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(How) Will the Euro Affect Inflation in the Czech Republic? A contribution to the current debate

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(How) Will the Euro Affect Inflation in the Czech Republic? A contribution to the current debate

Tomáš Slačik

Abstract

In the present study we analyse relevant macro- and microeconomic forces driving inflation in the Czech Republic with a particular focus on how these channels are likely to change in the wake of euro adoption. We employ an ARDL model combined with the Bayesian Model Averaging technique. In order to carry out this analysis, we also estimate the time-varying natural rate of interest purged from the risk premium. Our results suggest that the costs arising from the discontinuation of nominal trend appreciation of the koruna 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 ensuring that inflation hikes will remain limited after the euro adoption. The fulfilment of the Maastricht inflation criterion should not be enforced by policy measures that would artificially reduce inflation temporarily. The potential inflationary effect of the changeover cannot be eliminated altogether but it may well be substantially reduced by applying best practices based on the experience of current euro area participants.

JEL classification: F31, E31, E43;

Keywords: Euro adoption, inflation, natural rate of interest, Bayesian model averaging, Czech Republic.

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1 Introduction

The current lack of an official target date for euro adoption after the originally envisaged horizon of 2010 has been abandoned does not preclude an intense and prominent discussion on costs and benefits of the euro in the Czech Republic. It is noteworthy that this debate is very much focused on one particular aspect of the cost-benefit-analysis, namely the expected impact of the single currency on inflation. This might be due to the fact that concerns about higher inflation are a key issue for the general public. According to the Eurobarometer survey from November 2007, 74 % of citizens in the EU - New Member States (NMS)¹ and 80% of the Czech respondents agree with the statement that the euro will increase prices (see Eurobarometer 2007). Since inflation expectations might have a substantial impact on future inflation rates, it is essential for the sake of future price stability to provide the public with even-handed, understandable and transparent arguments in this debate.

Hence, against this background, the objective of the present paper is to contribute putting the discussion about the possible inflationary impact of the euro in the Czech Republic on a firm and solid footing. In other words, we aim at analysing in a qualitative and quantitative manner the channels through which the euro adoption might have an effect on inflation both in the short and long run, embedded in a framework of other cyclical, structural and external inflation factors.

It should be stressed in passing that the present study does not aspire to provide a comprehensive analysis of the net welfare gains of the single currency so as to work out whether or when to adopt the euro in a particular country. It rather intends to shed some light on the specific issue of whether euro adoption will have an inflationary impact which plays a central role in the discussion in the Czech Republic. In addition, it should be born in mind that we neither address the question whether higher inflation rates in a transition country might be justifiable (or even desirable) in as much as they are driven by the catching-up process. We are merely interested in the question whether the substitution of the national currency for the euro will or can lead to a higher one-off and/or trend inflation².

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

²An additional, related question would be whether also inflation volatility around the trend will increase in the wake of euro adoption. Although we do not address this issue explicitly in this paper it can be stated that the correlation between the

To begin with, the most apparent direct consequence implied by the accession to the euro area is the replacement of national notes and coins for their euro counterparts. Though this changeover should be a mere nominal event in theory, it can be argued that it does entail some risks of price increases. However, even if such risks can probably not be eliminated altogether they might be substantially reduced by best-practice policies as the experience of a number of current euro area members suggests. Moreover, we will argue that possible price rises brought about by the changeover would be of a rather limited and temporary nature, if in particular a sufficient degree of competition is ensured.

A much more important consequence of the euro adoption than the one-off changeover effect is the fact that joining a monetary union implies the abandonment of an autonomous monetary policy. 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. Hence, were the country in question to adopt, alongside with the euro, interest rates that would be too low to ensure internal equilibrium it would most likely have to face a higher trend inflation. In order to analyse this issue econometrically we first estimate 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 changeover and the abandonment of monetary autonomy, the most conventional argument put forward in the discussion on inflationary impacts of the euro centers around the substantial price level gap with respect to the euro area 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 euro area 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 qual-

level and the volatility of inflation is typically highly positive (see e.g. Fischer, Sahay and Végh 2002).

ity 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 which we do not quite share. In contrast, we believe that it is not the loss of a flexible exchange rate per se that implies a higher inflation rate. The size of the possible price hike 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.

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 lags (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 euro area 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.

The paper is structured as follows. The following section provides some stylized facts on the inflation development of the first-wave euro area countries and the three countries that have opted out. Against this backdrop the third section 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 the fourth section we first describe the data, the remaining variables and briefly the method employed, while details are provided in the appendices. Subsequently we sketch out the results. The last section wraps up and derives interesting policy conclusions.

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. The creation of the euro area 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.³ Greece joined the club on January 1, 2001.⁴ 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.⁵ Except for Denmark and the United Kingdom which negotiated an opt-out, the Treaty on European Union (Treaty) signed in Maastricht in February 1992 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 exchange-rate mechanism (now ERM II) for at least two years preceding the convergence examination without any break and without severe tensions.⁶ Sweden is the only 'old' EU member with no opt-out from the Treaty that has not yet introduced the euro, as it has not joined ERM II and has thus not met the exchange rate criterion. In contrast, Slovenia, Malta and Cyprus have been as of today the only NMS to enter the euro area.

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,

³Belgium, Germany, Ireland, Spain, France, Italy, Luxembourg, Netherlands, Austria, Portugal, Finland

⁴The initial 11 euro area countries and Greece will be referred to as EU-12.

⁵For details on the history of the Economic and Monetary Union see e.g. www.ecb.eu.

⁶See <http://eur-lex.europa.eu/en/treaties/dat/11992M/htm/11992M.html>

2002 on which date the cash changeover took place on the other. In case of countries that joined or will join the euro area later both effects, the one brought about by the abandonment of monetary policy autonomy and the one caused by the currency changeover, fall together.

After a visual inspection of Figure 1 depicting inflation paths for EU-11 it strikes that in most countries inflation⁷ eased somewhat or remained broadly stable in the year following the creation of the euro area. 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 euro area rose from 1.5% in the two years preceding the creation of the euro area to 1.9% between 1999 and 2006. In contrast, if we compare the same periods in two EU but non-euro area 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 some inflation acceleration. Such a conclusion, however, would be premature. Figure 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 euro area-counterpart, with some lag where appropriate. 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 2 the inflation paths in Denmark matched the one in Finland very closely. Correspondingly, very similar to the euro area 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 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 Britain⁸.

⁷Measured by a 12-month-moving average of the y-o-y change of the HICP.

⁸For a more detailed analysis of the inflation development before and after the euro adoption see e.g. ECB (2008).

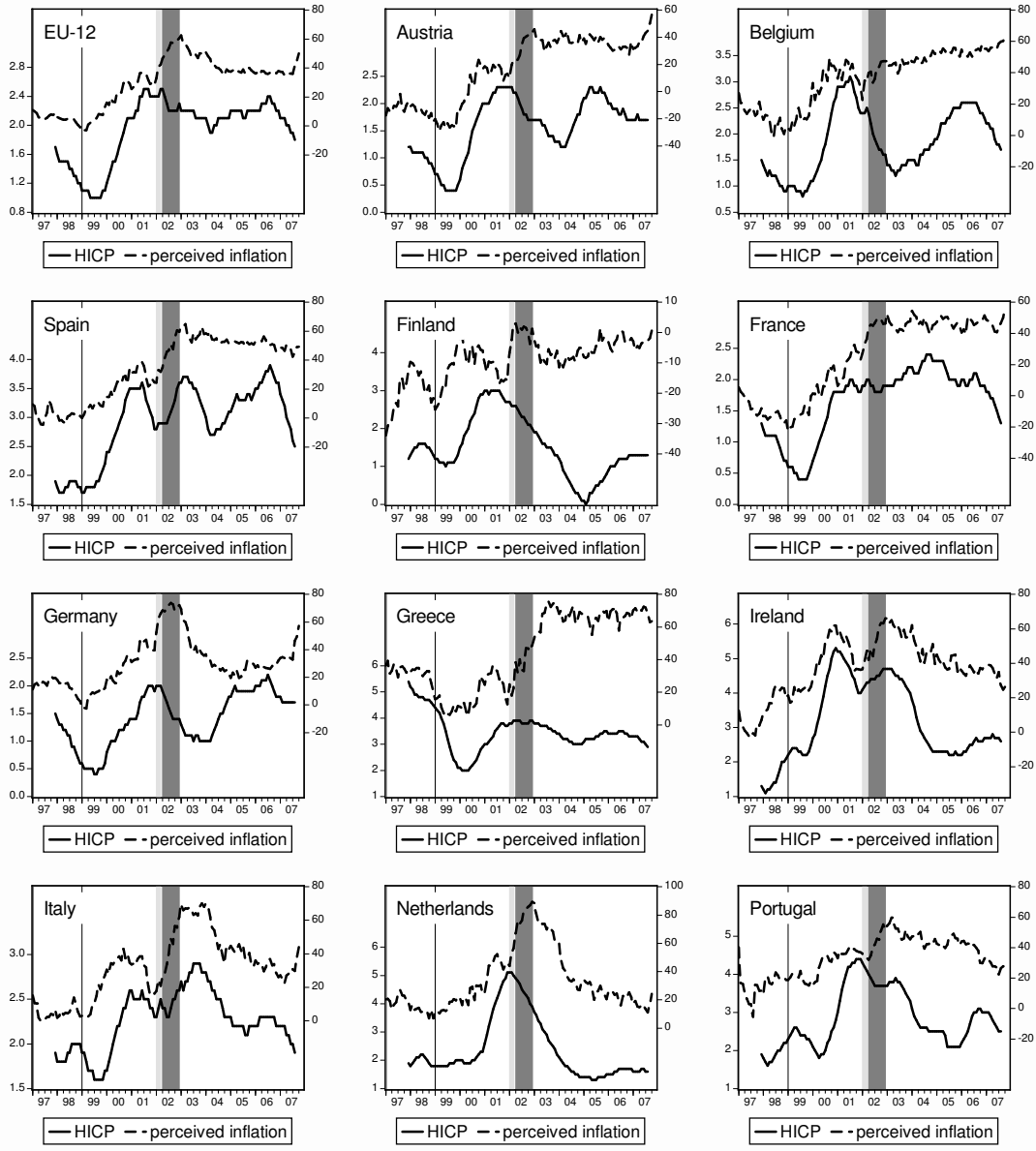


Figure 1: HICP (12-month moving average of the y-o-y change) and perceived inflation in the EU-11

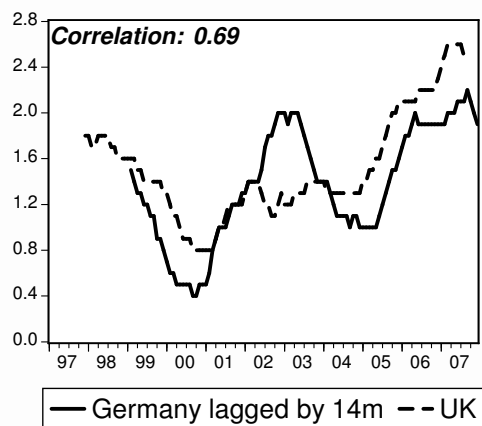
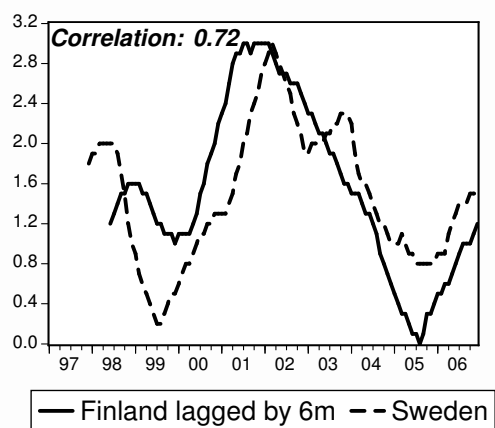
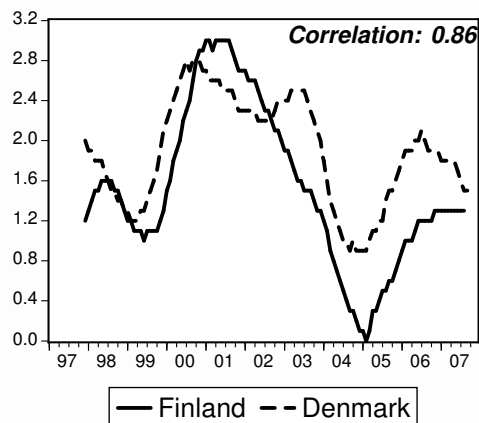


Figure 2: HICP (12-month moving average of the y-o-y change) in the non-EMU countries relative to selected EMU

Concerning the inflationary effect of the euro changeover in January 2002, a simple eyeballing of Figure 1 suggests what has been concluded by several studies after a much more sophisticated and comprehensive analysis⁹: Substantial and wide-spread price increases about which the public had been concerned prior to the introduction of euro coins and notes¹⁰ 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 euro area 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 2002¹¹ there were significant and persistent price increases in some specific branches, particularly in some services such as restaurants, cafes and hairdressers.¹² 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). As these price rises were particularly visible, they apparently affected perceived inflation much more strongly than their share in the consumer basket would have suggested.

In parallel to the debate on the impact of the euro area 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 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.¹³ The distance

⁹See e.g. Deutsche Bundesbank (2004), Eife (2006) or Hobijn et al. (2004).

¹⁰See Commission of the European Communities (2001).

¹¹The effect was probably largest in the Netherlands, 0.6 percentage points (see Folkertsma et al. 2002) and in Italy, 0.5 percentage points (Aucremagne et al. 2007).

¹²See http://ec.europa.eu/economy_finance/euro/faqs/faqs_16_en.htm

¹³There 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

between the two curves in Figure 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 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 did, indeed, accelerate 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, despite different exchange rate regimes 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. Moreover, both series are integrated of order one and the residuals of a mutual correlation are highly stationary suggesting that a spurious relationship between the series is rather unlikely.

Hence, although inflation in each respective country might have been driven to some extent by idiosyncratic factors (including euro adoption in Slovenia) it would be unfounded to mainly blame the euro for the recent inflationary development in Slovenia. 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 less¹⁴. According to the Bank of

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)

¹⁴Identical or even slightly lower impact of the euro on inflation in Slovenia was estimated also by, respectively, Eurostat and the Institute of Macroeconomic Analysis

Slovenia (2008) the rise of the inflation rate has been primarily fuelled by other determinants such as economic overheating¹⁵ 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 e.g. Eurobusiness 2007). While the actual impact of the euro on inflation thus seems to have been rather moderate, inflation perception literally skyrocketed in parallel and the gap between actual and perceived inflation has reached exceptional levels corroborating the experience in older euro area members.

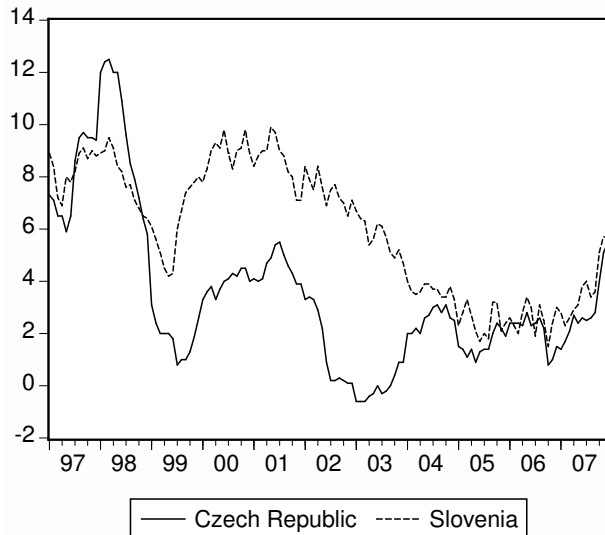


Figure 3: HICP (y-o-y change) in the Czech Republic and Slovenia

3 Factors Through Which Euro Adoption Might Impact Inflation

3.1 Long-term Factors

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

and Development of the Republic of Slovenia (IMAD) (see Bank of Slovenia (2008) for references).

¹⁵The output gap has been positive since early 2006. While it steeply widened until mid 2007 it seems to be closing ever since.

in the euro candidate country is substantially lower than in the euro area. 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 expensive (Bruha and Podpiera 2007), in the non-tradable sector it is particularly the Balassa-Samuelson (B-S) effect.¹⁶ 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. In other words, the rate of real exchange rate appreciation is given by the sum of the nominally appreciating exchange rate (the so called exchange rate channel) and the inflation differential between the home country and the euro area (i.e. the inflation channel). However, if nominal appreciation is no longer possible after the euro adoption it follows that the real exchange rate appreciation will take place completely via the inflation 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) 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.¹⁷ Mach (2003) goes even a dimension further and predicts that inflation will return to annual 10% from the first half of the 1990s.¹⁸

If our reading of the above argument sequence is correct it seems to assume that the recently observed rate of real exchange rate appreciation

¹⁶For a recent formal discussion of the Balassa-Samuelson effect see e.g. Égert (2007b).

¹⁷This 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.

¹⁸For a slightly different perspective of the same issue see Bruha and Podpiera (2007).

will be maintained also after euro area entry. In such a case, it indeed follows by definition that the inflation rate will have to rise by the current rate of nominal appreciation once the exchange rate channel is closed. However, to postulate a constant real appreciation is possible if and only if exchange rate movements are currently passed-through into consumer prices by 100% or if inflation started to accelerate for some other reason but the loss of the exchange rate channel. To judge whether this is a plausible assumption we have to investigate how high the exchange rate pass-through in the Czech Republic is and for which other reasons the introduction of the euro could lead to a discrete inflation acceleration.

Concerning the latter, a natural conjecture would be that the euro adoption will give a one-off boost to the speed of price level convergence due to higher transparency and comparability of prices. In the Czech Republic the price level of total goods reached nearly 74% of the euro area level in 2006, while the price level in services amounted to mere 42% of the euro area¹⁹. Consequently, there is certainly enough scope for price level convergence both in the tradable and the non-tradable sector. However, the literature seems to conclude almost unanimously that while price level convergence in the EU and the euro area 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). As Égert (2007a) documents and Figure 4 visualises the ranking of the price levels strongly correlates with the ordering of the countries in terms of GDP per capita. Hence, it follows that the price level is determined to a great extent by economic development and the price level convergence in the Czech Republic will thus gradually continue as the country becomes richer. However, it seems highly unlikely that the convergence process could receive a distinct 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? In general, the literature puts forward the following determinants of price level convergence (see e.g. Égert 2007 a,b): i) the shift to higher-quality goods as households grow richer; ii) changes in pricing to market practices suggesting that producers take into account the disposable income in a particular economy and adjust their price setting to the purchasing power of households; iii) the B-S-effect; iv) deregulations of administered prices and, finally, v)

¹⁹Source: New Cronos/Eurostat

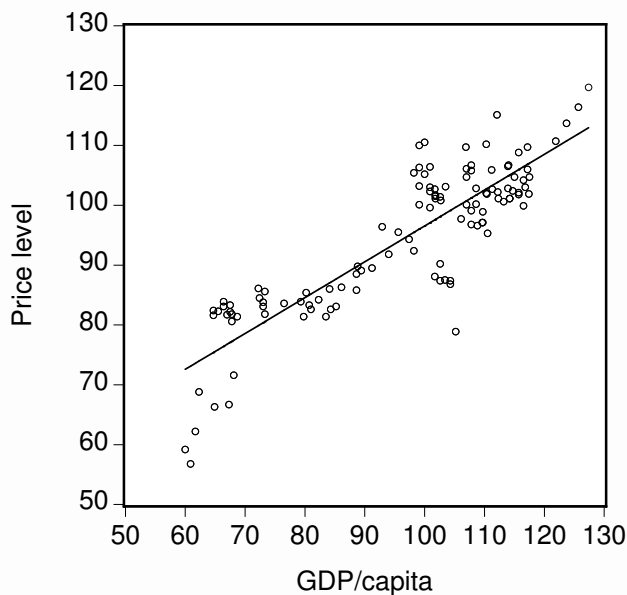


Figure 4: Price level vs. GDP/capita

the nominal exchange rate appreciation. In fact, the first three of these factors are at work irrespective of the monetary regime or the state of monetary integration. Therefore, they will not be affected by the euro adoption and are inappropriate as explanations for any discrete inflation acceleration.

Moreover, as far as the B-S-effect is concerned, empirical evidence on its significance in the Czech Republic seems mixed at best. While some older studies estimated a rather sizable B-S-effect²⁰, a whole plethora of more recent studies pioneered by Flek et al. (2002) found only a negligible impact of the B-S-effect on inflation (partially even negative, depending on the methodology). 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 Republic²¹. 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

²⁰For example, the B-S-effect contributed 4.3 percentage points to inflation in Golinelli and Orsi (2001) or 2.8 percentage points according to Sinn and Reutter (2001).

²¹For a recent discussion of the reasons for the small size of the B-S-effect which basically question the assumptions of the model see e.g. Égert (2007b).

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 compared to earlier results and range between 0.7 and 1.9 percentage points for the Czech Republic. Hence, if the strength of the B-S-effect were to increase in the future, inflation would accelerate relative to its euro area counterpart. But, to stress it again, such a development would not at all be the result of any particular monetary regime choice.

The next among the drivers of price level convergence listed above is the deregulation of administered prices. Their share in the consumer basket amounts to approximately 20% in the Czech Republic²², a number similar to the euro area level. In addition, inflation of regulated services has been almost persistently above the inflation rate of the overall consumer basket as can be seen in Figure 5. We can thus conjecture that increases of administered prices have a significant impact on the overall inflation. The outstanding question of interest in the context of the present paper is: how will regulated prices develop in the future, particularly after joining the euro area?

According to Égert (2007b), there are 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 Republic 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. Lastly, the level of liberalisation in regulated sectors is often subject to hot political debates in which some politicians might be reluctant to price increases and prefer to postpone them for later, especially until after elections.

²²Using the three-digit COICOP disaggregation level of the HICP from New Cronos and 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 protection, cultural services, rents and energy prices related to housing. 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 also be included in the definition.

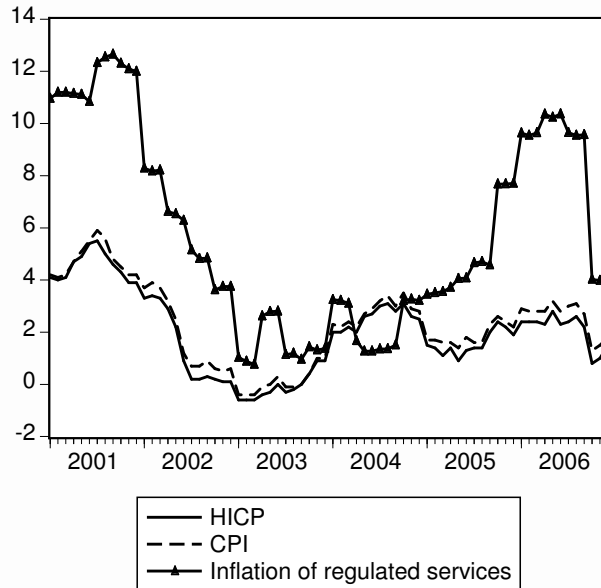


Figure 5: Inflation of price regulated services vs. HICP/CPI

While the first two arguments are not immediately related to the currency in circulation the quintessence of the last point does apply to the issue of euro adoption. Generally speaking, if a country keen on 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 - policy makers will be naturally looking for ways to take off further strain from the inflation rate. Hence, political representatives might have an incentive to strike inflation-curbing deals with the private sector and/or to postpone more than indispensable hikes of administered prices until after the euro adoption. In such a 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 for instance in Greece,²³ 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).²⁴ However, since 2001, the awareness in the euro area as regards

²³In Slovenia the government and the Bank of Slovenia helped to curb inflation with coordinated policy measures such as strict implementation of administered prices (increases not above headline inflation), counter-cyclical adjustment of excise duties on fuels and wage moderation based on a social agreement (see IMAD 2007).

²⁴Annual inflation in Greece soared from 2.1% in 1999 to 3.9% in 2002 which might

policy measures that might artificially lower inflation for some time, in particular in the context of the fulfilment of the Maastricht criteria, has risen substantially and the Treaty in fact provides good reason for alertness in this respect, given that it stresses that convergence must be found to be sustainable in order to arrive at a positive convergence assessment.

The last determinant of price level convergence we need to look at is the nominal exchange rate. While doing so we have to make a clear distinction between its role in the price level convergence process and its impact on inflation. A nominally appreciating currency increases the price level expressed in euros and thus speeds up the price level convergence. For the inflation development, however, it is not the nominal appreciation as such that matters but the pass-through of exchange rate movements on consumer prices. On the one hand, the higher the the exchange rate pass-through (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. Conversely, a lower ERPT implies a more significant convergence of price levels but provides less help in curbing inflation.

Égert (2007b) argues that the extent to which exchange rate changes are translated into domestic prices hinges primarily on the price setting practice of firms in the particular country, the composition of imports and expectations. 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.

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.

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 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. 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 the euro area by including an openness indicator defined along the lines of Égert (2007a) as the nominal exchange rate of the euro vis--vis the koruna multiplied by imports from the 'old' EU-15 Member States²⁵ 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.

3.1.2 Natural Interest Rate Too Low?

It is not only the exchange rate that is abandoned after euro area entry but in fact the autonomous monetary policy run by the national central bank.²⁶ 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 appropriate 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 ac-

²⁵Apart from the EU-12 defined above EU-15 contains also the UK, Denmark and Sweden.

²⁶To which extent the monetary policy can be autonomous in small and open economies shall be left aside.

celeration. 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 euro area then inflation might speed up. In order to understand 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 euro area?
2. 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 Horváth (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 from 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-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) aims at decomposing the time-series in a trend component, a cyclical component and an irregular component and is explained in detail in the Appendix. Whereas estimation results for an UCM for the Czech Republic are also reported in the Appendix in Table 3, Figure 6 depicts the estimated TVNIR graphically. 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 issue²⁷.

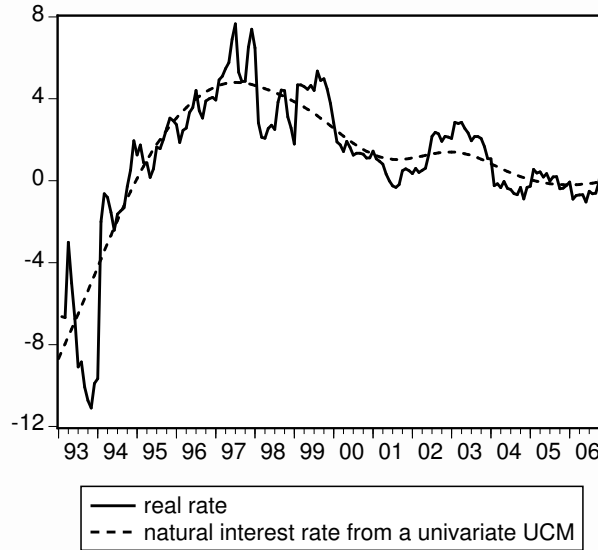


Figure 6: Real interest rate in the Czech Republic and the time-varying natural interest rate

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 euro area? - is addressed in Figure 7. The chart displays the difference between the actual real interest rate in the Czech Republic and in the euro area (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 univariate UCM for reasons explained above, for the euro area 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 euro area a multivariate version did work quite well and we came

²⁷For reasons explained in the Appendix we opted for a univariate version of the model. The estimation of a univariate UCM for industrial production looks significantly different than for the interest rate. 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 euro area. The discrepancy between the statistic features of the two series in our estimations thus provide an ex-post explanation for why they are difficult to “squeeze” into a multivariate model.

very close to the results obtained by Crespo 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 euro area reached some 2.5 percentage points on either side since 2001.

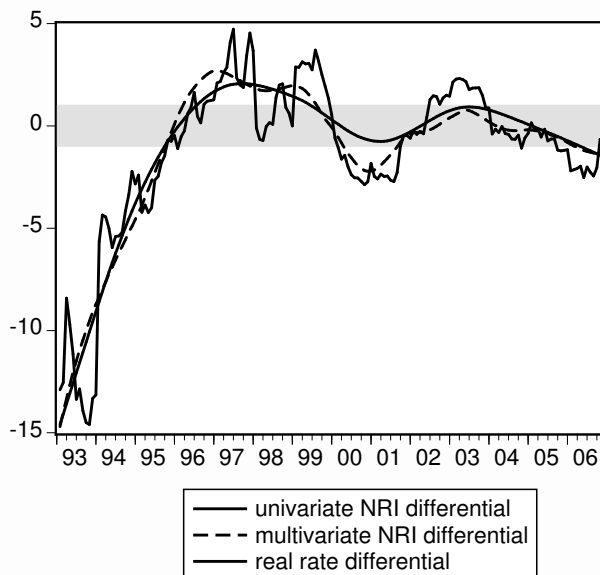


Figure 7: Real rate and NRI differentials between the Czech Republic and the EMU

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 euro area 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 [$E(\pi_{t+12}^{cr} | \{\pi_k^{cr}\}_{k=1}^t) - E(\pi_{t+12}^{emu} | \{\pi_k^{emu}\}_{k=1}^t)$]:

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

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 length²⁸ 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. Due to possible correlations between the error term γ_t and the regressors²⁹ 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 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 $\hat{\beta}_0$ and the random shock $\hat{\gamma}_t$. Hence, the nominal interest rate purged from the risk premium corresponds to

$$i_t^{adj} = i_t^{cr} - \hat{\beta}_0 - \hat{\gamma}_t \quad (2)$$

In Figure 8 we plot the adjusted real rate against the original series and for the sake of comparison also the real rate of the euro area. 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

²⁸At each t the maximum lag length is 12.

²⁹Inflation expectations, for instance, might be driven by the same factors as the risk premium.

might reflect the trend appreciation of the koruna vis-à-vis the euro.³⁰ In analogy to Figure 7 we show in Figure 9 the difference between adjusted real rate in the Czech Republic and the actual real rate in the euro area as well 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 euro area 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 euro area level in the last three years or so of the sample, the risk premium adjusted real rate has been well above the euro area rate since mid 2006.

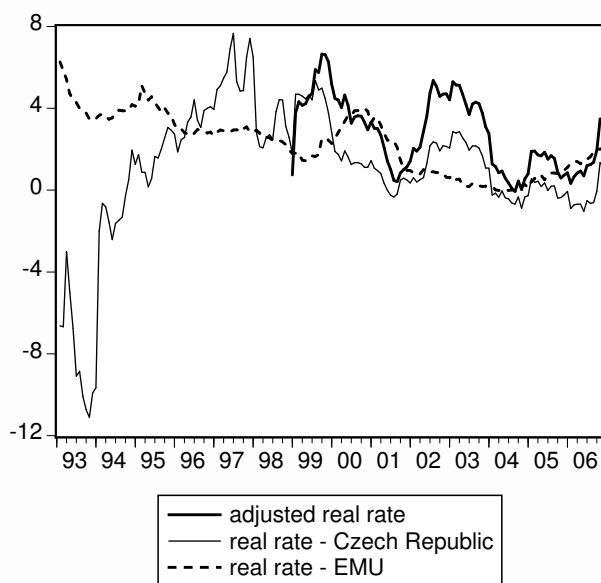


Figure 8: Adjusted and original real rate in the Czech Republic versus real rate in the EMU

Hence, the evidence provided so far suggests that despite a significant convergence of the Czech real and natural interest rates towards the euro area 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

³⁰Between June 1999 and December 2006 the koruna appreciated in nominal terms against the euro by more than 25%.

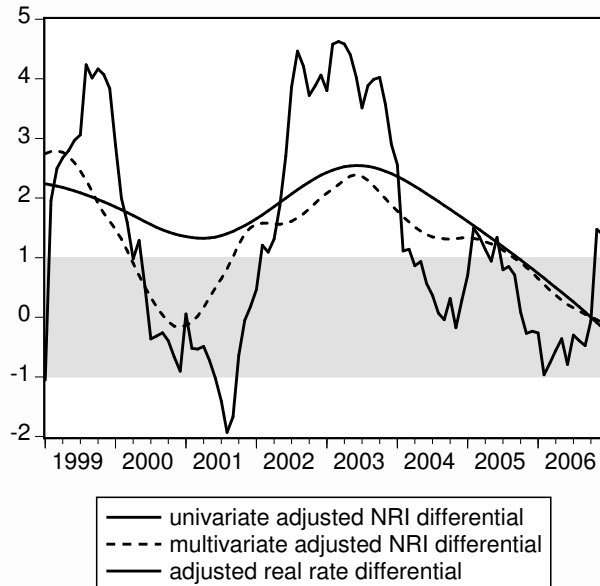


Figure 9: Adjusted real rate and NRI differentials between the Czech Republic and the EMU

path in the Czech Republic if the euro area 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 euro area 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 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).³¹

³¹Horváth (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.

3.2 One-off Factors Related to the Euro Changeover

As has been shown above, although the overall impact was moderate, in some countries the prices for some goods and services did increase noticeably after the changeover. In most cases, the effect was however transitory. Still, there were clear differences among the countries. Eife (2005) who focuses on the price setting behaviour around the changeover in Austria and Germany finds that in Austria, unlike in Germany, 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.³²

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 holds then one should observe significantly less frequent price changes before and after the changeover which was however 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.³³ On the other hand, however, there was a counteracting effect due to the higher sensitivity and

³²See 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.

³³This 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.

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 and are standardized in the sense that irrespective of the provider the consumer has quite a precise expectation of what he or 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, which 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.³⁴ In the light of the Slovenian experience, however, reliance should not be solely 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.³⁵ 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.

³⁴In 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.

³⁵While it was more than 12 months in Austria, in Germany the replacement period lasted only 2 months.

As we have seen in Figure 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 Lindn 2006). Hence, whereas the findings on these issues might not be 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.^{36 37}

In the Czech Republic, where a great part of the discussion on costs and benefits of the euro centres around the inflationary impact of the euro, a vast majority of the participants in the debate supports the hypothesis that the euro will lead to a marked inflation acceleration, i.a. also because of the change-over. So it is probably little surprising that 80% of the Czechs, 6 percentage points above the average in the New

³⁶Del 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.

³⁷For the preference ranking of trusted distributors of information see Eurobarometer (2007).

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 Lindn (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 Virn (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 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.³⁸

4 Econometric Evidence

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

³⁸There 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)

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.³⁹ 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.

To control for structural phenomena spelt out above, the pricing to market effect and the shift to higher quality goods, we include linearly interpolated data on GDP per capita at PPP⁴⁰. Concerning the B-S-effect, three different approaches have been employed in the empirical literature to derive its impact on inflation. Following Égert (2007b) we will use 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

$$\Delta p^{B-S} = (1 - \alpha) (\Delta prod^T - \Delta prod^{NT}). \quad (3)$$

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 sec-

³⁹The measures are defined as the relative gap between actual and, respectively, trend or potential gross domestic product at 2000 market prices.

⁴⁰Source: New Cronos. Since GDP per capita appears to be growing almost exponentially we defined this variable as 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

tor 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.

Apart from the two already mentioned alternatives which are supposed to capture the exchange rate pass-through the set of external variables includes 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.⁴¹ 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 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

$$\pi_t = \mu + \sum_{i=1}^p \gamma_i \pi_{t-i} + \sum_{j=0}^r \beta_j x_{t-j} + \varepsilon_t \quad (4)$$

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

⁴¹I would like to thank my mother, a tax adviser, for helping me out with the tedious work through the tax code.

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 $2^{\bar{K}}$, where \bar{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 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 (MC³) method explained in the Appendix.

There are, however, a couple of major problems with these two 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 so called Bayesian model averaging (BMA) methodology which proposes averag-

ing 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). A brief explanation of this method following closely Crespo-Cuaresma and Slacik (2007) is provided in the Appendix.

For each variable we sum the posterior probabilities of those models visited by the MC³ 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, our prior on the inclusion probability of each variable is 0.5.⁴² 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 regressors 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.

4.2 Estimation Results

The application of the BMA methodology in combination with the MC³ 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 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.

⁴²There 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 $2^{\bar{K}-1} / 2^{\bar{K}} = 0.5$. Intuitively it means that there is a 50:50 chance for a new variable to belong to the true model or not.

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

Variable	BMA			Frequentist check		
	$E(\beta_i/Y)$	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 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 (4)(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.⁴³ The results changed slightly though, if we employed the natural rate of interest obtained by Horváth (2007) as can be seen in Table 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

⁴³We did not experiment only with various alternative definitions of the variables but also with the number of variables initially included.

than that, the remaining variables showed up at the same lags and the coefficients 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 increasing income of households seem to be accelerating 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. Unless we interpret the first lag of the inflation variable as a backward-looking inflation expectation, our results provide very little support for expectations 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.⁴⁴

⁴⁴The 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	$E(\beta_i/Y)$	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 2: Estimation of the ARDL model employing NRI obtained by Horváth (2007)

5 Summary and Policy Conclusions

The debate on costs and benefits of the euro in the Czech Republic is very much focused on one particular aspect – the impact of the common currency on inflation. Against this background, in the present study we analyse in a qualitative and quantitative manner the channels through which the euro adoption might have an effect on inflation both in the short and long run, embedded in a framework of cyclical, structural and external inflation factors. As we do not aspire to provide an overall analysis of the net welfare gains of the single currency but concentrate on one specific issue of monetary integration in this paper, our findings do not allow to answer the question of when to adopt the euro in the Czech Republic.

After a brief sketch of some stylized facts drawn from current euro area members in the first module of the paper we proceed in the second one to a qualitative discussion of both, the long-term determinants impacting trend inflation as well as the one-off inflation risks related to the changeover. With respect to the former we argue that neither higher price transparency nor the usual continuous drivers of price level convergence can explain any discrete inflation jump after the adoption of the euro. However, for the sake of a sustainably stable inflation the fulfilment of the Maastricht inflation criterion should not be enforced by any non-standard policy measures. As regards inflationary risks of the changeover, we illustrate that there are better and worse ways to handle them and although they can probably not be eliminated altogether they might be substantially reduced by best-practice policies as the experience of a number of current euro area members suggests.

In the third and last building block of the paper we employ an ARDL model to test the relative importance of various inflation drivers. Since standard model selection criteria do not provide a clear picture, 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. An additional innovation of our paper is the estimation of 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 earlier empirical studies, in our results, too, the exchange rate pass through appears rather weak which implies a faster price level convergence in case of a currency appreciation but a relatively low dampening effect of nominal appreciation on CPI inflation. In addition, according to our estimates neither fiscal policy (including VAT changes), nor cyclical factors, nor the increasing wealth of households have an accelerating impact on inflation. In contrast, a recently observed inflation rates are highly significant for future inflation rates. This result, which might be interpreted as a backward-looking inflation expectations, suggests that a low inflation environment is very supportive for price stability in the future. Inflation also seems quite sensitive to deviations from the natural interest rate. Hence, a harmonization of business cycles between the Czech Republic and the euro area seem to be an important prerequisite to contain inflation volatility after euro adoption.

A final aspect is the euro's impact on perceived inflation. Given the loss aversion of economic agents which has been well documented by experimental economics it is certainly a challenging task to tackle the rife and deeply embedded opinion that the euro raises prices. To address this issue as effectively as possible, an even-handed, transparent and fair information campaign on the inflation aspects of the euro is key.

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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: 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 4: Risk premium estimation

Variable name	Definition
GDP/capita at PPP	m-o-m change of y-o-y rate
Productivity diff.	$(1 - \alpha) (\Delta prod^T - \Delta prod^{NT})$, $\alpha = 0.25$; productivity = gross value added/total hours worked (see text)
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	cyclical component from UCM/trend component from UCM. Alternatively, two 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	central gov. expend./GDP (y-o-y change)

Table 5: Basic variables: Definitions

6.1 Unobserved Components Models

The estimation of an unobserved components models is based merely on the statistical properties without imposing any economic theory. This method which goes back to Harvey (1989) 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

$$z_t = \mu_t + \phi_t + u_t; u_t \sim IID(0, \Sigma_u) \quad (5)$$

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

$$\mu_t = \mu_{t-1} + \kappa_{t-1} + \tau_t; \tau_t \sim IID(0, \Sigma_\tau) \quad (6)$$

$$\kappa_t = \kappa_{t-1} + \psi_t; \psi_t \sim IID(0, \Sigma_\psi) \quad (7)$$

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 parameters⁴⁵

$$\begin{pmatrix} \phi_t \\ \phi_t^* \end{pmatrix} = \rho \left[\begin{pmatrix} \cos \lambda & \sin \lambda \\ -\sin \lambda & \cos \lambda \end{pmatrix} \otimes I \right] \begin{pmatrix} \phi_{t-1} \\ \phi_{t-1}^* \end{pmatrix} + \begin{pmatrix} \omega_t \\ \omega_t^* \end{pmatrix} \quad (8)$$

where $(\omega_t, \omega_t^*) \sim IID(0, I \otimes \Sigma_\omega)$ 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 ($\alpha_t^f = E(\alpha_t | \{z_t\}_{t=0}^{t-1})$).⁴⁶

⁴⁵A univariate (non-stochastic) cycle can be most conveniently expressed as a mixture of sine and cosine waves: $\psi_t = \alpha \cos \lambda t + \beta \sin \lambda t$, $t = 1, \dots, T$. 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 $\psi_0^* = \beta$.

⁴⁶In contrast, smoothed estimates would exploit information contained in the entire sample ($\alpha_t^s = E(\alpha_t | \{z_t\}_{t=0}^T)$).

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.

6.2 Bayesian Model Averaging and MC³

A way to combine models with a proper weighting is the so called 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). In what follows we provide a brief explanation of this method following closely Crespo-Cuaresma and Slacik (2007).

In a situation where there are $M = 2^{\bar{K}}$ competing models, $\{M_1, \dots, M_M\}$ ⁴⁷ which are defined by the choice of independent variables, Bayesian inference about the parameter of interest β_i is based on its posterior distribution $P(\beta_i|\mathbf{Y})$ (that is, the distribution given the data, $\mathbf{Y} = \{y \ \mathbf{X}_K\}$),

$$P(\beta_i|\mathbf{Y}) = \sum_{m=1}^M P(\beta_i|\mathbf{Y}, M_m)P(M_m|\mathbf{Y}), \quad (9)$$

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

$$P(M_k|\mathbf{Y}) = \frac{P(\mathbf{Y}|M_k)P(M_k)}{\sum_{m=1}^M P(\mathbf{Y}|M_m)P(M_m)}. \quad (10)$$

The posterior model probabilities can thus be obtained as the normalized product of the integrated likelihood for each model ($P(\mathbf{Y}|M_k)$) and the prior probability of the model ($P(M_k)$). 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, (10) can be obtained directly and (9) can be computed. In particular, the expected value of β and its variance, $E(\beta|\mathbf{Y})$ and $\text{var}(\beta|\mathbf{Y})$ respectively, can be computed as follows

$$E(\beta_i|\mathbf{Y}) = \sum_{m=1}^M E(\beta_i|\mathbf{Y}, M_m)P(M_m|\mathbf{Y}), \quad (11)$$

$$\text{var}(\beta_i|\mathbf{Y}) = \sum_{m=1}^M [\text{var}(\beta_i|\mathbf{Y}, M_m) + E(\beta_i|\mathbf{Y}, M_m)^2]P(M_m|\mathbf{Y}) \quad (12)$$

$$-E(\beta_i|\mathbf{Y})^2 \quad (13)$$

The posterior mean and variance can thus be used to make inference while they explicitly take into account model uncertainty. Several methods have been proposed for approximating the expression in (10) when the cardinality of the model space makes the problem intractable (see Raftery, (1995) for an excellent description of these methods).

Leaning on the work of Madigan and York (1995) and the recent empirical application to the determinants of currency crises in Crespo-Cuaresma and Slacik (2007) we also used in our two model selection

⁴⁷ \bar{K} is again the number of regressors.

strategies that consider the entire, but intractable model space (the just spelled out BMA and the previously described SIC-strategy) the Markov Chain Monte Carlo Model Composite (MC³) algorithm. This method implements the Random Walk Chain Metropolis-Hastings algorithm in the model space as follows. In a given replication r of the algorithm, a candidate model M^{r+1} is randomly drawn from a set which consists of all models that contain either one variable more or one variable less than the model active in that replication (M^r). The proposed model, M^{r+1} , is accepted with a probability given by

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

which is just the Bayes factor comparing M^r and M^{r+1} if equal prior probability is assumed across models, so that $P(M^r)$ and $P(M^{r+1})$ 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.