurring in the following week as compared to a crisis happening in the period delimited by day seven and day thirty. The indicator remains practically constant from January to mid-May, and starts increasing dramatically on May 16th, reflecting a strong change in the perceptions of investors on the potential timing of a devaluation. The increase is strong and sustained until May 28th, and from that day onwards the indicator slowly decreases to a low level, comparable with the pre-crisis period. The indicator performs therefore extraordinarily as a (very-)short-term in-

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[5]The source of the data used in this study is Datastream.

[6]The results are mostly similar if different pairs of short-term interest rates are used. The resulting indicator using maturity combinations of 7, 30 and 90 days presents similar properties to those presented in this section. These results are available from the authors upon request.
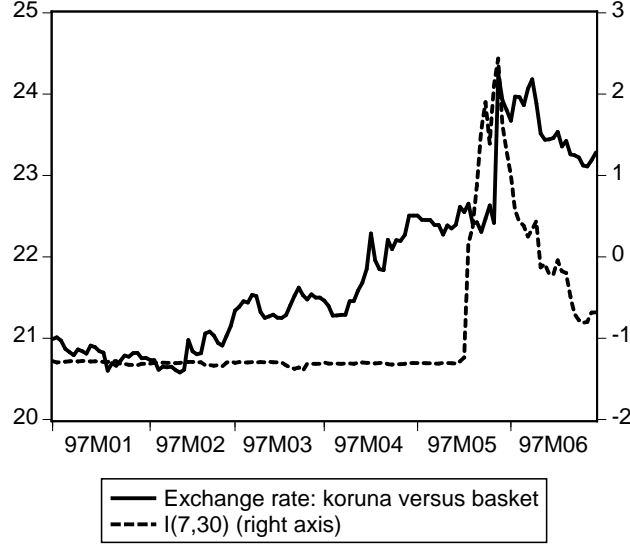
Figure 1.2: Czech Koruna: Exchange rate, January 1997-June 1997 and $I_{7,30,t}$

dicator of the crisis, and could be used ex-post as a device for dating the *de facto* occurrence of the crisis (since, *de jure*, the change of the exchange rate regime would be the corresponding indicator).

Alternatively, we also made use of the daily one month forward rates $(e^f_{t,30})$ for the koruna/USD exchange rate, which are available for the period under study, in order to get real-time estimates of $\rho_{t,30}$ in the framework of the covered interest rate parity, given by

$$\rho_{t,30} = \frac{(1 + i_{t,30})}{(1 + i^*_{t,30})} - \frac{e^f_{t,30}}{e_t}. \tag{1.11}$$

Using these estimates we can include the dynamics of the risk premium in the indicator. Since due to lack of data we do not have estimates of $\rho_{t,7}$ for the pre-crisis period, let us assume that the dynamics of $\rho_{t,7}$ are similar to those of our estimate of $\rho_{t,30}$, although the level may be different, so that we assume $\rho_{t,30} < \rho_{t,7}$ in order to avoid negative probabilities, keeping the assumption $\delta_t = 1$. [7] The resulting indicator is plotted in Figure 1.3. Although the long-run dynamics of the indicator are not affected by this correction, it should be noted that the resulting estimate of

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[7]Seven-days forward exchange rates for the crisis period (but not for the pre-crisis period) are available, and confirm that the resulting risk premium for seven days was higher than $\rho_{t,30}$.

the risk premium for the Czech Republic is far from being constant for the period considered. In particular, the speculative attack that lead to the crisis is identified as a strong increase of the risk premium. It was therefore not possible to foresee in advance if the inclusion of this correction would affect the properties of the indicator. We also computed the corresponding indicator for the German case, using a synthetic forward rate (since forward rates for the koruna/DM are not available for the period) obtained from other forward cross-rates and the resulting graph is similar to Figure 1.3.[8]
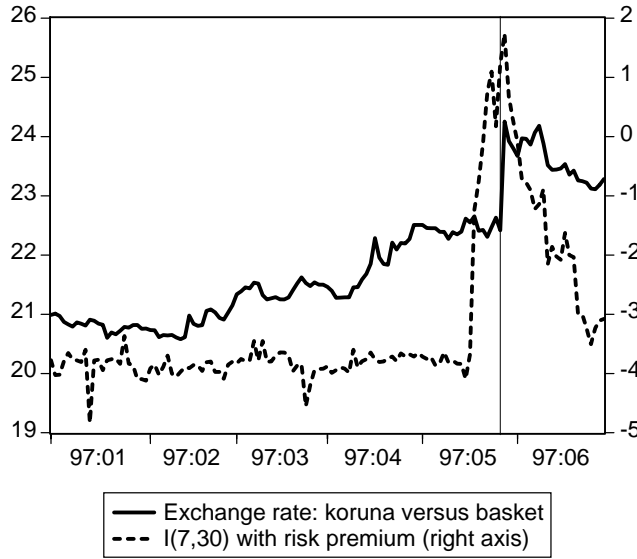
Figure 1.3: Czech Koruna: Exchange rate, January 1997-June 1997 and $I_{7,30,t}$ with risk premium adjustment

The results concerning the leading indicator properties of the log-ratio are qualitatively similar to those obtained without the risk premium adjustment, although the indicator series is now relatively more volatile in the pre-crisis period.

We also performed the correction based on inflation expectations as follows. We used the realized inflation rate at time t as a proxy for $E(\Delta p_{t,7}|\Omega_t)$, and for each period we estimated different models in order to obtain forecasts of $\Delta p_{t,30}$ (Δp_{t+1} in monthly notation), $E(\Delta p_{t,30}|\Omega_t)$ using data up to time t. This correction is not

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[8]Detailed results are available from the authors upon request.
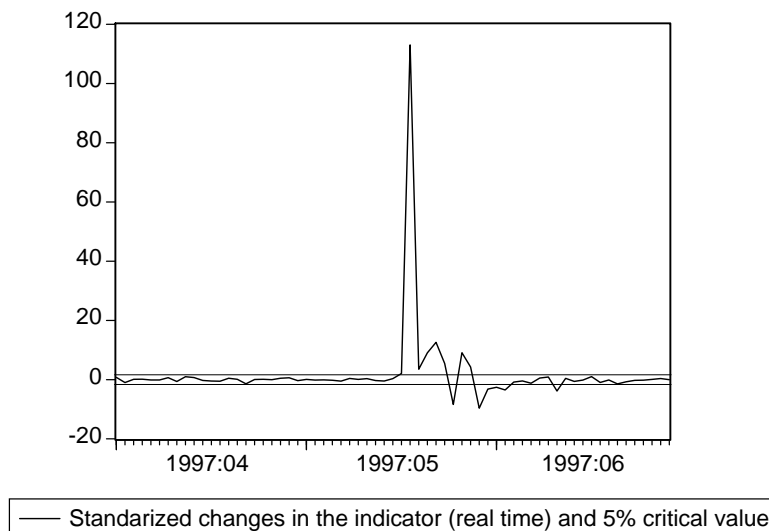
Figure 1.4: Standardized $I_{7,30,t}$ changes: Czech Republic

able to overcome the negative probabilities if simple autoregressive processes with and without deterministic trends are used as forecasting models. Since these parsimonious models tend to forecast future inflation relatively well, the correction is not able to render positive values of $\eta_{t,30}$ unless the data generating process assumed for inflation contains nonlinear deterministic trends that systematically produce strong disinflationary forecasts.[9]

Until now, no reference has been made to the size of the change in the indicator leading to a crisis signal. While several methods can be used to evaluate the threshold leading to significant signals, an extremely simple one based on the standardization of changes in the indicator seems to perform well. In Figure 1.4 we present the changes in the indicator at each period t standardized using the average change and standard deviation realized up to period $t - 1$. We start the exercise in April 1997 based on indicator changes ranging from January 1997 and we also plot the 5% critical values corresponding to a standard normal distribution. The first significant change takes place on April 16th, and no false signal is sent before the crisis.

While the aim of the indicator proposed is to serve as a short-term leading indicator for exchange rate crises, we study whether also in tranquil times relative changes

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[9]Detailed results on the correction are available from the authors upon request.

of the indicator contain information about future changes in the exchange rate. This will be done by performing a simple out-of-sample forecasting exercise for the pre-crisis sample. The forecasting abilities of an autoregressive model on the first difference of the (log) exchange rate of the Czech koruna against the USD and the DM will be compared to those of a simple vector autoregressive (VAR) model including changes in the exchange rate and the indicator with risk premium adjustment.

The forecasting exercise is carried out as follows. Using data from January 1st to April 1st, 1997, an autoregressive model is estimated for the log changes in the exchange rate, together with a VAR for the vector of log changes in the exchange rate and the first difference of the indicator. In both cases, the lag length of the model is chosen so as to minimize AIC for the sample. Using the estimated models, out-of-sample forecasts are obtained for 1 to 30 (working) days ahead, and the forecasting errors are computed by comparing the forecasts with the real data. The observation corresponding to April 2nd is added to the sample, the models are estimated again and new out-of-sample forecasts are obtained. This procedure is repeated until all available observations have been used. In our case, since we are interested in the informational content of the indicator in the pre-crisis period, the full sample used ranges from January 1st to May 20th. With the forecasts, we compute the root mean square forecasting error (RMSFE) for each forecasting horizon, defined as $\text{RMSE}_h = \sqrt{\sum_{n=1}^{N_h}(e_n - \hat{e}_n)^2/N_h}$, where N_h is the number of h-steps ahead forecasts computed, e_n is the actual value of the exchange rate and $\hat{e}_n$ is the corresponding forecast.

Table 1.1 presents the results of the forecasting exercise. For each exchange rate considered, the improvement of RMSFE for the VAR model over the simple autoregression is presented for different forecasting horizons. The results of the corresponding Diebold-Mariano test (Diebold and Mariano, 1995) for equality of forecasting ability is also reported in each case. The forecasting abilities of the model including information on the indicator are superior to those of the autoregressive model for all forecasting horizons in the case of the US dollar, with marginal improvements on the forecasting error averaging 0.75% over the 30-day forecasting horizons. Although the improvement is very modest, it should be pointed out that the particular exchange rate regime of the Czech Republic for the period considered limited significantly the volatility of exchange rate movements for the period. The improvements are furthermore statistically significant for forecasting horizons of 5-days ahead and longer. The results are not so positive for the DM, where our indicator does not

seem to possess information on future exchange rate developments for quiet periods. For the one-month forecast horizon the model including the indicator improves in RMSFE over the simple AR model, albeit not significantly.

Exchange rate: Czech koruna/USD			
Steps ahead	RMSFE difference	Diebold-Mariano test	Observ.
1	-1.271 %	-0.705	34
5	-0.998 %	-1.845**	30
10	-0.408 %	-1.687**	25
15	-0.686 %	-1.955**	20
20	-0.757 %	-2.496***	15
25	-0.773 %	-1.370*	10
30	-1.267 %	-3.474***	5
Exchange rate: Czech koruna/DM			
Steps ahead	RMSFE difference	Diebold-Mariano test	Observ.
1	0.156 %	1.053	34
5	0.062 %	0.866	30
10	0.052 %	0.679	25
15	0.050 %	0.787	20
20	0.027 %	0.509	15
25	0.066 %	0.867	10
30	-0.044 %	-0.967	5

The column "RMSFE difference" is the difference between the RMSFE of the VAR model and the AR model as percentage of the RMSFE of the AR model. The column "Diebold-Mariano test" refers to the Diebold-Mariano test for equal forecasting error (Diebold and Mariano, 1995). *(**)[***] stands for significance at the 10%(5%)[1%] level.

Table 1.1: Pre-crisis forecasting exercise: VAR vs. AR

1.3.2 The 1998 crisis in Russia

The Russian Central Bank announced in November 1997 that, starting January 1998, the ruble would be targeted at a central rate of 6.2 rubles/dollar, with a fluctuation band of $\pm$ 15%. However, the volatility of the ruble/dollar exchange rate was minimal in the months preceding the crisis (the standard deviation of percentage changes in the exchange rate was 0.002 in the period January-August, 1998). On August 17th, 1998, the Russian government announced the devaluation of the

ruble by the end of the year, defaulted on its government debt and declared a 90-day moratorium on foreign debt. On August 26th the Russian Central Bank declared that the fixed exchange rate could not be supported any longer and on September 2nd, 1998 the Russian ruble was floated.[10]

Using the corresponding interbank interest rate data for Russia and the US, we construct the indicator for the dynamics of the relative probability of a crisis occurring in seven days as compared to the crisis taking place in the interval delimited by seven and thirty days. The same problems as for the case of the Czech Republic come up if the indicator proposed by Collins (1984) is used, since the probability ratio turns negative in some periods due to the downward-sloping term structure of Russian interbank rates. Figure 1.5 presents the Russian ruble/USD exchange rate together with the indicator $I_{7,30,t}$ for the period April-September, 1998, after assuming $\delta_t = 1$, $\rho_7 = 0$ and adding a constant (one in this case) to the denominator of the expression in (1.10) so as to avoid negative implied probabilities. The shaded area delimits the period of time starting with the announcement of the devaluation and ending with the floatation of the ruble.
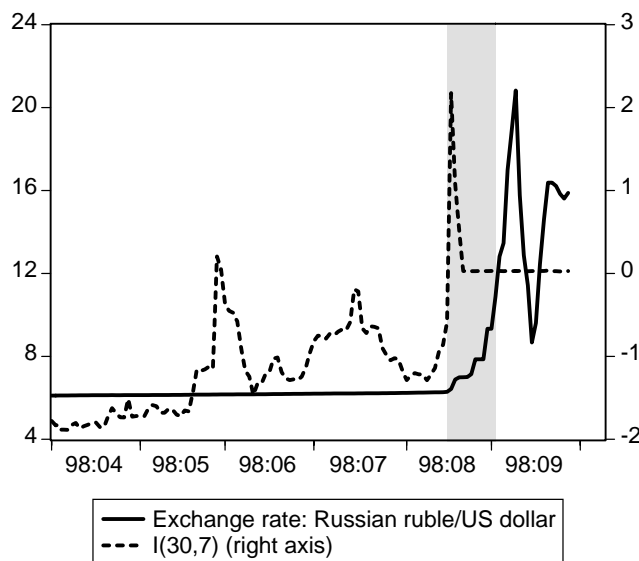
Figure 1.5: Russian ruble: Exchange rate, April 1998-October 1998 and $I_{7,30,t}$

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[10]For an excellent account of the Russian crisis, see Kharas et al. (2001)

The first relevant feature of $I_{7,30,t}$ is the fact that it has a positive trend in the period under study. This implies that investors systematically changed their expectations of the timing of an exchange rate crisis in the months preceding the actual occurrence of the Russian crisis. In this sense, as the crisis approached, they tended to consider the event increasingly imminent. Apart from this medium-run trend in $I_{7,30,t}$, the indicator presents relevant increases in the end of May, mid July and a global peak following the announcement of the devaluation, which precedes the change in the exchange rate regime by seven (working) days. The first peak, on May 28th, takes place right after the Central Bank increased key interest rate to 150 % and is followed by a series of interventions in the coming days (involving the expenditure of \$1 billion in reserves) in a successful attempt to defend the ruble (see e.g. Chiodo and Owyang, 2002). The indicator declines in the following days, and follows the positive trend that dominates the full period. The second signal of a shift in expectations to an imminent devaluation takes place starting in early July. The start of the increase coincides with the Russian parliament's postponement of the policy reforms needed to qualify for IMF loans. Expectations of crisis timing shift away from one week with the final approval of an IMF emergency loan to Russia in mid-July. Finally, our indicator increases dramatically in the period August 10-18 in parallel to the collapse of the stock and bond markets (August 13) and in spite of Boris Yeltsin's declarations that "there will be no devaluation" of the ruble following an emergency parliamentary session on August 14. Our indicator only stabilizes on August 21, when the Russian crisis can already be felt in markets all around the world. Although our indicator peaks when the crisis is already being felt, the increase in $I_{7,30,t}$ is strong up to seven days before the devaluation announcement. The dynamics of the indicator remain unchanged if the inflation expectation correction is carried out, or if the 90-day interest rate is used as the long edge of the maturity comparison. [11]

Figure 1.6 presents the standardized changes in the indicator computed with information up to period t for each observation. The indicator data used for the standardization starts in January 1997, and the values corresponding to the sample under study are presented in the figure together with the 5% critical values. The two peaks of the indicator which did not result in a crisis do not actually appear significant at the 5% level, while the crisis signal does.

The minimal volatility of the ruble/US dollar exchange rate in the pre-crisis period makes an out-of-sample exercise such as the one carried out for the Czech Republic

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[11]Detailed results are available from the authors upon request.
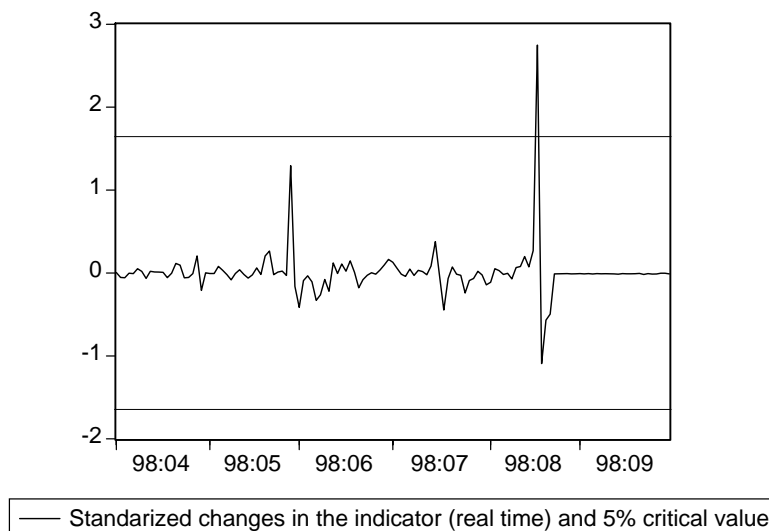
Figure 1.6: Standardized $I_{7,30,t}$ changes: Russia

useless in this context. To sum up, our indicator is able to identify both speculative pressures that were successfully combated by the central bank in the pre-crisis period, and starts signalling the occurrence of the impending crisis six days before the official announcement of the devaluation.

1.4 Summary and conclusions

The increased frequency and strength of currency crises in recent years has motivated researchers in both public and private institutions to develop effective early-warning systems for currency crises. A vast majority of existing approaches uses similar macroeconomic variables to forecast the timing of financial distress. In our opinion, fundamental data are perfectly suited to identifying the set of potentially vulnerable countries and, possibly, to explaining crises after the fact. However, in our opinion, the desired forecasting instrument needs to focus strongly on market sentiment as it is the participants on foreign exchange markets who eventually trigger a crisis. Investor sentiment is much more sensitive to short-term news and incoming signals than to underlying long-term fundamentals. Along these lines, recent research suggests that market-based forecasting tools possess fair predictive power and usually outperform alternative instruments in terms of accuracy.

Thus, based on simple economic theory and exploiting the term structure of relative interest rates, we constructed a very short-term early-warning indicator to evaluate relative probabilities of a crisis occurring in different time horizons. Subsequently, we applied the indicator to data from two recent Eastern European currency crises: the Czech koruna crisis in 1997 and Russia ruble crisis in 1998. We found that our indicator performs extraordinarily as a (very) short-term predictor of a crisis in both considered cases. We also provided evidence that the indicator contains extra information about future short-run exchange rate changes.

In principle, the indicator put forward can be interpreted as a *dating* instead of a *predicting* mechanism for currency crises. While at least part of the dynamics preceding the breakdown of the peg in both cases studied are to be interpreted as attempts of the respective central bank to defend its currency, the econometric analysis carried out in the paper shows however that the dynamics of the indicator contains useful information about future movements of the exchange rate. In that sense, our indicator can be seen as a measure of stress in the foreign exchange market. This is better illustrated in the case of the Russian crisis, where a positive trend in the behaviour of the indicator can be interpreted a a continuous increase of the crisis risk, and the medium-run dynamics is able to identify speculative attacks prior to the crisis.

As seen in the Russian case study, even false alarms from the market need to be taken seriously by central bankers and governments and all available short-term measures should be implemented. Likening our indicator to a thermometer, a high body temperature does not necessarily imply serious illness, but is always a reason for concern. In that sense, our indicator is useful for monetary policy institutions as an extra signalling instrument to complement long-run warning mechanisms. Several improvements to the methodology used in this piece of research can be implemented to refine the indicator. Among possible avenues of research, using information on the time-varying nature of interest rate volatility to proxy for developments in the risk premium may lead to improvements in the signalling properties of the estimator.

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Chapter 2

On the Determinants of Currency Crises: The Role of Model Uncertainty[1]

2.1 Introduction

Over the course of the last couple of decades several parts of the world have experienced rather harsh financial market crises, sometimes repeatedly, and mostly accompanied by painful real shocks. The very last wave of such turmoil, initially triggered on the US (subprime) mortgage market, has exemplified that financial market turbulences are not confined only to the developing and emerging economies. Moreover, the recent tensions have clearly unveiled challenges financial stability authorities and policy makers have to face in the age of ever deeper and more global markets. Most importantly, diminishing barriers to capital flows and instant information distribution increase the potential sudden evasiveness of capital. As evidenced by the shocking promptness with which the US mortgage malaise extended from one corner of the financial market to another, crises can spread swiftly between different types of markets in geographical and technical terms.

One of the most frequent targets of speculators is the currency market and substantial devaluations of the currency under attack generally imply severe consequences for the respective economy. Against this backdrop it is not surprising that both in the academic literature and in the private sector a variety of empirical attempts has been undertaken to predict currency crises. Following the pioneering indica-

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[1]This chapter is based on a joint work with my adviser, Prof. Dr. Jesús Crespo-Cuaresma.

tor approach by Kaminsky, Lizondo and Reinhart (1998) a whole plethora of early warning systems for currency crises has been developed. Some of the rather recent approaches employ innovative methodologies such as Markov switching models (see e.g. Abiad, 2003 or Chen, 2005) or financial market tools (see e.g. Malz, 2000 or Crespo-Cuaresma and Slacik, 2007) to predict currency attacks.

The vast majority of the empirical literature assesses the effect of various potential determinants on the probability of a currency crisis using limited dependent variable - logit or probit - models. The discrete crisis variable is regressed on a set of fundamental indicators, such as, inter alia, current account and government balances, exchange rate overvaluation or liquidity ratios . The choice of regressors is typically inspired by the three generations of theoretical models on balance-of-payment crises. In one of the most recent empirical contributions on this topic Bussière (2007) overhauls the usually static specification, in which, moreover, all regressors tend to enter at the same lag. He thus extends the usual set of explanatory variables by including several lags of the regressors as well as of the dependent binary crisis variable. He finds that there are several variables significantly affecting the probability of a crisis in a dynamic logit model. However, the impact of the indicators ranges between short-run (4-6 months) e.g. for the liquidity measures to very long-run (2 years) in case of over-appreciation of the exchange rate. In addition, his results indicate that past crisis episodes increase the probability of a new attack, particularly in the short run.

Notwithstanding substantial variations in the literature on early warning systems with respect to methodology, data as well as results, there is one general caveat which applies to all existing binary choice models. Given that there is no unique theoretical framework linking the potential set of determinants with the realizations of currency crises, the issue of model uncertainty surrounding both the choice of variables and the estimates obtained deserves to be treated seriously. Model uncertainty can be explicitly taken into account using Bayesian statistical techniques, in particular with the use of the Bayesian model averaging (BMA) methodology which proposes averaging of the parameter values over all (relevant) alternative models using posterior model probabilities as respective weights to evaluate the relative importance of different variables (see Raftery, 1995 for a general discussion and Sala-i-Martin *et alia*, 2004, Fernandez *et alia*, 2001, or Crespo-Cuaresma and Doppelhofer, 2007 for applications to economic growth regressions).

The different theoretical settings used to explain different crises episodes give rise to

alternative sets of potential explanatory variables (with intersections which are not necessarily empty) for the probability of a crisis occurring. The so-called *first generation* models (Krugman, 1979, Flood and Garber, 1984) concentrate on bad economic policy leading to unsustainable developments of some fundamental macroeconomic variables. The abandonment of the fixed exchange rate regime is then precipitated by the eventual exhaustion of the central bank's foreign reserves. The *second generation* of currency crises models (see for instance Obstfeld, 1994), explains crises as the consequence of self-fulfilling expectations in theoretical settings with multiple equilibria. In contrast, the *third generation* of models (Krugman, 1998) explains the outbreak of a currency run as a symptom of accumulated problems in the banking and financial sector. In the theoretical setting, government guarantees aimed at attracting foreign investment lead to a bubble on the asset market that eventually bursts and creates the crisis. Obviously, given the different theoretical nature of the ultimate cause of the currency crises in the different generations of models, the potential empirical determinants to be included in econometric studies vary strongly depending on the theory used to select covariates.

The objective of the present paper is to revisit binary-variable models for currency crises based on macroeconomic fundamental data by explicitly taking into account model uncertainty. In particular, we want to work out to what extent model uncertainty puts the robustness of the explanatory variables of the logit models championed in the literature (e.g. Bussière and Fratzscher 2006 or Bussière 2007) under strain. On the one hand, our results indicate that the usual macroeconomic variables used in empirical studies of currency crisis are very fragile determinants of the occurrence of such episodes. On the other hand, if we redefine the crisis indicator as to give a signal for observations up to one year prior to the crisis, several variables appear as robust determinants of these crisis periods. Financial market indicators and the deviations of the real exchange rate from a linear trend present very high posterior model inclusion probabilities and thus can be considered robust determinants of crisis periods.

The remainder of this chapter is structured as follows: Section 2.2 sketches the Bayesian model averaging procedure. In section 2.3 the data are described and variables defined. Section 2.4 presents the results on the extent to which model uncertainty matters, while section 2.5 concludes.

2.2 Dealing with model uncertainty: Bayesian model averaging

The binary variable we are interested in modelling takes value one if a currency crisis occurs in period t ($y_i = 1$) and zero if no currency crisis is observed ($y_i = 0$). A stereotypical regression aimed at assessing the effect of a set of variables $\{\mathbf{x}_j\}_{j=1}^{K}$ on the probability of a currency crisis occurring is given by

$$\mathrm{P}(y_i = 1 | \{\mathbf{x}_j\}_{j=1}^{K}) = F(\mathbf{X}_K \beta), \tag{2.1}$$

where $F(z)$ will typically be a logistic function ($F(z) = (1 + e^z)^{-1}$) or the distribution function of a normal random variable ($F(z) = \Phi(z)$), $\mathbf{X}_K = (x_1 \ \ldots \ x_K)$, which is a subset of $\mathbf{X}_{\bar{K}} = (x_1 \ \ldots \ x_{\bar{K}})$, containing all possible regressors ($\bar{K} > K$ of them), and $\beta = (\beta_1 \ \ldots \ \beta_K)'$. In principle, many candidate variables can be proposed as potential covariates in (2.1).

So far, the literature tends to concentrate on an arguably tiny subset of this model space. Model averaging techniques propose averaging over all these alternative models using Bayes factors so as to evaluate the relative importance of different variables as determinants of the occurrence of a currency crisis. In the situation where there are M competing models, $\{M_1, \ldots, M_M\}$, which are defined by the choice of independent variables, so that $M = 2^{\bar{K}}$, Bayesian inference about the parameter of interest, β_i is based on its posterior distribution (that is, the distribution given the data, $\mathbf{Y} = \{y \ \mathbf{X}_K\}$),

$$\mathrm{P}(\beta_i | \mathbf{Y}) = \sum_{m=1}^{M} \mathrm{P}(\beta_i | \mathbf{Y}, M_m) \mathrm{P}(M_m | \mathbf{Y}), \tag{2.2}$$

where the posterior probabilities $\mathrm{P}(M_k | \mathbf{Y})$ are given by

$$\mathrm{P}(M_k | \mathbf{Y}) = \frac{\mathrm{P}(\mathbf{Y} | M_k) \mathrm{P}(M_k)}{\sum_{m=1}^{M} \mathrm{P}(\mathbf{Y} | M_m) \mathrm{P}(M_m)}. \tag{2.3}$$

The posterior model probabilities can thus be obtained as the normalized product of the integrated likelihood for each model ($\mathrm{P}(\mathbf{Y} | M_k)$) and the prior probability of the model ($\mathrm{P}(M_k)$). Notice that for the simple case $m = 2$ the posterior odds for a model against the other can be readily written as the product of the Bayes factor and the prior odds. Further assuming equal priors across models, the posterior odds are equal to the Bayes factor ($\mathrm{P}(\mathbf{Y} | M_2) / \mathrm{P}(\mathbf{Y} | M_1)$). The Bayes factor, in turn, can

be accurately approximated (see Leamer, 1978, and Schwarz, 1978) as

$$\frac{\mathrm{P}(\mathbf{Y}|M_2)}{\mathrm{P}(\mathbf{Y}|M_1)} = N^{(k_1-k_2)/2}\left(\frac{Lik_2}{Lik_1}\right), \tag{2.4}$$

where N is the number of observations, k_j and Lik_j are respectively the number of parameters and the likelihood of model j. This simple approximation allows us to compute (2.3) and the corresponding statistics based on (2.3).

This implies that for a given prior on the model space, the posterior distribution of β can be obtained as a weighted average of the model-specific estimates weighted by the posterior probability of the respective models. If the cardinality of the model space is computationally tractable, (2.3) can be obtained directly and (2.2) can be computed. In particular, the expected value of β and its variance, $\mathrm{E}(\beta|\mathbf{Y})$ and $\mathrm{var}(\beta|\mathbf{Y})$ respectively, can be computed as follows

$$\mathrm{E}(\beta_i|\mathbf{Y}) = \sum_{m=1}^{M} \mathrm{E}(\beta_i|\mathbf{Y}, M_m)\mathrm{P}(M_m|\mathbf{Y}), \tag{2.5}$$

$$\mathrm{var}(\beta_i|\mathbf{Y}) = \sum_{m=1}^{M} [\mathrm{var}(\beta_i|\mathbf{Y}, M_m) + \mathrm{E}(\beta_i|\mathbf{Y}, M_m)^2]\mathrm{P}(M_m|\mathbf{Y}) - \mathrm{E}(\beta_i|\mathbf{Y})^2. \tag{2.6}$$

The posterior mean and variance can be used to make inference on the quantitative effect of changes in the covariates on the probability of a currency crisis explicitly taking into account model uncertainty. Several methods have been proposed for approximating the expression in (2.3) when the cardinality of the model space makes the problem intractable. The *leaps and bounds* algorithm, the use of Markov Chain Monte Carlo Model Composite (MC^3) methods or the use of Occam's window are possible methods of setting bounds to the number of models to be evaluated when computing (2.3) (see Raftery, 1995, for an excellent description of these methods).

In our empirical application we will use a simple MC^3 algorithm to evaluate the posterior distribution based on the work of Madigan and York (1995), also used recently by Fernández *et alia* (2001) in the framework of cross-country growth regressions.[2] This Markov Chain Monte Carlo method implements the Random Walk Chain Metropolis-Hastings algorithm in the model space as follows. In a given replication s of the algorithm, a candidate model M^{s+1} is proposed, which is randomly drawn from the group of models composed by the model which is active in that

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[2]Koop (2003) also describes the method thoroughly.

replication (M^s), the same model with an extra variable added to the specification and the same model with a variable removed. The proposed model is accepted with a probability given by

$$\alpha(M^s, M^{s+1}) = \min \left[\frac{\mathrm{P}(\mathbf{Y}|M^{s+1})\mathrm{P}(M^{s+1})}{\mathrm{P}(\mathbf{Y}|M^s)\mathrm{P}(M^s)}, 1 \right],$$

which is just the Bayes factor comparing M^s and M^{s+1} if equal prior probability is assumed across models, so that $\mathrm{P}(M^s)$ and $\mathrm{P}(M^{s+1})$ cancel out in the expression above. This algorithm is repeated a large number of times, and the sums defined above are computed for the group of models replicated, which will tend to cover model subspaces with the highest posterior probability.

In the same fashion, posterior inclusion probabilities for the different variables can be obtained by summing the posterior probability of models containing each variable. This measure captures, thus, the relative importance of the different covariates as determinants of the occurrence of a currency crisis and can be interpreted as the probability that a given variable belongs to the true specification.

2.3   Data and variable descriptions

2.3.1   Data description

The early warning system for currency crises dealt with in this paper is derived from a binary-variable model based on macroeconomic fundamental data, in the spirit of the classical contributions by, for instance, Frankel and Rose (1996). Since currency crises are events which occur seldom, in this type of models it is necessary to pool country/time data in order to increase the number of observations and obtain sufficient degrees of freedom. Naturally, this procedure implicitly imposes the assumption of parameter homogeneity across countries and in the time dimension. The resulting first requirement on our sample thus was that the crises episodes considered be sufficiently homogeneous, that is, characterized by a similar development of fundamentals. In addition, however, it was also desirable in this context to employ the same data source as a recent benchmark study using a 'standard' binary-variable approach (that is, without explicitly dealing with model uncertainty) in order to be able to figure out the value added by our model averaging procedure.

For these reasons, we decided to use as a yardstick for comparison the dataset of one of the most recent papers on this issue by Bussière (2007), who exercised great

care in constructing a sample sufficiently homogenous so that common fundamental development driving the crises may be expected. Against this backdrop the overall sample consists of a pool of observations on 27 countries recorded from January 1994 to March 2003 and contains approximately 1400 observations[3]. Observations prior to 1994 are taken out of the sample to avoid biases emanating from hyperinflationary experiences in Latin American countries and the early years of transition towards a market economy in Eastern European economies.[4]

The dependent binary variable is defined to equal one if a crisis occurs and zero otherwise. Although in the common understanding a currency crisis might be associated predominantly with a dramatic devaluation of the exchange rate, the literature on early-warning mechanisms usually tends to employ a broader definition of currency distress by using the concept of exchange market pressure. Although the latter is not uniformly defined in the literature it is usually a weighted average of some combination of the change of the real or nominal exchange rate, the country's foreign reserves and the real interest rate. The dependent variable is thus computed in two steps. First, the exchange market pressure index ($EMPI_{i,t}$) for country i at time t is defined as

$$EMPI_{i,t} = \omega_{RER}\left(\frac{\Delta RER_{i,t}}{RER_{i,t-1}}\right) + \omega_r\left(\Delta r_{i,t}\right) - \omega_{res}\left(\frac{\Delta res_{i,t}}{res_{i,t-1}}\right),$$

where RER stands for the real effective exchange rate, r is the short-term real interest rate and res the level of international reserves. In the next bout this continuous variable is transformed into a binary index which equals one whenever $EMPI_{i,t}$ exceeds the threshold of the country-specific mean ($\overline{EMPI}_i$) plus twice its standard deviation (σ_{EMPI_i}),

$$CI_{i,t} = \begin{cases} 1 \text{ if } EMPI_{i,t} > \overline{EMPI}_i + 2\sigma_{EMPI_i}, \\ 0 \text{ otherwise.} \end{cases}$$

The choice of the explanatory right-hand side variables in (2.1) is motivated by the theoretical literature on currency crises on the one hand and by the results of the

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[3]The countries included in the sample are Argentina, Brazil, Chile, Colombia, Ecuador, Mexico, Peru, Venezuela, China, Hong Kong, India, Indonesia, Korea, Malaysia, Philippines, Singapore, Thailand, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Russia, Slovak Republic, Slovenia, Turkey

[4]Bussière and Fratscher (2006) tested for slope homogeneity in a very similar dataset by comparing out-of-sample forecasts based on the parameter homogeneity assumption. From the good forecasting performance they conclude that the same parameter vector is suitable for different countries and episodes. In contrast, the sample used by Peltonen (2006) which contains also data on crises from the 1980s suggests significant difference between Latin America and Asia.

existing empirical early warning models on the other. Table 2.1 lists the complete final set of variables, different combinations and transformations of which are used in the estimations below.

Variable name	Definition	Details
Exchange rate (trend deviation)	$REERDEV_{i,t} = \frac{REER_{i,t} - Trend_{i,t}}{Trend_{i,t}} * 100$	Trend defined as a simple linear trend
Lending boom	$LB_{i,t} = \left(\frac{CPS_{i,t}}{GDP_{i,t}} - \frac{\overline{CPS}_{i,t-24}}{\overline{GDP}_{i,t-24}}\right) * 100$	$\frac{\overline{CPS}_{i,t-24}}{\overline{GDP}_{i,t-24}} = \frac{1}{12}\sum_{k=0}^{11} \frac{CPS_{i,t-24-k}}{GDP_{i,t-24-k}} * 100$; $CPS_{i,t}$ = credit to the private sector
Short-term debt/reserves	$STDR_{i,t} = \frac{STD_{i,t}}{RES_{i,t}} * 100$	$STD_{i,t}$ = short term debt; $RES_{i,t}$ = international reserves
Total debt/reserves	Analogously to the previous one	Locational or consolidated definition
Current account balance	$\frac{CA_{i,t}}{GDP_{i,t}}$	
Govnernment balance	$\frac{GB_{i,t}}{GDP_{i,t}}$	
Financial contagion	$CONT_{i,t} = \sum_{j=1}^{N-1} EMPI_{i,t} * Correl_{i,j}$	$Correl_{i,j}$ = correlation of equity market returns between country i and country j
Datastream index, total market	12-months percentage change	Broad index
Datastream index, banks	12-months percentage change	Sub-index
Datastream index, financial institutions	12-months percentage change	Sub-index
GDP growth rate	Yearly growth rate of GDP	

Table 2.1: Basic variables: Definitions

The exchange rate variable is supposed to capture any excessive real overvaluation of the currency, which would be expected to increase the risk of devaluation. It is defined as the deviation of the real exchange rate from a linear trend. Since data on non-performing loans are barely available for under-reporting reasons, the lending boom indicator is meant to serve as a proxy and is defined as the deviation of the credit to the private sector ($CPS_{i,t}$) from a one year average with a two year lag. The short-term-debt-to-reserves ratio (and analogously the total debt indicator) reflect the so called Greenspan-Guidotti rule which states that reserves should cover entirely the amount of external debt that can be sold short-term by investors in case of an attack. A rise of this indicator can thus stem from either a rise in debt or a fall of reserves and should render a crisis more likely. The total debt indicator is defined analogously for two different definitions: the locational (lc) and the consolidated concept (cc).[5] The set of explanatory variables further contains the current account and government surpluses, both normalized with the respective country's GDP. The sign of these two indicators is expected to be negative as the higher the surplus (the lower the deficit) the lower should be the probability of an attack. Since Bussierè and Fratscher (2006) show that contagion across countries is only significant via the financial and not via the trade channel, only the former was taken into account in Bussière (2007). Financial interlinkages of a country i with all other countries in the sample are modelled as the average of the other countries' $EMPI_{j,t}$ ($j = 1$ to $N - 1$, $j \neq i$) weighted by the correlation of equity market returns in country i and country j. Intuitively, the parameter attached to this variable should show up positive in the estimation results. The three subsequent Datastream indices, a broad market index and two sub-indices on banks and financial institutions, account for the predictive power of financial markets. They are defined as a 12-months percentage change of each stock index and are expected to enter with a negative coefficient. Finally, the year-on-year GDP growth is included as higher economic growth should reduce the government's temptation to devalue on its currency, e.g. in order to gain competitiveness. [6]

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[5]The locational banking statistics gather data on international financial claims and liabilities of bank offices resident in the reporting countries on a gross (unconsolidated) basis, including those vis-à-vis own affiliates. In contrast, the consolidated concept covers claims reported by domestic bank head offices, including the exposures of their foreign affiliates, and are collected on a worldwide consolidated basis with inter-office positions being netted out. For details see Bank for International Settlements (2003).

[6]Further details on the construction of the variables and the intuition behind their choice can be found in Bussière (2007)

2.4 Empirical results: How much does model uncertainty matter?

2.4.1 Results for the "crisis occurrence" indicator

Following Bussière (2007), we present results based on three types of specification. Firstly, we deal with a purely static model, where lags of the dependent variable do not appear as extra regressors in the model, although all explanatory variables are evaluated with one month lag with respect to the crisis variable. We then address dynamic models, which on top of the exogenous set of variables employed in the static model also include up to six lags of the crisis index as explanatory variables. Finally, the most general specification includes up to 24 lags of six selected variables $(REERDEV, LB, STDR, \frac{CA}{GDP}, CONT, GROWTH)$.[7]

In Table 2.2 we report the results of the BMA exercise for the static case, where all specifications in the model space have been evaluated in order to compute posterior inclusion probabilities and posterior expected values of the parameters.[8] We also deal explicitly with the issue of potential multicollinearity among the regressors. The first two columns of the table show the posterior expected values of the parameters corresponding to each variable (first column) and the posterior inclusion probabilities (second column) for the BMA exercise using all variables in Table 2.1. Under the header *Static uncorrelated* the results are presented for the BMA exercise after taking out variables whose correlation with some other explanatory variable was equal to or greater than 0.5 (both total debt indicators and one of the Datastream indices are the variables which do not enter this exercise).

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[7]Bussière (2007) also estimates models with fixed effects and reports that the hypothesis that all country fixed effects are equal to zero can be rejected, but admits that the p-value of the test is close to 10%. Conditional logit models are also estimated by Bussière (2007) for both the static and the dynamic model, with results which are very close to those from the model where no fixed effects were used.

[8]In order to keep the table readable, we do not report the posterior variances of the parameters, which are available from the authors upon request.

Variable	Static		Static uncorrelated		Bussière (2007) static	
	E(β_i|**Y**)	Inc. Prob.	E(β_i|**Y**)	Inc. Prob.	Simple static	Fixed effects
Exchange rate, dev. from trend	0.007658	0.001285	0.021211	0.001445	[0.025; 0.042]	[0.032; 0.045]
Lending boom	0.004109	0.002581	0.008854	0.002312	[0.006]	[0.005; 006]
Short-term debt/reserves	0.000456	0.001335	0.000976	0.001278	[0.003; 0.004]	[0.007; 0.009]
Total debt/reserves (lc)	0.000252	0.001266			[0.002]	[0.004]
Total debt/reserves (cc)	0.000333	0.001701			[0.003]	[0.006]
Current account balance	−0.012997	0.001434	−0.032725	0.001473	n.s.	[−0.141; −0.077]
Govnernment balance	0.020414	0.000947	0.048560	0.000945	[0.082]	n.s.
Financial contagion	0.023883	0.011678	0.051947	0.009797	[0.043]	[−0.015; 0.043]
Datastream index, total market	−0.002986	0.008047	−0.012907	0.033148	n.s.	[−0.008]
Datastream index, banks	−0.003206	0.023155			n.s.	[−0.014]
Datastream index, financial institutions	−0.003414	0.029542	−0.011913	0.095018	n.s.	[−0.015]
Growth rate	−0.007179	0.000846	−0.014045	0.000838	[−0.052]	n.s.

Table 2.2: BMA results: Static model

These posterior expected values of the parameters can be compared with the results reported in Bussière (2007), which are shown in the fifth column for the simple static model and in column six for the static model with fixed effects. Since Bussière alternates the set of included variables to avoid multicollinearity we report here the range in which his (significant) estimates fall ($n.s.$ stands for non-significant, if no estimate on at least the 10%-level was available). Two facts call attention when considering the results in Table 2.2. First of all, the posterior expected parameter values have mostly the expected sign. The probability of a crisis thus tends to increase with the lending boom, debts relative to reserves, the contagion indicator and the deviation of the exchange rate from its trend. In contrast, robust growth and rising market indices and current account surpluses reduce the risk of a currency attack. The only somewhat counter-intuitive result, consistently confirmed in all estimations, is the positive sign of the government balance variable.[9]

However, the lack of robustness of the relationships under study shows up when considering the posterior inclusion probabilities reported in Table 2.2. Since we assign equal prior probability to all models when computing the posterior model averaged objects, our prior on the inclusion probability of each variable is 0.5.[10] After observing the data, the probabilities of including each variable decreases strongly with respect to the prior, with none of the posterior probabilities being higher than 10%. To put it differently, the model with the greatest posterior probability (in fact one that is very close to 1) implies a constant crisis probability which is not country or time-specific (that is, the model including only a constant).

Table 2.3 is constructed in the same manner as Table 2.2 for the case of the dynamic model, including lags of the dependent variable. With the exception of the government balance variable, all variables show up again with the expected signs which coincide with those obtained by the benchmark study, when they are significant. However, except for the market indices, this time our coefficients appear to be substantially smaller in magnitude than Bussière's (2007). The posterior inclusion probabilities are once more well below the 0.5 threshold. In other words, the inclusion of six de-facto new variables does not lead to any improvement of the ex-

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[9]However, it should be borne in mind that the sample for all estimations starts in 1994, and it is a well known fact that first generation models generally fail to explain crises in the 1990s. Second and third generation models might actually get some support by this somewhat surprising result (see for example Krugman, 1996 and Bussière, 2007).

[10]There are $2^{\bar{K}-1}$ models including a given variable and $2^{\bar{K}}$ total models, so the prior inclusion probability of a given variable is $2^{\bar{K}-1}/2^{\bar{K}}=0.5$.

planatory power of macroeconomic fundamentals. Bussière finds that the dependent variable is significant only at lag 5 and 6 in both models, with and without fixed effects. The interpretation of this result is that crises sometimes hit in two waves such that the first attack is often followed by a second bout within a short time distance. In this context, it is also interesting to note that all the coefficients of the lagged crisis index in our and Bussière's regressions enter with a positive sign. Hence, past crises tend to increase the likelihood of repeated attacks, a result which is not quite obvious ex-ante. On the one hand, a country that has experienced a crisis may be deemed more vulnerable by investors which would speak for a positive sign. On the other hand, however, two arguments can be proposed why crises in the past might reduce the probability of an attack in the future. In the short run, after a currency run there is not much speculative capital left to be withdrawn. Moreover, in the longer run, one can argue that the country previously hit has improved its vigilance and supervision mechanisms which should render a repeated crisis less likely.

Variable	Dynamic		Dynamic uncorrelated		Bussière (2007) dynamic	
	$E(\beta_i|\mathbf{Y})$	Inc. Prob.	$E(\beta_i|\mathbf{Y})$	Inc. Prob.	Simple static	Fixed effects
Crisis index, lag 1	0.582846	0.003311	0.584356	0.003343	n.s.	n.s.
Crisis index, lag 2	0.034475	0.000722	0.037394	0.000722	n.s.	n.s.
Crisis index, lag 3	0.032963	0.000723	0.036480	0.000723	n.s.	n.s.
Crisis index, lag 4	0.357034	0.001151	0.359204	0.001157	n.s.	n.s.
Crisis index, lag 5	0.776955	0.015559	0.778136	0.015779	[1.636; 1.909]	[1.298; 1.615]
Crisis index, lag 6	0.760693	0.014288	0.761846	0.014475	[1.344; 2.036]	[1.229; 2.057]
Exchange rate, dev. from trend	0.007840	0.001315	0.007729	0.001296	[0.037; 0.056]	[0.042; 0.599]
Lending boom	0.004101	0.002564	0.004100	0.002567	n.s.	n.s.
Short-term debt/reserves	0.000455	0.001331	0.000457	0.001329	[0.003; 0.004]	[0.009; 0.011]
Total debt/reserves (lc)	0.000252	0.001263			[0.0017]	[0.005]
Total debt/reserves (cc)	0.000332	0.001694			[0.003]	[0.007]
Current account balance	−0.013013	0.001436	−0.013017	0.001437	n.s.	[−0.082]
Govnernment balance	0.020321	0.000945	0.020402	0.000947	n.s.	n.s.
Financial contagion	0.023864	0.011573	0.023900	0.011740	n.s.	[−0.056; 0.052]
Datastream index, total market	−0.002978	0.007926	−0.002984	0.008090	n.s.	n.s.
Datastream index, banks	−0.003200	0.022722			[−0.011]	[−0.013]
Datastream index, financial institutions	−0.003408	0.028988	−0.003407	0.029626	[−0.011]	[−0.016]
Growth rate	−0.007056	0.000843	−0.007230	0.000846	n.s.	n.s.

Table 2.3: BMA results: Dynamic model

In order to account for a general dynamic structure in the model, Bussière (2007) regresses in a standard logit model (without fixed effects) the dependent variable on six chosen explanatory variables ($REERDEV, LB, STDR, \frac{CA}{GDP}, CONT, GROWTH$) which are all lagged by 1 to 24 months. This series of regressions thus provides him with 24 different models and 144 different coefficients from which the author draws the conclusion that *"some variables have a very short-term impact, such as the short-term debt to reserve ratio, some have both a very short-term and a longer term impact (such as the contagion variable), some have a short- to medium-term impact (such as the lending boom), some always seem to have an impact (such as the exchange rate), while for growth and the current account, no impact can be detected"* (Bussière, 2007, page 26). We conducted a different exercise at this point and constructed the BMA procedure using as explanatory variables six lags of the crisis variable and 24 lags of all 12 variables listed in Table 2.1, all at the same time. Hence, this setting contains 294 potential explanatory variables which imply 2^{294} (more than 3×10^{88}) different models over which we have to average. Given the fact that, with the current technology, this does not appear possible in a lifetime[11], we used the MC^3 approach described above to evaluate the posterior objects.

In table 2.4 we confine ourselves to reporting only the results for the lags of each variable with the highest posterior inclusion probability.[12] Focusing on the coefficients in the second column one can note that some of the signs now have changed into an unexpected direction. The government surplus, which used to carry a counterintuitive positive coefficient now has got the "right", negative sign, while more robust growth, higher current account surpluses and lower lending suddenly and counter-intuitively increase the probability of a crisis - at least for the lags with the highest inclusion probability. As if this was not puzzling enough, the sign of the coefficients is not uniform for all lags but rather alternates from positive to negative for all variables. Interestingly enough, the fluctuation pattern looks to a great extent similar to the one derived by Bussière (2007). In his estimations growth, for instance, only has the expected negative sign for lags 1 to 8 and 16 to 19. Similarly, current account surpluses lagged by more than 11 months increase the probability of a crises. The latter is also more likely the lower was the lending boom 18 months

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[11]If it took only 0.001 second to estimate one model the whole calculation would last 1.009×10^{78} years. Although the reasoning put forward above could also imply that interactions between long and short-term variables play an important role in unwinding currency crises, due to the extra computational burden imposed by the use of cross-products, we do not embark in this type of exercise in the present study.

[12]The complete set of results is available from the authors.
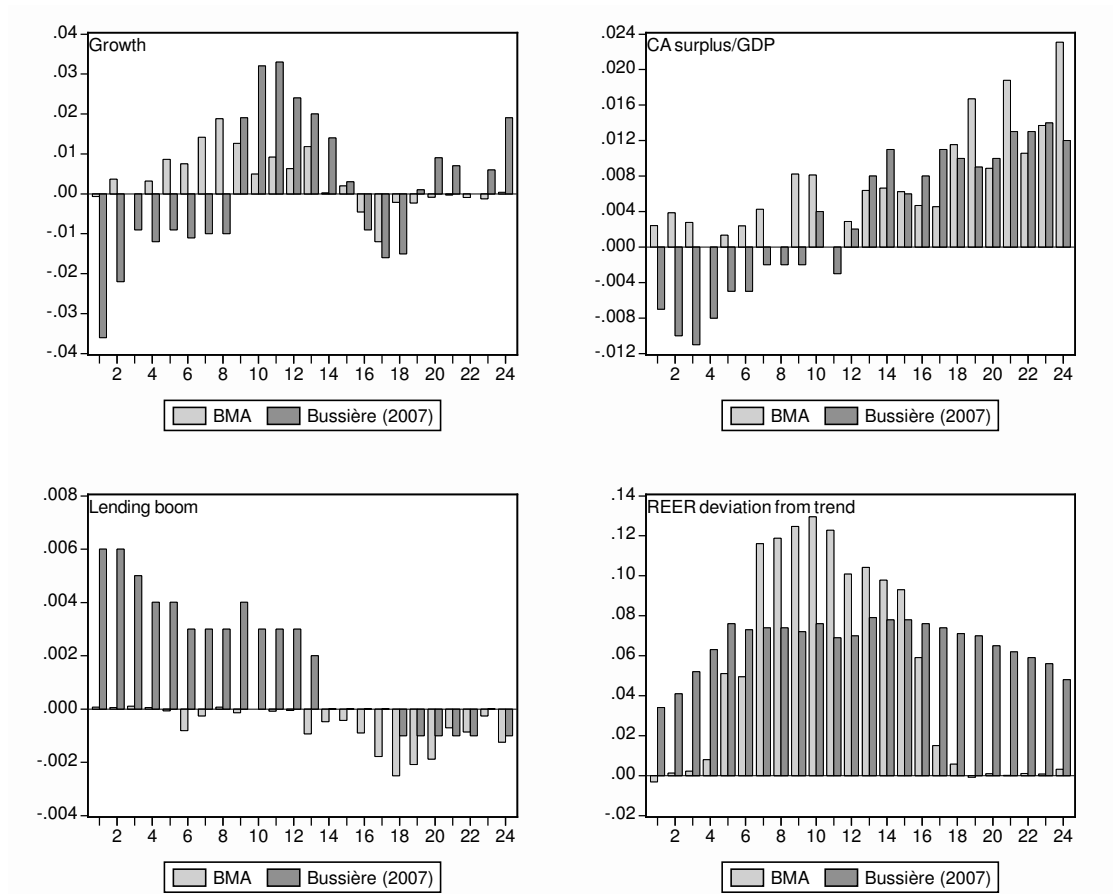
Figure 2.1: Estimated parameters at different lag lengths for selected variables

or more ago. It has to be added, however, that growth, current account and the lending boom from lag 13 on are not significant (see Figure 2.1, which presents some of the parameters estimated by BMA against Bussière's results).

Among the remaining variables which carry the same sign as in the previous calculations (for the lag with the highest posterior inclusion probability at least) it strikes that the effect of the lagged crisis binary variable is again the most robust at lag 5. In addition, the effect of the exchange rate deviation from trend is now almost twenty times bigger than in Tables 2.2 and 2.3. This is because the coefficient of the exchange rate variable shows a strong bell-shaped form, rising strongly between lags 4 and 10 and decreasing sharply after that. This contradicts somewhat Bussière's

Variable	$E(\beta_i|\mathbf{Y})$	Max. inc. prob.	at lag
Crisis index	0.47073	2.08×10^{-5}	5
Current account balance	0.023066	3.18×10^{-5}	24
Govnernment balance	-0.017251	1.30×10^{-5}	18
Growth rate	0.008615	1.69×10^{-5}	5
Lending boom	-0.000927	1.02×10^{-5}	13
Financial contagion	0.035614	9.55×10^{-5}	5
Datastream index, banks	-0.005556	4.22×10^{-5}	24
Datastream index, financial institutions	-0.005589	6.07×10^{-5}	24
Datastream index, total market	-0.010482	0.000154	24
Exchange rate, dev. from trend	0.129465	0.894149	10
Short-term debt/reserves	0.00066	1.98×10^{-5}	7
Total debt/reserves (cc)	0.000435	3.53×10^{-5}	23
Total debt/reserves (lc)	0.000183	1.93×10^{-5}	22

Table 2.4: BMA results: Dynamic model with lagged explanatory variables

results according to which the exchange rate effect seems much more homogenous and significant for all lags. Lastly, it may also be pointed out that all market signals seem to be most symptomatic of tension on the exchange rate market 2 years in advance, which is not quite easy to interpret either.

As can be seen in the third column of Table 2.4 which displays the lag with the maximum posterior inclusion probability for each variable all values but one are far beyond good and evil. Only the deviation of the exchange rate from trend at lag 10 shows up with a posterior inclusion probability above the prior of 0.5. Although the importance of the variable is clear, by no stretch of imagination we can think of any plausible explanation for the fact that only the tenth lag appears robust, and even less so if considering the fact that the second highest inclusion probability for this variable (at lag 9) is more than ten times smaller. We thus argue that it is just a matter of coincidence and that also in this exercise fundamentals have proven to have no systematic and robust explanatory power for currency crises.

2.4.2 Results for the "crisis period" indicator

The results presented above are based on a crisis index which indicates a crisis in a particular month if the continuous exchange market pressure index exceeds a certain threshold in that month. In other words, a model based on this definition

of a crisis attempts to predict the exact timing of a crisis in a given country. As we have shown, if we employ this crisis definition and address model uncertainty in a Bayesian manner we, unlike Bussière (2007), find virtually no robustness of the potential explanatory variables. The model-based results by Bussière (2007), however, do not perform too well in terms of prediction. It is argued in Bussière (2007) that, by trying to predict the exact month of a crisis, the model attempts to achieve something that may simply be infeasible. In order to address this caveat the time window of the crisis definition is extended to a whole year. Hence, a crisis signal is now issued not only if a strong depreciation occurs within a month but if the $EMPI$ exceeds the threshold in any of the successive 12 months. The corresponding (transformed) crisis indicator (TCI) is thus

$$TCI_{i,t} = \begin{cases} 1 \text{ if } \exists k \in 1, \ldots, 12 \mid CI_{i,t+k} = 1, \\ 0 \text{ otherwise.} \end{cases}$$

If the reason for the middling explanatory power in our results is the narrow definition of a crisis and the difficulty of predicting the exact timing of such episodes, then this broader definition should improve the inclusion probabilities of our explanatory variables. It should be noticed that in this case we are giving more relevance to the explanatory power for differences between countries, as opposed to within countries.

Analogously to our exercise for the original index, we estimate models within static and a dynamic specification classes using now this transformed crisis index.[13] The results are presented in Tables 2.5 and 2.6. For comparison we again report the intervals of significant parameter values obtained by Bussière (2007).[14]

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[13]Note that the extension of the time window implies a certain information loss. In the dynamic panel in this new setting the dependent variable thus has to be lagged by 12 months.

[14]Note that the comparability is limited in the dynamic setting since we, unlike Bussière, include the stock indices in this specification.

Variable	Static		Static uncorrelated		Bussière (2007) static	
	E($\beta_i|\mathbf{Y}$)	Inc. Prob.	E($\beta_i|\mathbf{Y}$)	Inc. Prob.	Simple static	Fixed effects
Exchange rate, dev. from trend	0.157588	1	0.15754	1	[0.092; 0.104]	[0.100; 0.131]
Lending boom	0.007909	0.044759	0.007805	0.041711	[0.007]	[0.006]
Short term debt/reserves	0.001784	0.010563	0.001018	0.008125	[0.004]	[0.009; 0.012]
Total debt/reserves (lc)	-0.001541	0.003943			n.s.	[0.008]
Total debt/reserves (cc)	0.000848	0.016333			[0.005]	[0.013]
Current account surplus	-0.025566	0.005702	-0.025418	0.005445	n.s.	[-0.083; -0.061]
Government surplus	-0.047456	0.003069	-0.04733	0.002895	n.s.	[0.243]
Financial contagion	0.071689	0.999994	0.071657	1	[0.063; 0.070]	[0.071; 0.084]
Growth rate	-0.025984	0.002907	-0.026202	0.002818	[-0.063; -0.04]	[-0.059; -0.026]

Table 2.5: BMA results for *TCI* as dependent variable: Static model

Variable	Dynamic		Dynamic uncorrelated		Bussière (2007)	
	$E(\beta_i|Y)$	Inc. Prob.	$E(\beta_i|Y)$	Inc. Prob.	Simple static	Fixed effects
Lagged dependent variable	-0.40697	0.003468	-0.400762	0.001675	n.s.	n.s.
Exchange rate, dev. from trend	0.160483	1	0.157799	1	[0.099; 0.109]	[0.103; 0.134]
Lending boom	0.006963	0.016235	0.007558	0.028585	n.s.	n.s.
Short term debt/reserves	0.00125	0.006489	0.001003	0.007343	[0.004]	[0.009; 0.012]
Total debt/reserves (lc)	-0.000956	0.002355			n.s.	[0.008]
Total debt/reserves (cc)	0.000832	0.012289			n.s.	[0.013]
Current account surplus	-0.02254	0.003624	-0.024533	0.004854	n.s.	[-0.070; -0.081]
Government surplus	-0.064848	0.009033	-0.06368	0.007149	n.s.	n.s.
Financial contagion	0.062504	0.999167	0.06708	1	[0.061; 0.067]	[0.075; 0.086]
Data stream index, total market	-0.00682	0.031511	-0.007013	0.499063		
Data stream index, banks	-0.005494	0.201715				
Data stream index, institutions	-0.006329	0.737342				
Growth rate	-0.015798	0.001164	-0.026202	0.001412	[-0.052]	[-0.029; -0.039]

Table 2.6: BMA results for *TCI* as dependent variable: Dynamic model

Most of the inclusion probabilities remain well below the prior threshold of 0.5. However, it strikes that in both static regressions the inclusion probabilities have improved dramatically for the real effective exchange rate deviation from a linear trend and for the financial contagion variable. Both variables now have a posterior inclusion probability close to one. Moreover, they both have the expected sign and in terms of magnitude come very close to the parameters obtained by Bussière. The same holds true also in the dynamic setting, where, in addition, also the included Datastream stock index (institutions) shows substantial explanatory abilities.

These results suggest that, while the exact timing of a crisis may indeed be unpredictable, differences in macroeconomic and financial variables still contain information about differential degrees of currency crisis exposure. Moreover, for the probability of a "crisis period" merely two groups of variables seem to matter: the deviation of the real exchange rate from trend on the one hand, and financial market indicators on the other. These results enforce the hypothesis that each currency crisis is eventually triggered by the behavioral change of financial market participants, who seem to care to some extent about a handful of macroeconomic variables and to a great extent about the (herding) behaviour of their colleagues.

2.5 Summary and conclusions

The dominant majority of early warning mechanisms for currency crises employs some version of fundamental-based binary choice models. To our knowledge, none of the papers on the subject tackles the issue of model uncertainty in currency crisis model explicitly. In the present paper we have explicitly taken into account model uncertainty in the framework of a binary choice model. By means of Bayesian model averaging we estimate the coefficients for each variable as weighted averages over the alternative models from the model space, where the weights correspond to the posterior probability of each model. In order to figure out the value added by this approach as opposed to "standard" logit regressions we have used the same data set as one of the most recent studies on the subject by Bussière (2007).

If the discrete dependent variable is constructed so as to predict the exact month in which a crisis may happen our conclusions are twofold. On the one hand, we have found that coefficients mostly have the expected signs coinciding with the benchmark study. On the other hand, however, our principal quality gauge, the posterior inclusion probability (the sum of posterior probabilities of all models containing a particular variable), unveils the lacking robustness of the relationships between re-

gressors and the dependent variable. These results imply that at least in this setting the best model to explain a currency crisis is a mere time and country-unspecific constant. Our results, therefore, indicate that none of the usual macroeconomic fundamental variables is a robust determinant of a currency crisis for the definition and sample used. The results improve considerably if we consider defining "crisis periods" instead of crisis occurrences. Defining crisis periods as observations up to one year prior to the crisis, we find that real exchange rate developments and financial variables are able to robustly explain differences in the probability of a country experiencing such episodes.

Since our sample starts in 1994 it could well be that episodes of currency distress included in the sample are crises rather of the second and third generation type. In such a case it would not be surprising that fundamental data show only limited explanatory power. To turn the argument around, the fundamentals should play a much more significant role in a sample covering the first generation type of crises. Exactly along these paths we are planning to conduct our future research.

A finer way of testing the different theoretical frameworks proposed by the three generations of currency crises models would imply grouping variables by theory and computing the joint inclusion probability of these groups of variables. The construction of groups of variables by theory could be handled in the BMA framework using the proposal by Brock, Durlauf and West (2003) of using a hierarchical prior in order to sort variables into theories or thematic indicators (see also the recent contribution by Doppelhofer and Weeks, 2007, for the concept of jointness of determinants in the BMA framework). Although we did not follow this approach in the paper, we propose it as a potentially fruitful path of further research.

An interesting issue that has not been directly tackled in the paper and that would deserve further scrutiny is the possibility of nonlinear effects in form of interactions among the potential determinants of crises. Developments in some relevant variables may just be responsible for preparing the ground for imbalances that end up a currency crisis when triggered by an unsound development in an additional variable. The use of interaction terms in a BMA setting could assess the importance of this type of effects.

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Chapter 3

(How) Will the Euro Affect Inflation in the Czech Republic?

3.1 Introduction

A lot has been said about the costs and benefits of monetary unions in general and about the common European currency in particular. One of the disadvantages of euro adoption typically most feared by the general public is a sizable rise of the price level. According to the Eurobarometer survey from November 2007, 74 % of citizens in the EU - New Member States (NMS)[1] agree with the statement that the euro will increase prices (see Eurobarometer 2007). An important role in forming inflation expectations related to the introduction of the euro is played by media as well as influential and trustworthy institutions and politicians. Inflation expectations, on their part, might have a substantial impact on future actual inflation rates. Therefore, in the discussion on the possible inflationary impact of the common currency it is essential to provide the public with balanced, understandable and transparent arguments. Using the case study of the Czech Republic the objective of the present paper is thus to put the discussion on a firm footing by analysing in a qualitative and quantitative manner the channels through which the euro adoption might have an effect on inflation embedded in a framework of other cyclical, structural and external inflation factors. The fact that the Czech Republic does not have an official target date after the originally envisaged horizon of 2010 has been postponed does not preclude an intense and partially very prominent discussion on pros and cons of the euro that primarily centres around the expected inflationary impact. In contrast,

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[1] By NMS we mean the following countries: Poland, the Czech Republic, Hungary, Slovakia, Slovenia, Estonia, Latvia, Lithuania, Malta, Cyprus, Bulgaria and Romania.

the presumably relatively long run-up period to the euro allows a thorough and conscientious preparation in which use can be made of the experience and expertise of the current Eurozone countries. A profound discussion is certainly a substantial part of this process.

Joining a monetary union might not only bring about a discrete change in inflation expectations. In the first place it implies the abandonment of an autonomous monetary and exchange rate policy, two important adjustment channels. It is natural to conjecture that the interest rate set for a union consisting of not entirely homogeneous economies might deviate from the interest rate that would be convenient for a single country, with all consequences for inflation. In order to analyse this issue econometrically we estimate as a side-product the time-varying real natural interest rate by means of an unobserved components model based on Harvey (1989). Horváth (2007) is to our best knowledge the only study so far to estimate the (nominal) natural interest rate for the Czech Republic using various specifications of Taylor-type rules and our results can thus provide a useful comparison. In addition, however, in the spirit of Crespo-Cuaresma et al. (2004) we also estimate the natural rate of interest purged from the risk premia which is probably the first attempt to do so for the Czech Republic.

Notwithstanding the abandonment of monetary autonomy, the most conventional argument put forward in the discussion on inflationary impacts of the euro centres around the substantial price level gap with respect to the Eurozone which is characteristic for a catching-up economy as the Czech Republic. Price level convergence is carried out via nominal exchange rate appreciation on the one hand, and a positive inflation differential relative to the Eurozone on the other. It is usually argued that the latter is driven by typical transition phenomena such as lower productivity growth in the non-tradable sector relative to the tradable sector (e.g. Holman 2006) or gradual shifts to higher quality goods (Bruha and Podpiera 2007). These authors conclude that if the nominal appreciation channel is closed after fixing the exchange rate to the euro inflation will inevitably have to accelerate in order to keep the pace of real appreciation. We will have a close look at this line of argument but we do not quite share this opinion and believe that the size of the possible inflation rise ascribable to the lack of nominal appreciation will depend essentially on the extent with which exchange rate movements are passed through onto consumer prices. If this exchange rate pass-through is strong then the nominal appreciation has a significant dampening effect on inflation thus rendering a peg more costly.

On January 1, 2007 Slovenia joined as the first NMS the European Monetary Union and was thus also the first country to introduce the euro as book money and cash on the same date. Hence, barring the effects related to abandonment of autonomous monetary policy we also have to investigate whether the changeover of coins and notes can have some impact on inflation. Though this should be in theory a mere nominal event it turns out that there are several luring risks for the changeover to be taken advantage of. However, we will also argue that even though these risks can probably not be eliminated altogether they might be substantially reduced as the experience of some current Eurozone members suggests.

After a qualitative discussion we test the relative importance of those inflation channels through which the euro might strike along with other relevant cyclical, external and structural inflation factors based on Égert (2007a,b) in an autoregressive distributed lag (ARDL) model. Since standard model selection criteria do not lead to consistent results, in line with the recommendations made by Yang (2004) we combine all models under a proper weighting scheme by employing the Bayesian Model Averaging technique. In order to reduce the intractably large model space we use the Markov Chain Monte Carlo Model Composite algorithm. The results suggest that a harmonization of the business cycle with the Eurozone and a low inflation environment are essential for a smooth inflation path. However, the latter should not be enforced by non-standard measures such as withholding necessary adjustments of regulated prices. In contrast, consistently with the existing literature we do not find much evidence either for the Balassa-Samuelson-effect or a strong exchange rate pass-through.

This chapter is structured as follows. The following section provides some stylized facts on the inflation development of the first-wave Eurozone countries and the three countries that have opted out. Against this backdrop section 3.3 seeks to identify and flesh out those relevant inflation forces that are likely to be affected in the wake of the euro adoption via both the common monetary policy as well as the cash changeover. Anticipating the subsequent econometric test we also describe the operationalization of these channels for the empirical estimation and the data. In section 3.4 we explain the method, describe the remaining variables included in the model and sketch out the results. The last section wraps up and derives interesting policy conclusions.

3.2 Stylized facts

Before we delve into the channels through which the euro might affect inflation in the Czech Republic it is helpful to have a brief look at some stylized facts regarding the European economic and monetary unification process and its effects on inflation. In June 1988 the European Council approved the objective of establishing an Economic and Monetary Union (EMU). The necessary revision of the legal framework was provided in the Treaty on European Union (Treaty) which was signed in Maastricht in February 1992 and came into force on November 1, 1993. The actual birth of the EMU in the sense of a single monetary policy under the responsibility of one European Central Bank (ECB) dates back to January 1, 1999. On that day the third and final stage of the economic and monetary unification process was launched with the irrevocable fixing of the exchange rates of the currencies of the 11 initially participating Member States (henceforth also referred to as EU-11) and the euro thus became their legal tender.[2] Greece joined the club on January 1, 2001.[3] One year later, on January 1 2002, euro coins and banknotes were put into circulation and thus replaced the 12 national currencies which had, since 1999, represented only different denominations of the single currency.[4] Except for Denmark and the United Kingdom which negotiated an opt-out the Treaty obliges all current and future EU members to adopt the euro as their legal tender after meeting the so called Convergence (or Maastricht) criteria. Besides the criterion on price stability, government finances and long-term interest rates each euro-candidate must have participated in the new exchange-rate mechanism (ERM II) for at least two years preceding the examination of the situation without any break and without severe tensions.[5] To this end, Sweden as the only 'old' EU member without an opt-out from the Treaty has not yet introduced the euro as it has not met the exchange rate criterion. In contrast, Slovenia, Malta and Cyprus have been as of today the only NMS to extend the Eurozone.

Hence, if we want to collect some stylized facts on the inflation impact of the euro in the 12 initial countries we have to make a clear distinction between the period following the launch of the common currency and monetary policy on January 1, 1999 (and January 1, 2001 in case of Greece, respectively) on the one hand, and the period after January 1, 2002 on which date the cash changeover took place on the

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[2]Belgium, Germany, Ireland, Spain, France, Italy, Luxembourg, Netherlands, Austria, Portugal, Finland

[3]The initial 11 EMU countries and Greece will be referred to as EU-12.

[4]For details on the history of the Economic and Monetary Union see e.g. www.ecb.eu.

[5]See http://eur-lex.europa.eu/en/treaties/dat/11992M/htm/11992M.html

other. In case of countries that joined or will join the Eurozone later both effects, the one brought about by the abandonment of monetary policy autonomy and the one caused by the currency changeover, fall together. Moreover, for reasons that will be addressed in more detail below, the combination of these two effects might turn out relatively even more pronounced if compared to the preceding run-up period.

After a visual inspection of Figure 3.1 depicting inflation paths for EU-11 it strikes that in most countries inflation[6] eased somewhat or remained broadly stable in the year following the realisation of EMU. In 2000 the cycle reached a turning point and inflation rose until it peaked in the course of 2001 followed by a slowdown, differently pronounced in the respective countries. Also the subsequent path suggests a recognizable cyclical pattern up until 2007 similar in most EU-11 countries. The average inflation rate in the Eurozone rose from 1.5% in the two years preceding the EMU creation to 1.9% between 1999 and 2006. In contrast, if we compare the same periods in two EU but non-Eurozone countries, Sweden and the UK, inflation remained stable in the former at 1.5% and declined from 1.7% to 1.1% in the latter. Thus, these numbers might evoke the hypothesis that in general the common monetary policy has led to inflation acceleration. Such a conclusion, however, would be premature. Figure 3.2 displays three panels in which the inflation cycle in three non-euro EU countries, Denmark, Sweden and the UK is drawn against the inflation paths in an appropriate EMU-counterpart, with some lag if needed. In most EU-11 countries the inflation cycle happened to head downwards to its local minimum prior to 1999 and made a U-turn after that. Therefore it is not surprising that inflation appears to have accelerated in the wake of a common monetary policy. However, as can be seen in the first panel of Figure 3.2 the inflation paths in Denmark matched the one in Finland very closely. Correspondingly, very similar to the Eurozone performance the average inflation in Denmark thus rose from 1.6% in 1997-1998 to 2% between 1999 and 2006. In Sweden and the UK where the cycle appears to be lagged by some 6 and 14 months (second and third panel of Figure 3.2), respectively, the downward paths to the local minimum corresponds to the period 1997-1999 in case of Sweden and 1997-2000 for the UK. If we compare the average inflation in these two respective periods to the average inflation in the second half of the sample we find that price increases accelerated by 0.5 percentage points in Sweden and 0.2 percentage points in Britain.

Concerning the inflationary effect of the euro changeover in January 2002 a simple eyeballing Figure 3.1 suggests what has been concluded by several studies after

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[6]Measured by a 12-month-moving average of the y-o-y change of the HICP.
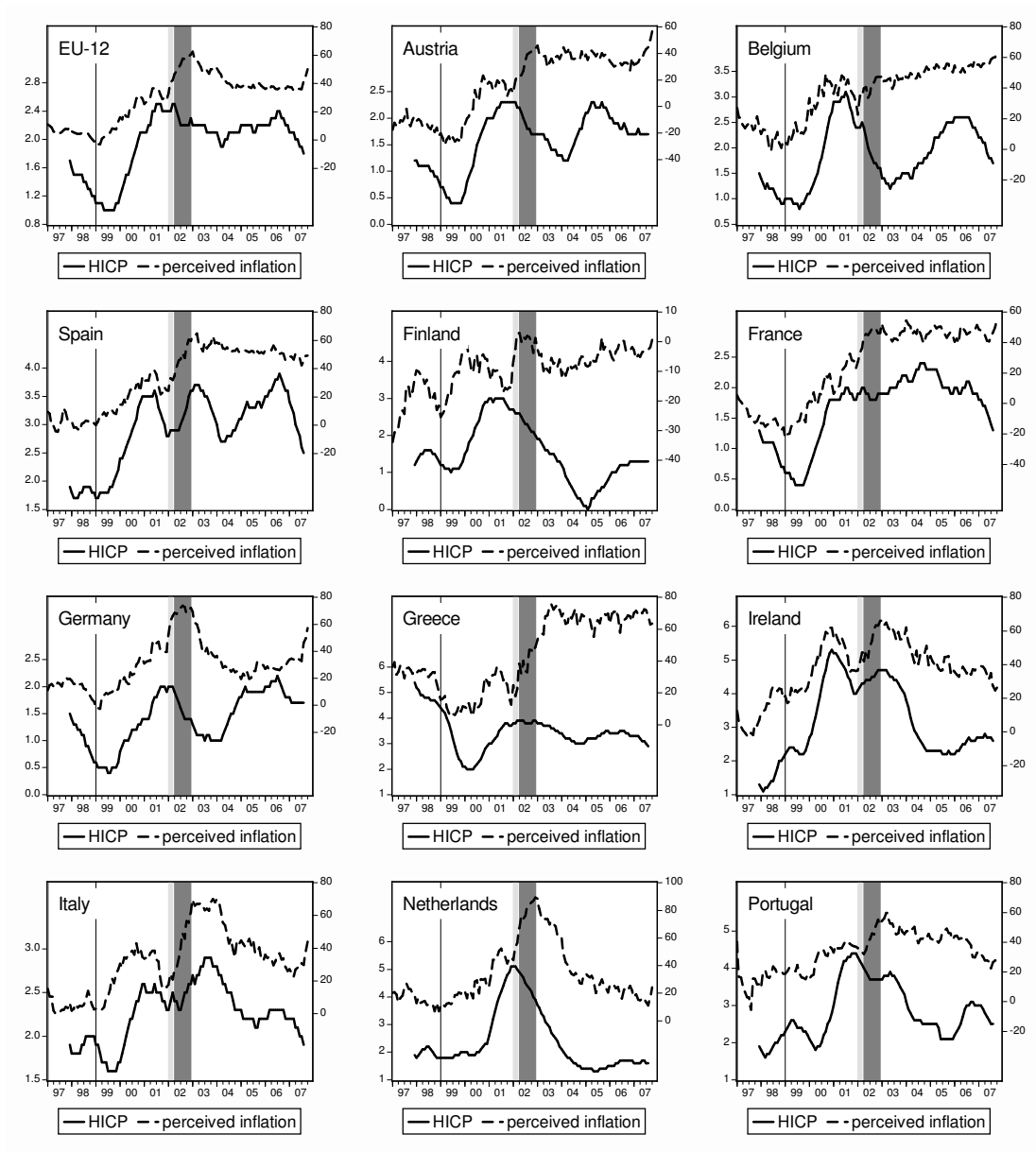
Figure 3.1: HICP (12-month moving average of the y-o-y change) and perceived inflation in the EU-11
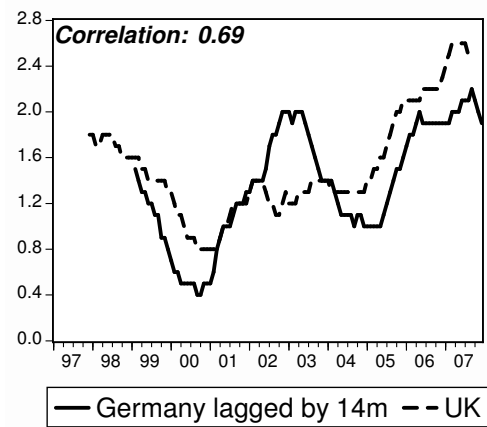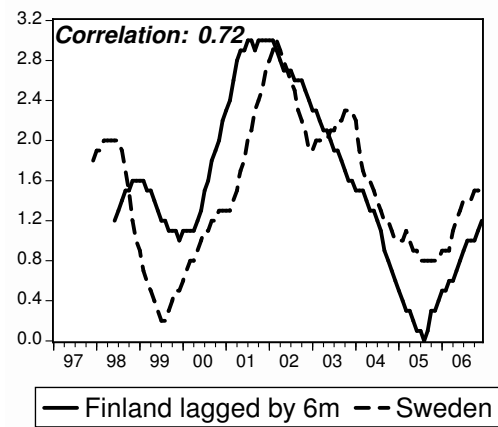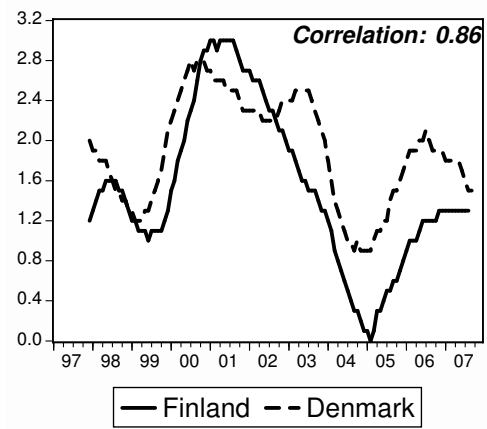
Figure 3.2: HICP (12-month moving average of the y-o-y change) in the non-EMU countries relative to selected EMU

a much more sophisticated and comprehensive analysis[7]: Excessive price increases about which the public had been concerned prior to the introduction of euro coins and notes[8] did not occur. Even though there was a slight inflation acceleration in some countries and also in the EU-12 as a whole in the first three months of 2002 (light grey window), the impact was transitory and in most countries price increases slowed down again within a couple of weeks (dark grey window). Hence, on balance, there is no evidence underpinning the notion that the euro changeover had a remarkable impact on the overall consumer price inflation. Moreover, as Hobijn et al. (2004) note, the inflation experience in the initial 12 EMU countries was not very different from that in Denmark, Sweden, and Britain. However, while the European Commission estimates that the changeover impact on the overall HICP inflation ranged between 0.12 and 0.29 percentage points in 2002[9] there were significant and persistent price increases in some specific industries, particularly services such as restaurants, cafes and hairdressers.[10] Therefore, while the average annualized monthly inflation rate in the EU-12 restaurant and cafe sector over the period January 1995 through March 2004 was 2.8%, in January 2002 inflation in that sector went up to 15.6% (see Hobijn et al. 2004).

In parallel to the debate on the impact of the EMU on inflation the euro gave rise to an unprecedented break in the consumers' perception of price increases that sharply contradicts actual inflation figures. This can be seen in Figure 3.1 which displays apart from the HICP also the so called balance statistic of the inflation perception survey carried out among consumers by the European Commission on a monthly basis. The balance statistic is calculated as the difference between a weighted proportion of respondents stating that prices have risen and those stating that prices have fallen or stayed about the same.[11] The distance between the two curves in Figure 3.1 thus measures the size of the perception gap. While in most countries inflation perception followed the actual inflation rather closely prior to the changeover in January 2002 the gap widened substantially immediately thereafter and continued

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[7]See e.g. Deutsche Bundesbank (2004), Eife (2006) or Hobijn et al. (2004).

[8]See Commission of the European Communities (2001).

[9]The effect was probably largest in the Netherlands, 0.6 percentage points (see Folkertsma et al. 2002) and in Italy, 0.5 percentage points (Aucremanne et al. 2007).

[10]See http://ec.europa.eu/economy_finance/euro/faqs/faqs_16_en.htm

[11]There are six possible answers to the question: "How do you think that consumer prices have developed over the last 12 months?". A(1) "risen a lot"; A(2) "risen moderately"; A(3) "risen slightly"; A(4) "stayed about the same"; A(5) "fallen" and A(6) "do not know". The balance statistic is thus computed as: $B_{it} = A_{it}(1) + 0.5A_{it}(2) - 0.5A_{it}(4) - A_{it}(5)$ (See: http://ec.europa.eu/economy_finance/indicators/businessandconsumersurveys_en.htm)

to grow until early 2003. Then it either remained stable or shrank until end 2006 and seems to have been broadening again since. Aucremanne et al (2007) confirm the anecdotal evidence by showing that this gap is econometrically significant and that there was no such break in the inflation perception in 2002 in the control panel of those three countries that have opted out of the Monetary Union so far.

We can conclude this section by looking at the data from Slovenia which is often cited as the obvious evidence for the inflationary impact of the euro. Although prices actually went down in Slovenia m-o-m in January and February 2007 the y-o-y inflation rate, indeed, accelerated over the course of 2007 and amounted to 5.7% in December 2007. However, the popular view to blame the euro for this pick-up is questioned if we contrast the inflation development in Slovenia with the one in the Czech Republic. As can be seen in Figure 3.3, inflation rates in the Czech Republic and Slovenia have co-moved virtually over the entire sample period. While the level gap widened somewhat in 1999 and reached its maximum in 2003 since 2004 the gap has become infinitesimally narrow with a correlation 0.8 between the two series. Since the euro introduction in Slovenia in January 2007 the correlation amounted to as much as 0.94 as the inflation rate in the Czech Republic also rose over the course of 2007 and reached 5.5% in December. In fact, according to the study by Deloitte Consulting (2007) there was only a slight acceleration of inflation in Slovenia attributable to the euro introduction amounting to mere 0.3 percentage points or less. Hence, the rise of the inflation rate must have been primarily fuelled by other determinants such as the economic overheating and external factors such as oil and food prices. Particularly the latter have increased by twice as much as in other EU countries due to, as anecdotal evidence suggests, insufficient competition in the retail market (see Bank of Slovenia 2007 and e.g. Eurobusiness 2007). This latest experimental evidence thus appears to corroborate previous experience: while the actual impact of the euro on inflation seems to have been rather moderate, inflation perception literally skyrocketed in parallel and the gap between actual and perceived inflation has reached exceptional levels.
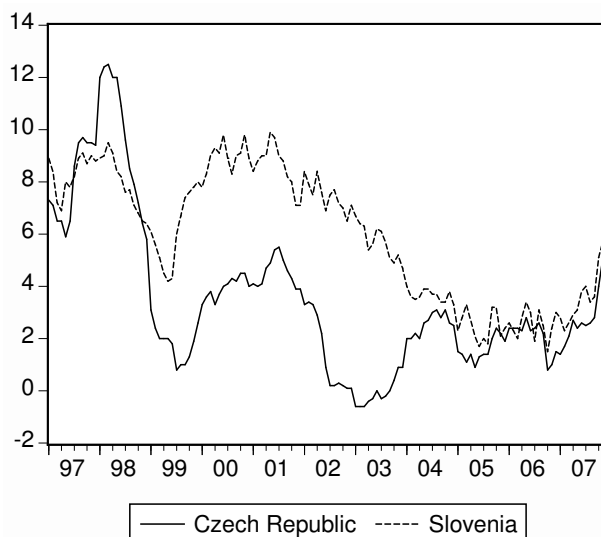
Figure 3.3: HICP (y-o-y change) in the Czech Republic and Slovenia

3.3 Factors through which euro adoption might impact inflation

3.3.1 Long-term factors

3.3.1.1 Price level convergence and nominal appreciation

The usual argument put forward for why the euro adoption is expected to increase inflation goes approximately like this: The relative price level in the euro candidate country is substantially lower than in the EMU. Therefore, there is a high potential for price catch-up both in the tradable and non-tradable sector. While in the former price convergence is driven for instance by a gradual shift to higher-quality-goods that are typically more expansive (Bruha and Podpiera 2007), in the non-tradable sector it is particularly the Balassa-Samuelson (B-S) effect.[12] At the same time the respective transition economy is often characterized by a trend nominal appreciation of its currency which in combination with the price catch-up process brings about a significant appreciation of the equilibrium real exchange rate. However, if the nominal appreciation is no longer possible after the euro adoption it follows that the real exchange rate appreciation will take place completely via the relative inflation

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[12]For a recent formal discussion of the Balassa-Samuelson effect see e.g. Égert (2007).

channel as it is the only one remaining. To put it with Suster et al. (2006):

In the long run it can be expected that after joining the euro area the inflation in Slovakia will be higher than inflation in the euro area by a margin which will be about equal to appreciation rate of the equilibrium real exchange rate of koruna against euro. (…) If the current annual appreciation rate of the equilibrium real exchange rate of koruna was maintained at the level of approximately 2-3% also after entry to the euro area, the inflation in Slovakia could be in a short run higher than in the rest of the euro area by such a difference.

Concerning the Czech Republic, Holman (2006), a Board Member of the Czech National Bank (CNB), estimated that if the Czech Republic had adopted the euro back then annual inflation would have hovered around 6% after rising by about 3 percentage points which corresponds to the average annual appreciation of the koruna.[13] Mach (2003) goes even a dimension further and predicts that inflation will return to annual 10% from the first half of the 1990s.[14]

If our reading of this argument sequence is correct, it makes two implicit assumptions. Firstly, it seems that a perfect pass-through of nominal exchange rate appreciation on consumer prices is assumed so that a x% nominal appreciation dampens the inflation rate by x percentage points. Secondly, this line of argument evokes the notion that in the equation $\Delta e^r = \Delta e + (\pi - \pi^*)$ where Δe is the annual change of the nominal exchange rate defined as euro per unit of domestic currency and $(\pi - \pi^*)$ stands for the annual inflation differential between the home country and the Eurozone, the path of the real exchange rate e^r on the left hand side is somehow exogenously determined and can only be pushed through either of the two channels on the right hand side. In other words, the argument assumes that price level convergence ($e\frac{p}{p^*}$, where e is the exchange rate (euro per unit of domestic currency) and p and p^* the price level in, respectively, the home country and the Eurozone) is not the result of the transition process but the driving mechanism per se. In our opinion, however, this sequence of arguments puts the cart before the horse and it is worth to think about it in more detail.

Certainly there is enough scope for price level convergence in the transition EU

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[13]This estimate was published in June 2006 after the annual inflation rate (measured by the CPI, the target index of the CNB) had reached 1.9% in 2005 and averaged 2.9% between January and May 2006.

[14]For a slightly different perspective of the same issue see Bruha and Podpiera (2007).

economies both in the tradable and the non-tradable sector. In the Czech Republic the price level of total goods reached nearly 74% of the EMU level in 2006, while the price level in services amounted to mere 42% of the Eurozone.[15] Apparently, price level convergence can only materialize through nominal exchange rate appreciation or a faster rise of the price level in the transition economy relative to the Eurozone. If one of the channels is abandoned in the wake of the euro adoption, the speed of price level convergence thus will be determined by definition solely by the relative inflation. Unlike the cited authors, however, we do not see the mechanism due to which price level convergence necessarily needs to keep the pace determined prior to the euro adoption and has thus be inevitably reflected in higher inflation rates after fixing the exchange rate.

Despite the theoretical concept of the law of one price an abundant empirical literature has confirmed that price levels significantly differ across the EU and converge rather sluggishly[16]. As Égert (2007a) documents and Figure 3.4 visualises the ranking of the price levels strongly correlates with the ordering of the countries in terms of GDP per capita. This observation suggests that the price level is determined to a great extent by economic development. Notwithstanding, one would expect that a reduction of trade barriers (e.g. as a result of the European integration process) and/or higher price transparency after the adoption of a common currency should enhance the arbitrage process and reinforce price level convergence. However, the literature seems to conclude almost unanimously that while price level convergence in the EU and the Eurozone was spurred by the establishment of a common market at the beginning of the 1990s the introduction of a common currency had no significant effect (see e.g. the discussion in Crespo-Cuaresma et al. 2007). Hence, it follows that the price level convergence in the Czech Republic will gradually continue as the country becomes richer in terms of GDP per capita. However, it seems highly unlikely that the convergence process could receive a discrete one-time boost after the euro adoption as a result of better price transparency.

But if it is not for higher price transparency are there any other forces which could discretely change in the wake of the euro adoption and make inflation suddenly accelerate? To answer this question in what follows we will have a close look at the factors behind the price level convergence process usually put forward in the

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[15]Source: New Cronos/Eurostat

[16]For instance Crucini and Shintani (2002) find substantial price level differences that persist over time while Taylor and Taylor (2004) argue that such differences can be explained by the existence of transaction costs. For further references see Crespo-Cuaresma et al. (2007).
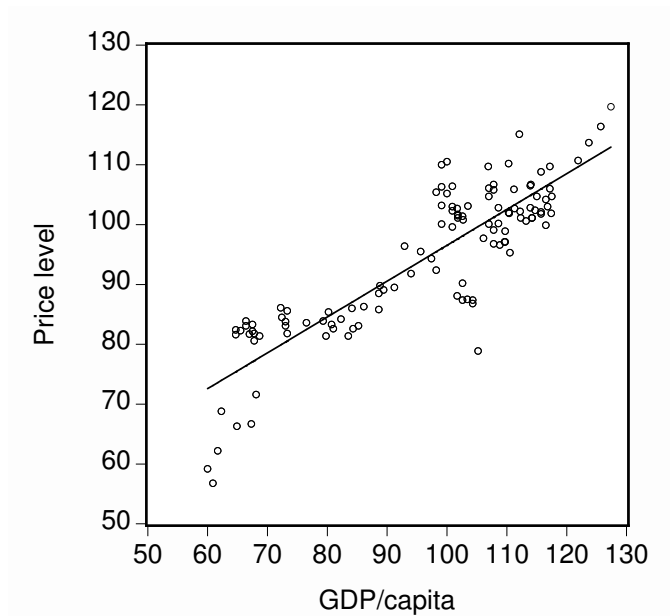
Figure 3.4: Price level vs. GDP/capita

literature such as the B-S-effect, deregulations of administered prices, the shift to higher-quality goods, changes in pricing to market practices and the nominal exchange rate appreciation. However, in the context of the present paper we are not only interested in the impact of these factors on price level convergence and inflation but, particularly, in how their inflation impact will change after the euro adoption.

Goods Prices
As Égert (2007a) describes in detail according to an extended Engel's Law richer growing households tend to spend less of their budget on food and their consumption pattern also typically shifts towards goods of higher quality usually manifested in brands. Naturally, higher-quality-goods are associated with higher prices. The shift towards better goods on the demand side is accompanied by a catch-up-process in terms of quality also on the domestic supply side. However, while this gradual preference change contributes substantially to the price level convergence it should not be reflected in higher inflation rates as it is not the same good that is getting more expensive (see Égert 2007b). Next Égert (2007a) documents that even prices of homogenous goods such as cars differ across euro area countries. This fact suggests that producers take into account the disposable income in the particular economy

and adjust their price setting to the purchasing power of households in that country. As the income gap closes in the wake of the convergence process also prices of those goods will rise and, in this case, push up also the inflation rate. In our estimation later on we will approximate these two effects by linearly interpolated data on GDP per capita at PPP obtained from the Eurostat New Cronos data base. It should be added that both the quality effect and the pricing-to-market-effect will take place irrespective of the exchange rate regime and should thus not be directly affected by the euro adoption.

Balassa-Samuelson-effect
Some of the older studies estimated a rather sizable impact of the Balassa-Samuelson-effect (B-S-effect) on inflation in the Czech republic, contributing e.g. 4.3 percentage points in Golinelli and Orsi (2001) or 2.8 percentage points according to Sinn and Reutter (2001). Then, however, Flek et al. (2002) found that the inflation impact of the B-S-effect was very low (partially negative, depending on the methodology) and thus negligible. As the topic has become vary fashionable a whole plethora of research studies has emerged more recently confirming the finding that the inflationary impact is rather low in the Central and Eastern European (CEE) transition economies in general and in the Czech Republic in particular. Égert, Halpern and MacDonald (2006) provide a very comprehensive survey of the literature on this issue and conclude that the inflation contribution of the B-S-effect is close to 0 in the Czech Republic. Admittedly, these results are somewhat puzzling given the massive productivity gains in manufacturing in these countries over the course of transition. Égert (2007b) puts forward several reasons for the small size of the B-S-effect which basically question the assumptions of the B-S-model: i) Productivity growth may not necessarily lead to a high wage growth in tradables, ii) due to imperfect homogeneity of labor wages may fail to equalise across sectors, iii) in some countries there have been large productivity gains also in the non-tradable sector and, iv) perhaps most importantly, productivity-driven inflation acceleration in the non-tradable sector is reflected in the overall prices index only to a small extent because of the low share of non-tradables in the overall consumer basket. While Kovacs (2002) believes that the low impact of the B-S-effect can already give an upper estimate that should become even weaker in the course of the catch-up process as productivity differentials moderate Flek et. al (2002) disagree. According to them the productivity growth in the tradable sector will accelerate as the catching-up process moves on because real convergence in the Czech Republic did not record any remarkable progress throughout the 1990s. One of the most recent updates by Égert (2007a) seems to support this hypothesis as his upper bound estimates have increased com-

pared to earlier results and range between 0.7 and 1.9 percentage points for the Czech Republic.

Three different approaches have been employed in the empirical literature to derive the size of the impact on inflation imputed by the B-S-effect. Following Égert (2007b) we will use in our econometric test the simple accounting framework which assumes that changes in the productivity differential are proportionately translated into the relative price change of non-tradables $\Delta p^{NT} - \Delta p^T = \Delta prod^T - \Delta prod^{NT}$. At the same time, the impact of non-tradable inflation (Δp^{NT}) in excess of tradable inflation (Δp^T) is established by the share of non-tradables in the CPI basket $(1-\alpha)$: $\Delta p = (1 - \alpha)(\Delta p^{NT} - \Delta p^T)$. Therefore, the part of inflation ascribable to the B-S-effect (p^{B-S}) can be written as

$$\Delta p^{B-S} = (1 - \alpha) \left(\Delta prod^T - \Delta prod^{NT} \right). \tag{3.1}$$

The quarterly data is obtained from ARAD, the CNB's time series database and the Czech Statistical Office (CSO). Productivity is calculated as gross value added in the respective sector divided by total hours worked. We approximate the tradable sector by manufacturing (D in the 17-sectoral decomposition of Eurostat) while the non-tradable sector consists of construction (F in the Eurostat nomenclature) and market services (G-K in the Eurostat nomenclature). Services where prices are regulated such as agriculture, energy, water supply and public administration are excluded because prices there do not necessarily react to productivity changes in a market manner. The parameter $(1 - \alpha)$ thus amounts to 0.25 (see Égert (2007a). If the B-S-effect plays a significant role in the Czech Republic then the productivity differential should have a positive impact on inflation. However, also for the B-S-effect it should be borne in mind that it is completely independent of the exchange rate regime. In other words, even if there was a strong inflationary effect stemming from relatively higher productivity gains in the tradable sector it will take place irrespective of the currency used in that country.

Regulated prices
While prices of most goods and services have been set free and their inflation rates reflect the market forces, there is still quite a considerable share of goods and services for which prices are still administered. Following the broad concept of price regulated service categories proposed by Égert (2007b) our definition includes refuse and sewerage collection, medical, dental and paramedical services, hospital services, passenger transport by railway and by road, postal services, education, social pro-

tection, cultural services, rents and energy prices related to housing.[17] Using the three-digit COICOP disaggregation level of the HICP from New Cronos/Eurostat the share of administered prices in the consumer basket amounts according to this definition to approximately 20% in the Czech Republic, a number very similar to the euro area level and we can thus conjecture that increases of administered prices will have a significant impact on the overall inflation. Moreover, this impact is reinforced by the fact that the inflation of regulated services has been almost persistently above the inflation rate of the overall consumer basket as can be seen in Figure 3.5. The chart thus suggests that if regulated prices are adjusted the modification is substantial in relative terms.
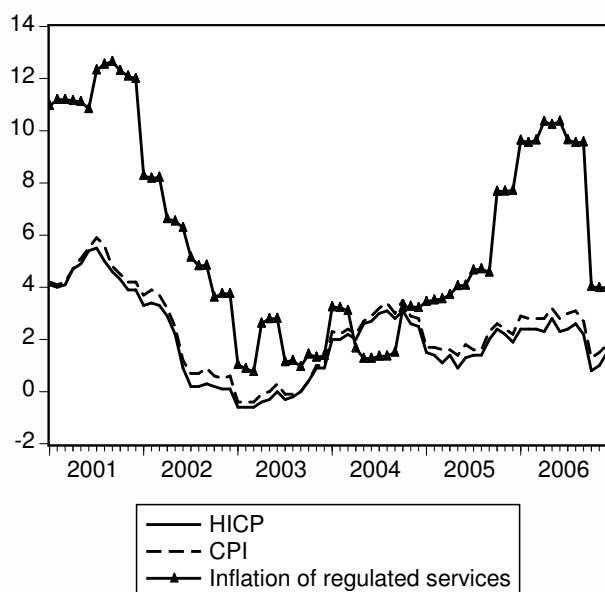
Figure 3.5: Inflation of price regulated services vs. HICP/CPI

Égert (2007b) puts forward three main reasons why regulated prices are likely to rise faster than the overall CPI also in the future. Firstly, prices of some of the regulated categories might still be below the cost recovery ratio (e.g. rents) and/or below the EU level such as electricity and gas which reached in the Czech Repub-

[17]Though price regulation mostly concerns services there are also some goods such as pharmaceutical products, alcoholic beverages or tobacco whose prices are not entirely market determined and could be included in the definition.

lic in 2006 some 70% of the EU-level. Secondly, the capital stock of some sectors such as railways or public transportation is outdated and urgently needs massive investments. Unless the latter can be fully financed from public subsidies and/or EU Structural and Cohesion funds price increases will be inevitable. Lastly, the degree of regulation which is closely interlinked with the ownership structure and level of liberalisation in those sectors is often subject to hot political debates. Some politicians subject to the political business cycle might be reluctant to price increases and prefer to postpone them for later, especially after elections. At the same time in some parts of the political spectrum there might be quite a strong resistance to privatisation or market liberalisation, if they are possible in the particular sector, which would increase efficiency and mitigate price increases.

The last point is certainly also strongly related to the question about how much regulated prices will be affected by the euro adoption. Generally speaking, if a country interested in adopting the euro does not meet the Maastricht price stability criterion with a convenient margin - which might be the case in transition CEE for reasons described above - it is natural that policy makers will not be particularly keen on putting the inflation rate further under strain by adjusting regulated prices more than necessary. Hence, since sustainability of the inflation criterion is not formally required it might provide incentives for the country's political representatives to postpone more than indispensable hikes of administered prices until after the euro adoption. If the latter was the case, inflation might indeed appear to have accelerated as a consequence of the euro introduction which would then be the easiest patsy to blame. A similar materialization of this incentive structure seems to have happened in Greece, where the fulfilment of the inflation criterion prior to the euro adoption was helped by cuts in indirect taxes and gentlemen's agreements between the government and commercial and industrial enterprises, as well as service providers (see ECB 2000).[18]

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[18]Annual inflation in Greece soared from 2.1% in 1999 to 3.9% in 2002 which might have been, to some extent caused by the non-renewal of the gentlemen's agreements. For instance the Athens News Agency reported on August 22, 2002 in 'Finance minister pledges intensified efforts against inflation':

"This new and worrying surge in inflation signals a clear-cut failure of the government's policy of 'gentlemen's agreements (...) These agreements are based on an unbridled operation of markets and a lack of effective checks and measures against violators", Dimitris Papadimoulis, economic and social affairs spokesman of the Coalition of the Left and Progress. He added that state-owned companies had often led the pack in unjustified rate rises, setting an example for the rest of the market, which meant that price hikes were transferred down to the consumer.

Nominal exchange rate and the exchange rate pass-through (ERPT) on Consumer Prices

If talking about the nominal exchange rate we have to distinguish clearly between its role in the price level convergence process and its impact on inflation. Nominal appreciation, for instance, increases the price level expressed in euros immediately (numerator in $\frac{ep}{p^*}$ with variables defined as above) and thus speeds up the price level convergence. For the inflation development, however, not necessarily the nominal appreciation as such matters but the pass-through of exchange rate movements on consumer prices. On the hand, the higher the ERPT and the higher the nominal appreciation rate, the bigger the dampening effect on the overall inflation. On the other hand, however, the higher the ERPT the less pronounced will be the effect of nominal appreciation on price level convergence as in $\frac{ep}{p^*}$ the rise in e is compensated by less strongly rising p as a result of the ERPT. Or conversely, a lower ERPT implies a more significant convergence of price levels but provides less help in curbing inflation.

The extent to which exchange rate changes are translated into domestic prices hinges primarily on the price setting practice and the composition of imports. The ERPT will be zero if firms set prices of imported goods in the local currency while the exchange rate movements will be completely reflected in the import price if firms invoice in the export country's currency. However, the eventual effect of exchange rate changes on consumer prices will vary depending on whether the country imports rather final goods included in the consumer basket or intermediate goods by which the ERPT on consumer prices will be watered down in the production process. In addition, Égert (2007b) argues that expectations are another important factor for the pass-through strengths. In an inflation targeting regime where expectations are anchored by credibly communicated inflation targets and not by the exchange rate itself the ERPT should be low. Moreover, Taylor (2000) coins the idea that the fact that the ERPT has declined over time and that it is typically lower in developed market economies is ascribable to generally lower inflation rates. In a microeconomic model he shows that with declining inflation people perceive cost changes as less persistent, adjust prices less frequently and thus incorporate exchange rate changes less often.

Given these arguments, how important is the exchange rate pass-through in the Czech Republic? On the one hand, it is certainly a rather small and open economy with a relatively high share of imported final and intermediate goods. On the other
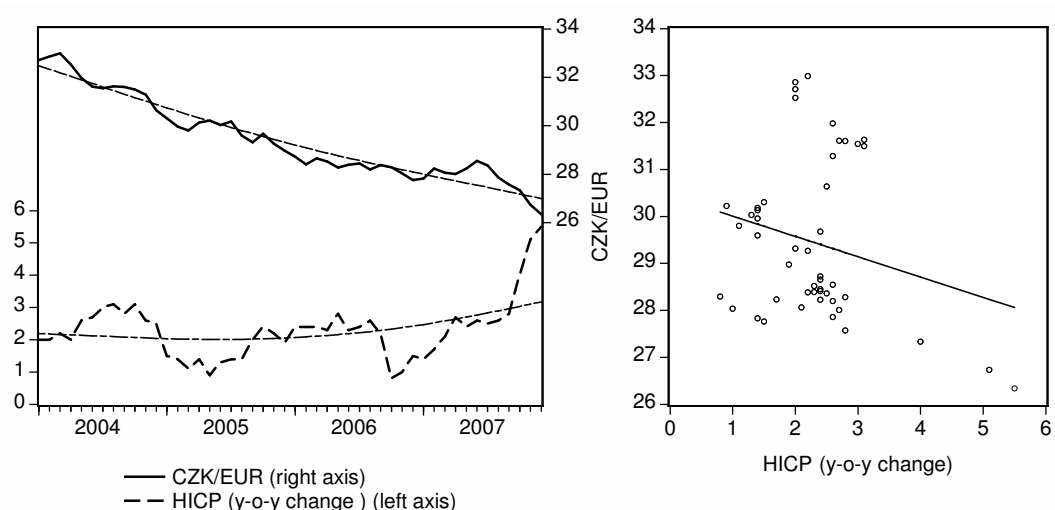
Figure 3.6: Exchange rate vs. HICP (y-o-y change) in the Czech Republic

hand, the CNB operates an inflation targeting regime with a long and successful track record and the Czech economy has experienced a strong disinflation over the course of the transition period. Before looking at empirical results Figure 3.6 provides a visual appetizer for the strength of the ERPT in the Czech Republic since 2004. It depicts the exchange rate (koruna vs. euro) and inflation development in two forms, as a line (with the corresponding Hodrick-Prescott-trend) and a scatter plot. As can be seen in the left panel, while there has been a clear appreciation tendency of the koruna vis-à-vis the euro, the inflation rate has drifted upward. The reverse trend of the two series has been particularly strong since mid 2007. The right panel provides another phrasing of the same content as it suggests a rather weak, and, worse still, negative relationship between exchange rate appreciation and the inflation rate.

Hence, it is not surprising that also empirical studies come to the conclusion that the ERPT in the Czech Republic is quite tenuous. While e.g. Ca' Zorzi et al. (2007) find that the one-year ERPT amounts to about 0.6 many other studies estimate it close to 0 or sometimes even negative (see e.g. Darvas (2001) or Coricelli, Jazbec and Masten (2004)). Coricelli, Égert and MacDonald (2006) report the average of non-negative, effective pass-through estimates from a literature survey reaching mere 0.2. In the context of the present paper we are interested in the pass-through of the koruna-euro-exchange rate. Therefore, we will approximate the pass-through from

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the Eurozone by including an openness indicator defined along the lines of Égert (2007a) as the nominal exchange rate of the euro vis-à-vis the koruna multiplied by imports from the 'old' EU-15 Member States[19] relative to GDP. An increase of this indicator due to an appreciation of the koruna and/or higher imports from the EU-15 would thus be expected to result in a lower inflation and should therefore carry a negative sign. The quarterly available data on imports and GDP were obtained from Eurostat and had to be linearly interpolated while Thomson Financial provides exchange rates on a daily basis and we took monthly averages. Alternatively to the openness indicator we will test the relevance of the exchange rate on inflation by including a simple annual rate of change of the exchange rate.

Exchange rate development
Strictly speaking, it is only secondary how relevant the exchange rate pass-through is now or has been up to now. This paper wants to shed some light on how much will inflation be affected after the euro adoption. Therefore, even if the substantial nominal trend appreciation of the koruna with respect to the euro had a considerable dampening effect on inflation at this moment it can not be concluded straightforwardly that inflation will rise by the same amount after joining the EMU. Between 1993 and 2007 the koruna gained on average some 1.4% against the euro in nominal terms while the appreciation was much stronger in the recent years (about 3% since 2000). The relevant question is though, how will the euro adoption (possibly also in other countries) affect those factors that stand behind the nominal appreciation. If, for instance, the value increase of the koruna vis-à-vis the euro was driven predominantly by speculative capital flows which will disappear after joining the EMU anyway, the costs of non-existent dampening effect on inflation would be low. On the contrary, if the appreciation was driven primarily by e.g. FDI it could be argued that since such capital flows will persist also after the euro adoption the inflationary effect will be rather high. Even then though, what would happen to FDI-flows if the Czech Republic was the only country in the region without the euro? Would not investors prefer the euroised neighbours? These are just some issues meant to exemplify the rather big amount of uncertainty related to this question.

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[19]Apart from the EU-12 defined above EU-15 contains also the UK, Denmark and Sweden.

3.3.1.2 Natural interest rate too low?

It is not only the exchange rate that is abandoned after the EMU entry but also the autonomous monetary policy run by the national central bank.[20] Unless the monetary union is a perfectly optimal currency area it is natural that the interest rate set for so many countries might not necessarily be the most convenient one for each particular member state. Hence, it could happen that with the introduction of the euro the country adopts too loose a monetary policy bringing about inflation acceleration. In other words, if the country's natural interest rate, defined along the lines of Laubach and Williams (2001) as the short-term interest rate at which output converges to potential and inflation is stable, is higher than the interest rate set for the Eurozone then inflation might speed up. In order to get some feeling for how big this risk is, we have to answer two questions:

1. Does the natural interest rate in the Czech Republic differ significantly from the one in the Eurozone?
2. If it does, how sensitive is inflation to deviations from the natural interest rate?

Since the capital/labor ratio which determines the relative price of capital changes over time, in transition economies typically increases, also the natural interest rate defined as above has to be time-varying (see Lipschitz et. al 2006). Different methods have been employed in the literature to estimate time-varying natural interest rates (TVNIR), ranging from various versions of the Taylor rule in Orphanides and Williams (2002), a simple macroeconomic model in Laubach and Williams (2001) where equilibrium interest rates and output gap are modelled as unobserved components to a DSGE model with sticky prices and wages in Smets and Wouters (2002). To our knowledge, there has been only one paper thus far estimating a nominal, time-varying policy neutral rate for the Czech Republic by Horvath (2007). Based on various forward- and backward-looking specifications of simple Taylor-type monetary policy rules he finds that the nominal policy neutral rate decreased form some 5% in 2001 to around 2% at the end of 2005 and increased subsequently to some 2.5% in 2006. Moreover, he concludes that the deviation of the actual nominal interest rate from the policy neutral (equilibrium) rate is a useful predictor for future level and change of inflation.

We, in contrast, will proceed in the spirit of Crespo-Cuaresma, Gnan and Ritzberger-

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[20]To which extent the monetary policy can be autonomous in small and open economies shall be left aside.

Grünwald (2004) and estimate a TVNIR for the Czech Republic in a parsimonious multivariate unobserved-components-model (UCM). The estimation is thus based merely on the statistical properties without imposing any economic theory. This method which goes back to Harvey (1989), who aims at decomposing the time-series in a trend component (μ_t), a cyclical component (ϕ_t) and an irregular component (u_t). The vector $z_t = (r_t, y_t, \pi_t)$ consisting of real interest rate r_t, output y_t and inflation π_t can thus be written as

$$z_t = \mu_t + \phi_t + u_t; u_t \backsim IID\left(0, \Sigma_u\right) \tag{3.2}$$

where the multivariate trend component μ_t follows a random walk with drift κ_t which follows a random walk itself:

$$\mu_t = \mu_{t-1} + \kappa_{t-1} + \tau_t; \tau_t \backsim IID\left(0, \Sigma_\tau\right) \tag{3.3}$$

$$\kappa_t = \kappa_{t-1} + \psi_t; \psi_t \backsim IID\left(0, \Sigma_\psi\right) \tag{3.4}$$

In addition, the error terms τ_t, ψ_t and u_t are assumed to be mutually uncorrelated. The cyclical component is defined as a sine-cosine wave with time-evolving parameters[21]

$$\begin{pmatrix} \phi_t \\ \phi_t^* \end{pmatrix} = \rho \left[\begin{pmatrix} \cos\lambda & \sin\lambda \\ -\sin\lambda & \cos\lambda \end{pmatrix} \otimes I \right] \begin{pmatrix} \phi_{t-1} \\ \phi_{t-1}^* \end{pmatrix} + \begin{pmatrix} \omega_t \\ \omega_t^* \end{pmatrix} \tag{3.5}$$

where $(\omega_t, \omega_t^*) \backsim IID\left(0, I \otimes \Sigma_\omega\right)$ are also assumed to be uncorrelated with other errors, the dampening factor $\rho \in \langle 0, 1\rangle$ and frequency $\lambda \in \langle 0, 2\pi\rangle$ are assumed to be time-invariant and equal across variables in z. The state space model formulation enables the application of the Kalman filter. Via the prediction error decomposition the Kalman filter opens the gate to the maximum likelihood estimation of the unknown parameters (Harvey 1989:100 et seq.). We will opt for filtered or real-time estimates which use information available up to $t-1$ while forming expectations on the unobservable state at time t ($\alpha_t^f = E(\alpha_t \mid \{z_t\}_{t=0}^{t-1})$.[22]

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[21] A univariate (non-stochastic) cycle can be most conveniently expressed as a mixture of sine and cosine waves: $\psi_t = \alpha\cos\lambda t + \beta\sin\lambda t$, $t = 1, ..., T$. The period of the cycle (time taken to go through the complete sequence of values) is given by $2\pi/\lambda$ while functions of α and β determine the amplitude and the phase. As Harvey(1989:39) shows this model can be rewritten ψ_t as the following recursion: $\begin{pmatrix} \psi_t \\ \psi_t^* \end{pmatrix} = \begin{bmatrix} \cos\lambda & \sin\lambda \\ -\sin\lambda & \cos\lambda \end{bmatrix} \begin{bmatrix} \psi_{t-1} \\ \psi_{t-1}^* \end{bmatrix}$ with $\psi_0 = \alpha$ and $\psi_0^* = \beta$.

[22] In contrast, smoothed estimates would exploit information contained in the entire sample ($\alpha_t^s = E(\alpha_t \mid \{z_t\}_{t=0}^{T})$.

In line with the literature we use ex-ante real interest rates defined as the prevailing three-month money market rate in t, available at Eurostat, minus inflation between $t-1$ and t. We used the longest available monthly inflation series on the CPI which can be retrieved from the OECD and our estimation sample thus ranged from January 1993 to December 2006. However, as a robustness check we alternatively employed a much shorter series (starting in January 1997) on CPI excluding food and energy and the HICP obtained from Eurostat without significant changes in the results. Output y we approximated by the logged seasonally adjusted industrial production which is available on a monthly basis from the OECD.

In order to estimate the model initial values for cycle determining parameters have to be specified. As noted above though the parameters λ and ρ are assumed to be identical for all variables and they should be set so that the cyclical component of the model corresponds to the business cycle. We tested λ-values that imply business cycle length ranging between two and six years and ρ-values between 0.7 and 0.99. unfortunately, for the Czech Republic we cannot confirm the finding made by Crespo-Cuaresma et al. (2004) for the EU that the estimated parameters are robust to the choice of the starting values of λ and ρ in a multivariate model. This might be due to not really synchronized cycles of the three variables over the course of the transition period and/or due to the substantially shorter time series which might be a problem as the simultaneous estimation reduces the degrees of freedom. However, estimation results are robust to the initial specification of the parameter values if a univariate approach is employed. Having to choose between these two second-best possibilities we decided to opt for the univariate, but robust approach. Moreover, estimation results improved if the trend component was assumed to be smooth with the variance of the error term τ_t equal to 0.

Figure 3.7 and Table 3.3 in the Appendix display the results of the estimation of the univariate unobserved components model for the Czech Republic. While the estimated dampening factor ρ is 0.841 the frequency parameter λ amounts to 0.32 implying a cycle length of about 19 months. The Durbin-Watson test statistic suggests that a first-order autocorrelation of the residuals should be no major issue. For comparison, it should be noted that for industrial production the results of the univariate estimation look significantly different. Whereas ρ is 0.941, λ is close to 0.15 which corresponds to a cycle length of nearly 3.5 years. These results seem to match much more closely the findings obtained by Crespo-Cuaresma et al. (2004) in a multivariate setting for the EMU. The discrepancy between the statistic features of the two series in our estimations might thus explain why they are difficult to squeeze into a multivariate model.
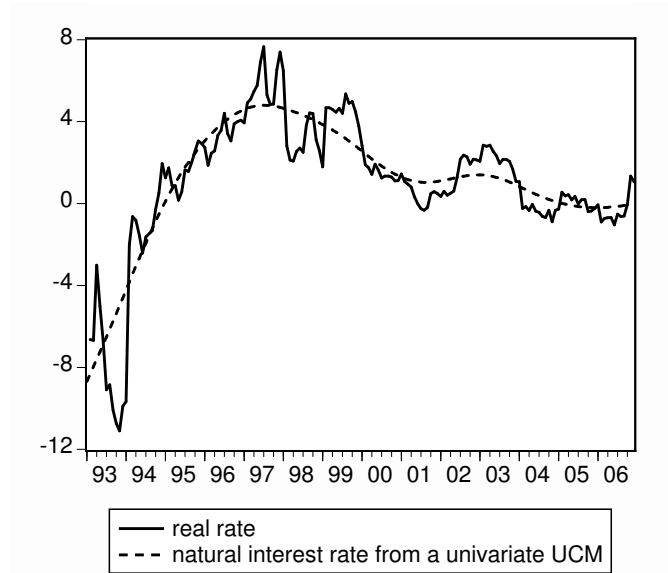
Figure 3.7: Real interest rate in the Czech Republic and the time-varying natural interest rate

At the beginning of this subsection we raised two questions. The first one - how much does the real interest rate in the Czech Republic deviate from the one in the EMU? - is addressed in Figure 3.8. The chart displays the difference between the actual real interest rate in the Czech Republic and in the EMU (the most volatile line) and two versions of the differential between the natural rate of interest (NRI) in both currency areas (the two smooth lines). While the NRI for the Czech Republic was calculated in a univarite UCM for reasons explained above, for the EMU we calculated the NRI both in a univariate UCM (blue/solid line) as well as in a multivariate UCM (red/dashed line) since for the EMU a multivariate version did work quite well and we came very close to the results obtained by Crespo-Cuaresma et al. (2004). Eyeballing the NRI differential it can be seen that a multivariate model tends to yield a smoothed cyclical component (the difference between the actual series and the trend) with slightly smaller amplitude. The figure suggests that since 2001 both measures of the NRI differential have stayed within the grey interval $[-1, 1]$. The maximum deviation of the actual Czech real interest rate from the EMU reached some 2.5 percentage points on either side since 2001.
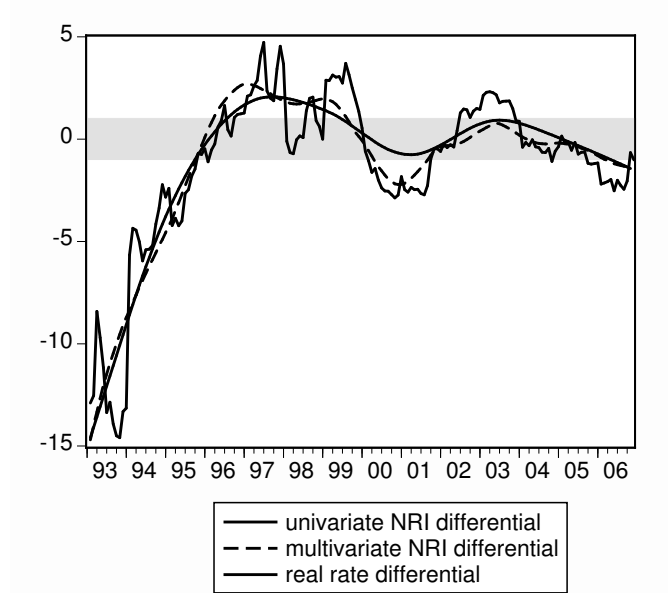
Figure 3.8: Real rate and NRI differentials between the Czech Republic and the EMU

What has been said so far is subject to one major caveat. The real interest rate series for the Czech Republic which was used in our calculations necessarily includes risk premia that reflect the markets' perception of risk, particularly perils related to exchange rate changes and possibly also to the fiscal position of the country. These risk premia, however, will disappear after the euro adoption. Therefore, if we want to get some feeling for how much the Czech real interest rate (actual or the NRI) will deviate from its EMU counterpart under a common currency for the sake of a fair analysis we first have to purge the real rate series from the risk premium. Crespo-Cuaresma et al. (2004) develop a method to eliminate the risk premium from the nominal money market rates from which a synthetic, risk-adjusted real rate series can easily be constructed.

In the model, the time-varying risk premium is defined as the residual part of the nominal interest rate spread with the EMU short-term interest rate that cannot be explained by differentials in inflation expectations and business cycle disharmony. Hence, the nominal interest rate spread between the Czech Republic and the EMU (s_t^{cr}) is regressed on the output gap differential $(g_t^{cr} - g_t^{emu})$ and the inflation expectation differential between the two regions $\left[E\left(\pi_{t+12}^{cr} \mid \{\pi_k^{cr}\}_{k=1}^{t} \right) - E\left(\pi_{t+12}^{emu} \mid \{\pi_k^{emu}\}_{k=1}^{t} \right) \right]$:

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$$s_t^{cr} = \beta_0 + \beta_1(g_t^{cr} - g_t^{emu}) \tag{3.6}$$
$$+\beta_2 \left[E\left(\pi_{t+12}^{cr} \mid \{\pi_k^{cr}\}_{k=1}^{t}\right) - E\left(\pi_{t+12}^{emu} \mid \{\pi_k^{emu}\}_{k=1}^{t}\right)\right] + \gamma_t$$

While we use as a measure for the output gap the cyclical component of industrial production obtained in the previous estimation of a univariate UCM, the inflation expectations are computed as 12 months ahead forecasts in an autoregressive model whose length[23] is reestimated at each point in time t so as to minimize the Akaike information criterion. In order to estimate the inflation expectations we thus had to split the sample into two parts and it turned out that the best performance was achieved if the first forecasts were computed for January 1998. Hence, the part of the sample between 1:1993 and 1:1997 was used to compute the inflation forecasts while for the actual estimation of equation (3.6) the second half of the sample between 1:1998 and 12:2006 was employed. Due to possible correlations between the error term γ_t and the regressors[24] the estimation was carried out by means of a two-stage least squares estimation in which lags ranging between six and twelve months of the output gap and inflation expectation differentials were used as instruments. Table 3.4 in the Appendix shows the results of this estimation.

According to the definition spelled out above the risk premium incorporated in the Czech nominal interest rate corresponds to the sum of the estimated constant $\widehat{\beta}_0$ and the random shock $\widehat{\gamma}_t$. Hence, the nominal interest rate purged from the risk premium corresponds to

$$i_t^{adj} = i_t^{cr} - \widehat{\beta}_0 - \widehat{\gamma}_t \tag{3.7}$$

In Figure 3.9 we plot the adjusted real rate against the original series and for the sake of comparison also the real rate of the EMU. It might surprise that the adjusted real rate has been persistently above the actual real rate since mid 1999. In other words, the figure suggests that there has been a negative risk premium in the Czech Republic which might reflect the trend appreciation of the koruna vis-à-vis the euro.[25] In analogy to Figure 3.8 we show in Figure 3.10 the difference between adjusted real rate in the Czech Republic and the actual real rate in the EMU as well

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[23]At each t the maximum lag length is 12.

[24]Inflation expectations, for instance, might be driven by the same factors as the risk premium.

[25]Between June 1999 and December 2006 the koruna appreciated in nominal terms against the euro by more than 25%.

as the two versions of the NRI-differentials. Whereas the adjusted real rate deviated at most by slightly more than 1 percentage point from the real rate in the EMU since the beginning of 2004, the two NRI-differentials have declined rapidly over the past couple of years and are currently close to 0. It also strikes that while the actual real rate was below the EMU level in the last three years or so of the sample, the risk premium adjusted real rate has been well above the EMU rate since mid 2006.
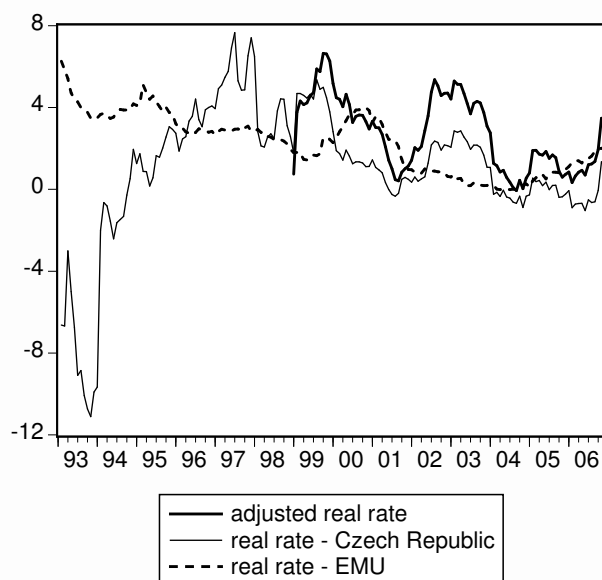
Figure 3.9: Adjusted and originial real rate in the Czech Republic versus real rate in the EMU

Hence, the evidence provided so far suggests that despite a significant convergence of the Czech real and natural interest rates towards the EMU-level (particularly if purged from the risk premium) there remains some little gap. Moreover, it is not quite unambiguous which sign this gap would have if the euro was adopted now. Nevertheless, in the next step we shall address the second question raised at the beginning of this subsection. Namely, would it actually matter for the inflation path in the Czech Republic if the EMU NRI deviated from its Czech counterpart? In other words, we are wondering how sensitive the inflation in the Czech Republic is to deviations from the natural rate. In order to get some feeling for how relevant possibly desynchronized natural rates between the EMU and the Czech Republic might be after the euro adoption we will approximate this future elasticity by the current sensitivity of the inflation to the deviation of the Czech real rate from the NRI. We

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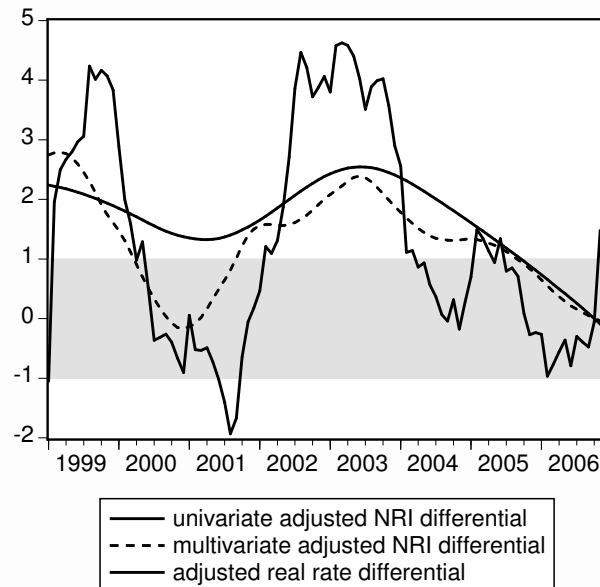
Figure 3.10: Adjusted real rate and NRI differentials between the Czech Republic and the EMU

will use both the original as well as the adjusted series although Crespo-Cuaresma et al. (2004) conclude that the adjusted series serves better as an indicator of inflation development. Notwithstanding, it should be emphasized at this point that the natural rate of interest is an unobserved variable and the estimation thereof depends essentially on the economic specification and the econometric method. A whole raft of literature documents that policy rules based on such unobserved variables, particularly if they are estimated in real time, are subject to a substantial level of uncertainty (see Crespo-Cuaresma et al. (2004) and other references therein). For this reason we will use in our regressions as an alternative the NRI estimated by Horváth (2007)[26].

3.3.2 Temporary factors related to the euro changeover

As has been shown above, although the overall impact was moderate, in some countries and some segments prices did increase substantially after the changeover. In

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[26]Horváth (2007), to which we are grateful for sharing his results with us, estimates nominal neutral rates which we will transform to real rates as described above.

most cases prices returned to the pre-changeover level and only in a few industries, particularly services such as gastronomy or cinemas, the effect was not transitory. However, there were big differences among the countries. Eife (2005) who focuses on the price setting behaviour around the changeover in Austria and Germany finds that unlike in the latter in Austria both the transitory and the persistent increases were practically absent (despite the substantial increase in perceived inflation) barring a few isolated exceptions such as hairdressing or language courses.

The most evident costs of a currency changeover are the so called menu costs, i.e. costs of replacing the price tags and menus. The higher the menu costs are the bigger the incentive for the firms to carry out intended price adjustments at one stroke with the changeover. Hence, price modifications in the weeks or months preceding the euro adoption will be postponed, those planned for the period after the changeover will be brought forward . Menu costs are the usual explanation the literature provides for the discrete price jumps that could be observed particularly in the service sector.[27] However, Eife (2005) concludes that price spikes in some sectors can barely be explained by menu costs for two reasons. First, the increases were too high so that they could be attributed solely to menu costs and second, if the menu cost argument washed than one should observe significantly less frequent price changes before and after the changeover which is not the case for most analysed goods and services categories. In contrast, according to the same study two factors appear to have been more important for transitory price changes. On the one hand, the so called 'initial confusion' argument seems to have been at play which says that the public is puzzled by unfamiliar coins and notes. In this situation similar to the initial confusion one experiences in a foreign country firms might be attempted to take advantage of this sort of a temporary increase of their market power.[28] On the other hand, however, there was a counteracting effect due to the higher sensitivity and awareness of the public and the media to price jumps. It thus turns out that big corporations which were certainly under a much harsher media pressure were much more hesitant to raise their prices and some of them, such as fast food chains in Austria, even welcomed the opportunity to grab the headlines for reducing the prices. Moreover, Eife (2005) finds that those services for which permanent price hikes could be observed are also typically sold by small, only locally active firms

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[27]See for example Hobijn et al. (2004) whose sticky price model generates a blip in inflation of the same magnitude observed in the data which can be explained by menu costs.

[28]This view is supported also by Ehrmann (2006) who found that the denomination of prices in a new currency had increased the information-processing requirements for consumers by more than for sellers. In addition, this wedge grows with the complexity of the currency conversion rates.

and are standardized in the sense that irrespective of the provider the consumer has quite a precise expectation of what she gets. Eife (2005) with reference to Tirole (1988) thus argues that these are emblematic characteristics of markets for which collusive behavior is typical and that the price increases were to a great extent driven by a collective shift to a new, higher equilibrium. Yet based on data for Austria that managed to create an environment in which both transitory and permanent price increases were rare Eife (2005) concludes some normative recommendations that should help render the changeover a mere nominal conversion event: To address the initial confusion a sufficiently long period of dual pricing is necessary.[29] In the light of the Slovenian experience, however, not too heavy a reliance should be placed on dual price display as an anti-inflation measure in itself, and a substantial emphasis should be put also on fair-pricing agreements between the government and the private sector (see Deloitte Consulting 2007). Moreover, as the findings made by Ehrmann (2006) suggest, the less complex the conversion rate the smaller the confusion of the consumers and thus the smaller the increase of firms' market power. To reduce the incentive for collective price adjustments within a short period of time and to make the transition more gradual the replacement should extend over several months.[30] In addition, to enhance the countervailing effect price observatories and hotlines operated e.g. by the statistical office or the central bank but also by private institutions could be set up or stimulated. Although the major threat for 'sinners' would probably be the attention by the media an enforceable legal framework would certainly also help to curb the incentive to abuse the changeover.

As we have seen in Figure 3.1 irrespective of the actual inflation path the perceived inflation rose in all countries following the introduction of euro coins and notes. Aucremanne et al. 2007 confirm formally the hypothesis that this perception break was significantly induced by the euro changeover. In contrast to the widespread opinion in the literature, they do not find any support for the finding that perceptions are systematically biased by frequently purchased goods and services (see e.g. Del Giovane and Sabbatini 2006) or that the perception gap was more pronounced for consumers with some specific socio-economic characteristics (see e.g. Fluch and Stix 2007 or Lindén 2006). Hence, whereas the findings on these issues might not be

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[29]In Austria, after the expiration of the compulsory period between October 2001 and February 2002 in which prices had to be displayed in Schilling and Euro dual pricing was still allowed after that. In Germany, in contrast, dual pricing was possible, not obligatory, prior to January 2002 but was no longer allowed after the 2-months transition period.

[30]While it was more than 12 months in Austria, in Germany the replacement period lasted only 2 months.

unambiguous, Traut-Mattausch et al 2004 provide strong experimental evidence suggesting that people's judgments on price trends are biased towards price increases if prices are denominated in the new and unfamiliar currency, the euro. These results are thus probably another manifestation of an anomaly that Samuelson and Zeckhauser (1988) called the 'status quo bias' - people's preference for the current state and their implied mistrust of new states. This phenomenon like other anomalies is a materialization of the so called 'loss aversion' - an asymmetry first documented by Kahneman and Tversky (1984) - which states that the disutility of losing an object is greater than the utility associated with acquiring it (see Kahneman et al. 1991). In the spirit of the status quo bias the disadvantages of a change thus loom higher in people's minds than the advantages of it. The size of this misalignment depends certainly crucially on the expectations people form on the advantages and disadvantages of the change. The expectation formation, in turn, is undoubtedly essentially influenced by the media, the institutions and the authorities people trust in such as high-rank politicians, the central bank or commercial banks.[31],[32] In the Czech Republic a great part of the discussion on costs and benefits of the euro centres around the inflationary impact of the euro and as has been already alluded to a vast majority of the (inter alia prominent) discussants supports the hypothesis that the euro will lead to a marked inflation acceleration. So it is probably little surprising that 80% of the Czechs, 6 percentage points above the average in the New Member States, believe that the euro will increase prices when it is first introduced and two thirds of the respondents are afraid of abuses and cheating on prices during the changeover (see Eurobarometer 2007). A balanced and detailed information campaign on the impacts of the euro changeover as well as a set of orchestrated measures proposed in the previous paragraph would certainly have an impact on these figures. Most likely they would tame those fears and accelerated inflation expectations because as Fluch and Stix (2007) and, in a slightly different phrasing, also Lindén (2006) document uninformed respondents tend to have significantly higher both inflation perception and expectation values than the informed ones. Obviously, it is impossible to control for the effects of the current or future information campaign. However, we will test in our regression how inflation expectations impact actual inflation rates. If this impact turned out to be rather strong as e.g. Paloviita and Virén (2005) show for the euro area in a VAR using inflation forecasts from the OECD then the numbers on expected inflationary effects of the euro should probably be addressed with more

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[31]Del Giovane and Sabbatini (2006) or Boeri (2004), for instance, provide evidence suggesting that there is a positive relationship between inflation perception and the attention the press pays to the phenomenon.

[32]For the preference ranking of trusted distributors of information see Eurobarometer (2007).

vigour well in advance of the euro introduction as they could end up in self-fulfilling equilibria with higher inflation rates. The data on qualitative inflation expectations are collected from the Joint Harmonised EU Programme of Business and Consumer Surveys regularly conducted by the European Commission's Directorate-General for Economic and Financial Affairs. Similarly to the perceived inflation the series on inflation expectations is a balance statistic calculated as the difference between a weighted proportion of respondents stating that prices will increase and those stating that prices will stay the same or fall.[33]

3.4 Econometric evidence

3.4.1 Method and data description

In this section we now proceed to test econometrically the relative importance of the factors collected so far along with other cyclical and external variables the choice of which we base primarily on the results obtained by Égert (2007a). Given the availability of the data and the auxiliary calculations described above the adjusted sample ranges from January 1998 to December 2006. All variables as used in the model are stationary according to the ADF test on conventional significance levels and if not available on monthly basis they were linearly interpolated.

In addition to the variables defined so far we included in the regression also the output gap for which we use three different measures: the cyclical component of industrial production relative to its trend resulting from the estimation of a unilateral unobserved components model described above and two output gap measures obtained from the European Commission.[34] The other cyclical variables are the annual rate of change of unit labor costs and the y-o-y change of central government expenditures relative to GDP. The set of structural variables contains three factors that have been spelled out above - the productivity differential, the annual change of regulated prices and the GDP per capita.[35] Concerning external factors apart from the

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[33]There are six possible answers to the question: "By comparison with the past 12 months, how do you expect that consumer prices will develop in the next 12 months? They will..." A(1) "increase more rapidly"; A(2) "increase at the same rate"; A(3) "increase at a slower rate"; A(4) "stay about the same"; A(5) "fall" and A(6) "do not know". The balance statistic is thus computed as: $B_{it} = A_{it}(1) + 0.5A_{it}(2) - 0.5A_{it}(4) - A_{it}(5)$ (See: http://ec.europa.eu/economy_finance/indicators/businessandconsumersurveys_en.htm)

[34]The measures are defined as the relative gap between actual and, respectively, trend or potential gross domestic product at 2000 market prices.

[35]Since GDP per capita appears to be growing almost exponentially we defined this variable as

two already mentioned alternatives which are supposed to capture the exchange rate pass-through we include also the y-o-y change of the crude oil price. Next we constructed two dummy variables for VAT increases and decreases, respectively. They equal 1 in those months in which the VAT rate was raised/lowered and also in those months in which a product or service group was moved from the preferential rate to the basic rate or vice versa. If two counteracting effects occurred simultaneously then we set that dummy equal to one for which the consumption basket weight of the goods and services in question was bigger. In all other months the dummies equal zero.[36] Finally, we include inflation expectations and three alternative measures for the deviation of the real interest rate from the natural interest rate on the grounds developed above. A neat list of all variables can be seen in Table 3.5 in the Appendix.

To account also for the impact of past inflation values on future inflation we will employ an autoregressive distributed lag model of the form

$$\pi_t = \mu + \sum_{i=1}^{p} \gamma_i \pi_{t-i} + \sum_{j=0}^{r} \beta_j x_{t-j} + \varepsilon_t \tag{3.8}$$

where π_t stands for inflation, x_{t-j} for a vector of regressors lagged by j. The error term ε_t is assumed to be serially uncorrelated and homoskedastic. Given the limited length of the sample we restrict the set of lags to $p \in \{1, 6, 12\}$ and $r \in \{0, 1, 6, 12\}$ which should still allow us to capture the contemporaneous as well as the short-, mid- and long-term effect of each variable. The estimations are carried out using OLS. However, to determine the lag length for each variable we first tried two alternative model selection techniques. The first one was a general-to-specific strategy based on hypothesis testing. We thus started out from the biggest model including all variables and all possible lags. Then we proceeded iteratively and in each round we eliminated the least significant regressor (=variable-lag-combination), the one with the highest p-value, until we were left with a parsimonious model containing only regressors significant on the five percent level. The second strategy selects iteratively from all possible models of which there are $2^{\overline{K}}$, where $\overline{K}$ stands for the number of initial regressors, the one with the lowest Schwarz Information Criterion (SIC). Since the maximum number of initial variables we tested was 13, the implied maximum number of regressors was 51 (= 4*13-1). If the computation of one iteration step took only, say, 0,001 seconds than it would take more than 71 000 years

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the month-on-month change of the y-o-y growth of GDP per capita which can be interpreted as the short term acceleration of the GDP/capita growth rate.

[36]I would like to thank my mother, a tax adviser, for helping me out with the tedious work through the tax code.

to go through all models. Therefore we had to reduce the intractable model space which we did by means of the so called Markov Chain Monte Carlo Model Composite (MC^3) method which was described in detail in Chapter 2 of the present dissertation.

As was also briefly alluded to in Chapter 2 there are some problems with such model selection strategies. Under the test based general-to-specific strategy the model space is strongly limited as it starts from the biggest model whose size is then iteratively reduced but it does not take into account possible model alternatives with variables that have been already excluded. Implicitly this procedure thus imposes a rather strong restriction as it presumes to know a limited model space in which the true model has to be included. In fact, however, this is not the case. Given the lack of an unambiguous theoretical framework that would uniquely determine which variables and lags are to be chosen in an equation attempting to explain inflation, we do not know either the true model or a restricted subset of all possible models from which the true model has to be recruited. Hence, we face a substantial level of model uncertainty that has to be taken into account. Although the strategy based on the SIC does in principal consider all perceivable models it did not perform much better. Namely, both strategies differ strongly in the models they choose and they do not converge to one model as one would expect owing to the fact that they are extremely sensitive to the definition of variables and the initial specification. Yang (2004) who compares hypothesis testing and model selection strategies both theoretically and empirically concludes that *"when model selection rules give very different answers, model combining is a better alternative approach for estimation and prediction. With a proper weighting, the large variability of the estimator from model selection can be substantially reduced."*

A way to combine models with a proper weighting is the Bayesian model averaging (BMA) methodology which we already employed and described in detail in Chapter 2 and to which the reader may refer. In a nutshell, the BMA algorithm proposes averaging of the parameter values over all (relevant) alternative models using posterior model probabilities as respective weights. In addition, for each variable we sum the posterior probabilities of those models visited by the MC^3 algorithm that include the respective variable. This sum, the so called posterior inclusion probability, is a measure that captures the relative importance of the different covariates as determinants of inflation. It can be interpreted as the probability that a given variable belongs to the true specification. Since we assign equal priors to all models, $P(M_i) = P(M_j)$ for all M_i, $M_j \in M$, our prior on the inclusion probability

of each variable is 0.5.[37] Moreover, we can also use the inclusion probability as a model selection criterion and combine the Bayesian and frequentist approach. In a Bayesian based 'frequentist check' we thus keep only those regresors whose model averaged parameters have an inclusion probability equal to or greater than the prior benchmark 0.5 and run a new estimation only with these regressors.

3.4.2 Estimation results

The application of the BMA methodology in combination with the MC^3 algorithm produced much more robust results than either of the two other methods. The selection procedure chose under most amendments the model presented in Table 3.1. We report under 'BMA' the parameter averages and inclusion probabilities of those regressors with an inclusion probability equal to or greater than 0.5. Under 'Frequentist check' the reader may find the usual 'frequentist' output from a regression run with those regressors that have been selected using inclusion probability as a model selection criterion.

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[37]There are $2^{\bar{K}-1}$ models including a given variable and $2^{\bar{K}}$ total models, so the prior inclusion probability of a given variable is $2^{\bar{K}-1}/2^{\bar{K}}$=0.5.

Dep. Var.:   inflation
Method: Least Squares
Incl. observations: 60 after adjustments

| | BMA | | Frequentist check | | | |
Variable	$E(\beta_i/Y)$	Inc. Prob.	Coefficient	Std. Error	t-Statistic	Prob.
const.			0.804654	0.147676	5.448780	0.0000
infl.(-1)	0.837428	1.000000	0.882788	0.045621	19.35046	0.0000
infl.(-12)	−0.427722	0.899787	−0.421428	0.081746	−5.155349	0.0000
prod. differential(-6)	−0.022989	0.794936	−0.022467	0.006121	−3.670748	0.0006
(realrate-NRI)(-12)	−0.318557	0.859325	−0.305976	0.077111	−3.967982	0.0002
infl. of reg. prices(0)	0.226622	1.000000	0.219477	0.035395	6.200828	0.0000
infl. of reg. prices(-1)	−0.187172	0.999976	−0.182571	0.036386	−5.017684	0.0000
infl. of reg. prices(-12)	0.041563	0.977844	0.046075	0.021219	2.171376	0.0345

			R^2	0.949488	AIC	0.487067
			adj. R^2	0.942688	SIC	0.766313
			DW	1.958128	Prob(F-stat)	0.000000

Table 3.1: Estimation of the ARDL model

As can be seen, except for the productivity differential the regressors have an interpretable sign and the estimates reveal interesting information on the determinants of inflation in the Czech Republic. While inflation appears highly persistent in the short-run, in the long-run high inflation rates tend to reduce future inflation. The latter is probably a result of the strong disinflation process that took place in the Czech Republic in the sample period. Another or additional interpretation of this result could be that restrictive measures of monetary policy implemented as a reaction to high inflation kick in with a 12-month-lag. Adjustments of regulated prices have a strong and long-lasting impact on overall inflation. Inflation typically rises in the month in which administered prices are raised but falls in the immediately following month. A weak, positive correlation between regulated and overall prices is found also for the 12-month-lag. The long-run effect of the regulated price inflation, which can be in an ARDL model computed as

$$\frac{\sum_{j=1}^{r} \beta_j}{1 - \sum_{i=1}^{p} \gamma_i}$$

where β_j and γ_i refer to the model specification in (3.8) (see Greene 2003), is 0.154 meaning that a 10% increase of regulated prices raises overall inflation in the long-run by 1.5 percentage points. In addition, if today's real interest rate deviates from the natural interest rate upwards by 1 percentage point the inflation in 12 months will be dampened by 0.3 percentage points. These results appear plausible and the only variable we struggle with is the productivity differential supposed to capture the B-S-effect. If the B-S-effect in the Czech Republic was an issue, the parameter should be at least positive as the higher productivity growth in the tradable sector relative to the non-tradable sector the higher should be the pressure on wages in the latter and thus on the overall CPI. This seems not to be the case in the Czech Republic and our results thus replicate the findings in the existing literature.

This resulting parsimonious model was rather robust to alternative definitions of the variables and also to the initial specification of the model.[38] The results changed slightly though, if we employed the natural rate of interest obtained by Horváth (2007) as can be seen in Table 3.2. Unlike in the previous model this time the BMA procedure did not select any of the two, not easily interpretable regressors - the productivity differential and the twelfth lag of the regulated prices inflation. Other than that, the remaining variables showed up at the same lags and the coefficients

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[38]We did not experiment only with various alternative definitions of the variables but also with the number of variables initially included.

have the same signs and similar magnitudes as before. Only the long-run effect of the inflation of regulated prices amounts now to 0.058. The inclusion probability of all other factors did not surpass the prior threshold to enter the final equation. Hence, according to our estimates neither fiscal policy (including VAT changes), nor cyclical factors, nor the richer growing households seem to be accelerating the inflation. Moreover, external factors also appear to matter very little. In particular, in line with the existing literature also in our results the exchange rate pass-through turns out to be rather weak so that the lack of the appreciating currency after the euro adoption should have no dramatic impact on inflation. In contrast, our results provide very little support for the exchange rate expectation being a relevant inflation determinant - a result that seems to contradict conclusions drawn by other studies on the euro area. In further research may be alternative measures of inflation expectations should be tested to shed more light on this issue.[39]

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[39]The modelling system of the Czech National Bank (CNB) assumes that inflation expectations are formed predominantly by the latest inflation developments and to a lesser extent by the forecast for future inflation in a ratio of 90:10 (see Box in CNB' Inflation Report, January 2005). Hence, it could be argued that in a way inflation expectations are captured by the lagged inflation variable.

Dep. Var.:   inflation
Method: Least Squares
Incl. observations: 56 after adjustments

| | BMA | | Frequentist check | | | |
Variable	**E(β_i/Y)**	**Inc. Prob.**	**Coefficient**	**Std. Error**	**t-Statistic**	**Prob.**
const.			1.400264	0.259484	5.396336	0.0000
infl.(-1)	0.730016	0.999697	0.736449	0.046076	15.98344	0.0000
infl.(-12)	−0.259026	0.999560	−0.252407	0.038709	−6.520565	0.0000
(realrate-NRI)(-12) (Horváth 2007)	−0.332356	0.999995	−0.320396	0.084548	−3.789534	0.0004
infl. of reg. prices(0)	0.232367	1.000000	0.231591	0.034529	6.707038	0.0000
infl. of reg. prices(-1)	−0.200496	0.999557	−0.201613	0.035031	−5.755355	0.0000
			R^2	0.945865	AIC	0.396836
			adj. R^2	0.940452	SIC	0.613838
			DW	2.309893	Prob(F-stat)	0.000000

Table 3.2: Estimation of the ARDL model employing NRI obtained by Horváth (2007)

3.5   Summary and conclusions

In the present chapter we analyse relevant macro and microeconomic forces driving inflation in the Czech Republic with a particular focus on the question how these inflation channels are likely to change in the wake of the euro adoption. After a qualitative discussion we employ an ARDL model. Since standard model selection criteria disagree, following the recommendations made in the theoretical literature for such cases we combine all models under an appropriate weighting scheme by using the Bayesian Model Averaging algorithm while we reduce the intractably large model space by the Markov Chain Monte Carlo Model Composite algorithm. Our paper is innovative in one more aspect since we estimate as a side-product the time-varying natural rate of interest by means of an unobserved components model both from the actual series as well as from a series purged from the risk premia.

We do not find much evidence underpinning the line of argument usually put forward according to which inflation will rise because of the absence of the appreciating koruna. In line with the existing literature also in our results the exchange rate pass through appears rather weak which implies a faster price level convergence in case of a currency appreciation but a negligible dampening effect on the CPI inflation. In other words, the costs attributable to the lack of koruna appreciation after euro adoption are likely to be rather low. Not surprisingly, a low inflation environment is very supportive for future inflation rates. Inflation also seems quite sensitive to deviations from the natural interest rate. Hence, an essential prerequisite for a smooth inflation development after the euro adoption is a harmonization of the inflation and the business cycles between the Czech Republic and the euro area, and thus consequently of the natural interest rates.

In addition, empirical results as well as anecdotal evidence from some current euro area members advise that for the sake of a sustainably stable inflation the fulfilment of the Maastricht inflation criterion should not be enforced by withholding price deregulations, reducing indirect taxes or by other non-standard policy measures such as gentlemen's agreements with the private sector or labor unions. As regards the inflationary effect of the changeover although the price increases that occurred were transitory in most cases, theory and past experience suggest that there are better and worse ways to handle the undoubtedly luring risks of abuses for price hikes. If possible, an easy conversion rate, a sufficiently long dual pricing and replacement period as well as fair-pricing agreements with the private sector proved helpful. Negative publicity in the media enhanced by the set-up of price observatories and

hotlines also reduce incentives to take advantage of the event. In contrast, inflation perceptions are a very different story and given the well documented loss aversion in people's minds it is a challenging task to tackle the rife and deeply embedded opinion that the euro raises prices. Probably not much more can be done but a balanced, transparent and fair information campaign on the inflation aspects of the euro.

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3.6   Appendix

Parameter estimates	
ρ	0.841
λ	0.326
Σ_u	0.829332
Σ_ω	0.035163
Residual analysis	
Std. error	0.96426
DW	1.841

Table 3.3: Univariate unobserved components model

Dep. Var.:	Interest rate spread btw. CR and EMu			
Method:	Two-Stage Least Squares			
Sample (adj.)	1999/M01 2006/M12			
Incl. observations:	96 after adjustments			
Instrument list	Output gap(-6 to -12), Inf. expectation diff.(-6 to -12)			

Variable	Coefficient	Std. Error	t-Statistic	Prob.
const.	-1.549415	0.279479	-5.543945	0.0000
Output gap	-26.84044	10.15148	-2.643994	0.0096
Inf. expectation diff.	0.463972	0.060047	7.726780	0.0000

R^2	0.658006
adj. R^2	0.650651
DW	0.079463

Table 3.4: Risk premium estimation

Variable name	Definition
GDP/capita at PPP	m-o-m change of y-o-y rate
Productivity diff.	$(1-\alpha)\left(\Delta prod^T - \Delta prod^{NT}\right)$, $\alpha = 0.25$;productivity $= \frac{grossvalueadded}{totalhoursworked}$ (see text for details)
Inf. of regulated services	y-o-y change, see text for details
Openness indicator	$e\frac{IM^{EU15}}{GDP}$ or, alternatively, y-o-y change of e; Note: $e = \frac{EUR}{CZK}$; $IM^{EU15}=$ import from EU 15
Deviation from the NRI	real rate - NRI; three alternatives for NRI : i) original ii) adjusted iii) Horváth (2007)
Output gap	$\frac{cyclicalcomponentfromUCM}{trendcomponentfromUCM}$ or, alternatively, two output gap measures from EC (see text)
ULC	y-o-y change
Crude oil price	y-o-y change
VAT increase	dummy; see text for details
VAT decrease	dummy; see text for details
Inflation expectations	EC; see text for details
Government expenditure	$\frac{centralgov.expend.}{GDP}$ (y-o-y change)

Table 3.5: Basic variables: Definitions

Appendix A

Deutsche Zusammenfassung der vorliegenden Dissertationsarbeit

Der Wechselkurs ist einer der Schlüsselpreise in jeder Volkswirtschaft. Er übt einen wichtigen Einfluss auf viele wichtige makroökonomische Variablen aus, indem er unmittelbar das Verhalten und den Entscheidungsprozess der Wirtschaftsagenten wie der Haushalte, Unternehmen, Regierungen und Finanzinstitutionen beeinflusst. Auf der einen Seite kann der Wechselkurs als ein bedeutender Anker und Stabilisator fungieren. Auf der anderen Seite birgt jedes Wechselkursregime vor allem vor dem Hintergrund der zunehmend globalisierten Kapitalflüsse seine charakteristischen Risiken. Die vorliegende Dissertation setzt sich in drei Kapiteln mit zwei spezifischen, wechselkursregimeverwandten Risikoaspekten auseinander - der Gefahr einer Währungskrise und der Möglichkeit einer durch die Euroeinführung hervorgerufenen Inflationsbeschleunigung, die in den meisten Eurokandidatenländern heftig diskutiert wird.

Im ersten Kapitel wird ein Model präsentiert, mit dessen Hilfe man die Zinsstruktur im In- und Ausland zur Schätzung der relativen, von den Marktteilnehmern wahrgenommenen Währungskrisenwahrscheinlichkeit heranziehen kann. Der entwickelte Frühwarnindikator signalisiert somit die Wahrscheinlichkeit einer Krise in den nächsten sieben Tagen relativ zu dem Rest des Monats. Anschließend wird gezeigt, dass ein auf diese Weise konstruierter Indikator mehrere signifikante Signale vor den Währungskrisen in Tschechien 1997 und in Russland ein Jahr später gesendet hätte.

Auch das zweite Kapitel beschäftigt sich mit dem Thema Frühwarnmechanismen für

Währungskrisen. Es wird aufgezeigt, wie man in einem Logit-basierten Frühwarnmodel im Unterschied zu der bisherigen Literatur explizit Modelunsicherheit berücksichtigen kann und wie dadurch die Ergebnisse beeinflusst werden. Mittels Bayesianischer Techniken wird somit sowohl für statische als auch dynamische Modelle die Robustheit der, in der Literatur vorgeschlagenen, Erklärungsvariablen getestet. Wenn man versucht, den genauen Krisenzeitpunkt zu erklären, stellt Modelunsicherheit die üblichen fundamentalen Krisendeterminanten in Frage. Die Resultate kommen denen in der existierenden Literatur nä her, wenn man nicht Krisenzeitpunkte, sondern Krisenzeitfenster erklärt. Bei einer solchen Definition der abhängigen Variable scheinen insbesondere die reale Überbewertung sowie Finanzmarktindikatoren einen starken Effekt auf die Wahrscheinlichkeit einer Währungskrise auszuüben.

Das letzte Kapitel widmet sich dagegen dem Thema einer potentiellen, durch die Einführung der Gemeinschaftswährung hervorgerufenen Inflationsbeschleunigung. Am Fallbeispiel der Tschechischen Republik werden somit inflationsrelevante makro- und mikroökonomische Faktoren analysiert mit einem Fokus auf die Frage, wie sich deren Einfluss auf die Inflation im Zuge der Euroeinführung verändern wird. Methodologisch bedient man sich dazu eines mit Bayesianischen Techniken kombinierten 'autoregressive distributed lag' (ARDL) Models. Als Nebenprodukt wird zum ersten Mal für die Tschechische Republik ein zeitvariabler, um die Risikoprämie bereinigter natürlicher Zinssatz geschätzt. Die Ergebnisse deuten an, dass Kosten, die durch den Verlust einer nominellen Trendaufwertung der Koruna gegenüber dem Euro entstehen, eher moderat sein dürften. Dagegen spielen niedrige Inflationsraten und eine möglichst gute Abstimmung der Wirtschaftszyklen in Tschechien und der Eurozone eine Schlüsselrolle für eine glatte Inflationsentwicklung nach der Euroübernahme. Nichtsdestotrotz, das Erfüllen der Maastrichter Inflationskriteriums sollte nicht durch außergewöhnliche politische Maßnahmen erzwungen werden. Der potentielle Inflationäre Effekt des physischen Bargeldtausches kann zwar wahrscheinlich nicht komplett eliminiert, aber wenigstens deutlich reduziert werden.

Appendix B

Curriculum Vitae - Tomáš Slačík

CURRICULUM VITAE - TOMÁŠ SLAČÍK

Personal

Date and place of birth	June 4, 1978 in Liberec, Czech Republic, Europe
Nationality	Czech
Civil status	Single, no children
Contact	Oesterreichische Nationalbank Foreign Research Division Vienna, 1090 Austria, Europe Email**:** tomas.slacik@oenb.at Tel: 0043 676 9759378

Education

University of Vienna Vienna, Austria	summer term 2007 - PhD-student, Economics Department
University of California (UCLA) Los Angeles, USA	2004-2005 PhD-student, Economics Department, 2nd year coursework
University of Mannheim Mannheim, Germany	since 10/2003 PhD-student, Center for Doctoral Studies in Economics and Management; 1st year coursework
Universal English College Sydney, Australia	10/2002-12/2002 *English for Academic Purposes (EAP)*
University of Leipzig Leipzig, Germany	1999-2002 Masters Degree in Economics (Diplom) Majors: Money and Currency, International Economic Relations, Marketing Grade: **1.3** (equivalent to High Distinction)
University of Bonn Bonn, Germany	1997-1999 Undergraduate studies in Economics (Vordiplom) Grade: **1.4** (equivalent to High Distinction)
Bilingual high school Liberec, Czech Republic	1992-1997 German and Czech graduation diploma Grade: **1.3** (equivalent to High Distinction)
	Since 2001 Various economic and econometric courses and workshops at the University of New South Wales in Sydney, Institute for Advanced Studies and Oesterreichische Nationalbank in Vienna, Dresdner Bank Frankfurt and Foundation for Economic Education in New York, USA.

Work experience

Oesterreichsiche Nationalbank Vienna, Austria	12/2005 - Foreign Research Division Country and financial market monitoring (Czech Republic, Serbia), research on financial crises, exchange rates, inflation
Deutsche Bank Frankfurt, Germany	10/2005-11/2005 Risk Management Department; Internship Credit scoring by means of discrete choice models, preparation of data
Bolsa de Comercio Cordoba, Argentina	7/2003-9/2003 Centro de Investigaciones Economicas; Research Visit Cost-benefit analysis of a potential monetary union in MERCOSUR
Institute for Economic Research Halle, Germany	4/2001-3/2002 Central- and Eastern Europe Department Diploma Thesis on *Unilateral euroisation as an alternative strategy for the central and eastern European countries during the EU-approach* (in German) Grade: **1.0** (the best grade possible)
Skoda Auto, VW Group Mlada Boleslav, Czech Republic	9/2000-10/2000 and 9/1999-10/1999 Department for Controlling / Client Services; Internship

Teaching experience

University of Applied Sciences Eisenstadt, Austria	Summer term 2007 Undergraduate course on economic and political transition in the Czech Republic
University of Mannheim Mannheim, Germany	Summer term 2004 Assistant lecturer in undergraduate microeconomics courses
University of Leipzig Leipzig, Germany	1/2000-3/2001 and summer term 2002 Assistant lecturer in undergraduate microeconomics courses, proofreading of textbooks and papers by Prof. Wiese
Univers Language School Liberec, Czech Republic	8/1996 – 8/1997 and 4/2002-6/2002 Founder, English and German Teacher

Other qualifications

Language skills
Czech (Slovak) (native), German (Graduation Diploma), English (EAP-Diploma Sydney; TOEFL 283), Spanish (fluent), Russian (basic)

Software skills
Microsoft Office, Eviews, SAS, Matlab, SPSS, STAMP, Scientific Workplace, LaTEX (working knowledge)

Scholarships and grants

Scholarship from the **German Research Foundation (DFG)** (9/2003-12/2003 and 10/2004-9/2005)

Grant by **McKinsey&Company** for the academic year 04/05 at the UCLA (10/2004)

Scholarship from the **EU** project *The Economic Geography of Europe: Measurement, Testing and Policy Simulations* (EGEMTPS) (1/2004-9/2004)

Grant by the **Foundation for Economic Education** for the Seminar on *The State of Civil Society*, New York, USA (8/2002)

Scholarship from the **German Academic Exchange Service (DAAD)** (1997-2002)

Social commitment

Volunteer at Camino Seguro (www.safepassage.org), a project providing education for children at the Guatemala City garbage dump (fall 2006); continuous fundraising campaign ever since

Tutoring of underprivileged Afro- and Latin-American children in East Los Angeles (winter and spring quarter 2005 at UCLA)

Counsellor at the summer camp of Caritas Bonn (Summer 1999)

Sponsorship of Plan Deutschland, Licht für die Welt, Misereor

References

Prof. Jesús Crespo-Cuaresma
University of Innsbruck
Universitätsstraße 15
6020 Innsbruck, Austria
Email: Jesus.Crespo-Cuaresma@uibk.ac.at
Tel.: +43 512 507 7350
Relationship to the referee: PhD advisor

Prof. Arnold C. Harberger
University of California, Los Angeles
Bunche Hall 8375
Department of Economics
UCLA
Box 951477
Los Angeles, CA 90095-1477
Email: harberger@econ.ucla.edu
Tel.: +1 310 825 1011
Relationship to the referee: my former professor at UCLA

Mr. Marek Mora
Deputy Vice Prime Minister for European Affairs
Office of the Government of the Czech Republic
nábřeží Edvarda Beneše 4
CZ-118 01 Praha 1
Email: mora.marek@vlada.cz
Tel.: +420 224 002 508
Relationship to the referee: supervisor of my Masters thesis at the University of Lepzig

Publications
Peer-reviewed articles

On the Determinants of Currency Crises: The Role of Model Uncertainty (with J. Crespo-Cuaresma). Focus on European Economic Integration 2/07. Oesterreichische Nationalbank, 2007

An "Almost-Too-Late" Early Warning Indicator for Currency Crises Uncertainty (with J. Crespo-Cuaresma). BOFIT DP, Bank of Finland, 2007; after revision re-submitted to the Journal of Comparative Economics

Predicting Currency Crises Using the Term Structure of Relative Interest Rates: Case Studies of the Czech Republic and Russia Uncertainty (with J. Crespo-Cuaresma). Focus on European Economic Integration 1/07, Oesterreichische Nationalbank, 2007

Serbia: Country Profile and Recent Economic Developments (with K. Michal) in Focus on European Economic Integration 2/06, Oesterreichische Nationalbank, 2006

Non-peer-reviewed articles

Is There a Case For Argentina to Push For a Currency Union in MERCOSUR? A Cost-Benefit Analysis and Resulting Conclusions for the Economic Policy, Centro de Investigaciones Economicas, Bolsa de Comercio, Cordoba, Argentina, 2004

Unilaterale Euroisierung als eine alternative Wechselkursstrategie in den mittel- und osteuropäischen Ländern, IWH-Sonderheft, Institut für Wirtschaftsforschung Halle, Germany, Oktober 2002 (in German)

Working papers and unpublished articles
Talking About My Generation: Currency Crises and Their Determinants (with J. Crespo-Cuaresma).

(How) Will the Euro Affect Inflation in the Czech Republic?

Early Warning Systems for Currency Crises Under Model Uncertainty (with J. Crespo-Cuaresma).

Financial Markets as Predictors for Currency Crises (with J. Crespo-Cuaresma and A. Gersl)

The Introduction of Euro in the EU Acceding Countries: "Standard" EMU Accession or Euroisation? (with Z. Cech and M. Mora), mimeo

Press articles
Mocny, nikoli vsemocny trh in Hospodarske Noviny, June 14, 2007 (in Czech)

Euro bez unie in Euro, Nr. 50, December 10, 2001 (in Czech)

Presentations in conferences and seminars
- Sixth Emerging Markets Workshop, Bofit, Helsinki, May 2008
On the determinants of currency crises: The role of model uncertainty

- Staff Seminar, Oesterreichische Nationalbank, Vienna, April 2008
(How) Will the euro affect inflation in the Czech Republic?

- Graduate and Staff Seminar, Economics Department, University of Vienna, January 2008
On the determinants of currency crises: The role of model uncertainty

- Course on 'Challenges on the Road to the Euro: Inflation, Exchange Rate Stability and Credit Growth, Joint Vienna Institute, Vienna, September 2007
An "almost-too-late" warning mechanism for currency crises

- Graduate and Staff Seminar, Economics Department, University of Vienna, June 2007
An "almost-too-late" warning mechanism for currency crises

- Czech National Bank, Prague, June 2007
An "almost-too-late" warning mechanism for currency crises

- 11th Annual International Conference on Macroeconomic Analysis and International Finance, Crete, May 2007
An "almost-too-late" warning mechanism for currency crises

- Brown Bag Seminar, Economics Department, University of Mannheim, November 2006
An "almost-too-late" warning mechanism for currency crises

- Staff Seminar, Oesterreichische Nationalbank, Vienna, March 2006
An "almost-too-late" warning mechanism for currency crises