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# Labour Market Performance and Macroeconomic Policy: The Time Varying NAIRU in the CR

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## 1. Introduction

During the second half of the 1990s the Czech economy experienced a sharp increase in the unemployment rate. Were these movements caused by structural changes, worsening labour market performance, or just by the changing business cycle position? Answering this question has direct implications for both monetary and fiscal policies.

Czech National Bank currently uses the deviation of actual output from its potential level as a basis for decisions on interest rates. Analogously, the Ministry of Finance uses potential output for determining the structural and cyclical component of the budget deficit. The potential output level depends, among other influences, on labour market functioning. This is reflected by the "equilibrium" rate of unemployment, which can be approximated by the "non-accelerating inflation rate of unemployment" (NAIRU), i.e. the rate of unemployment at which inflation remains constant. However, the approaches of the above-mentioned institutions to potential output determination either do not deal with the notion of the NAIRU explicitly, or deal with it in a form that remains open to further discussion.[1]

A crucial question that has to be addressed is whether the NAIRU remains stable over time. According to Gordon (1997), the decrease in the NAIRU in the USA during the 1990s was associated with relatively weak trade unions, a relatively low minimum wage and a slight decline in labour's income share. By contrast, Krugman (1994) or Blanchard (1999) find rather the opposite tendencies in Western Europe – caused, among other influences, by an inappropriate institutional design of labour market institutions *vis-à-vis* supply shocks – which contributed to the constant unemployment growth since 1960.[2]

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[1] At present, the NAIRU is involved in the CNB's modeling process rather implicitly, i.e. it could be argued that it is behind each value of the potential output growth rate, which is subject to explicit modeling using the Kalman filter. This, however, makes it more difficult to assess the direct impact of labour market dynamics on economic performance. The approach of the Ministry of Finance is based on modeling potential output using a two-factor Cobb-Douglas production function with total factor productivity, where employment is adjusted for the NAIRU (see http://www.mfcr.cz). As we argue later in the text, the methodology applied there for NAIRU approximations is still subject to methodological disputes.

Although at this stage of the research we are not sure about the precise role played by these particular aspects of labour market functioning in the Czech Republic, we can infer that, in this case too, increases in the NAIRU would signal diminishing labour market efficiency and, subsequently, a slowdown in potential output growth.[3]

This paper is organized as follows. In Section 2 we introduce the NAIRU model and its properties implied by the theory. Section 3 describes the data. In Section 4 we present model specifications in a testable form. The estimation results are presented in Section 5, and the last section summarises our conclusions.

## 2. The Model

Unfortunately, the NAIRU is not directly observable. Some combination of economic and statistical reasoning must therefore be used to estimate it from observable data. Several estimates of the NAIRU have already been made for the Czech Republic. They differ, however, in the economic models used and also in estimation methodology.

Vašíček and Fukač (2000) use a backward-looking model of inflation expectations for an open economy, employing the Kalman filter. We will show later in the text that assuming pure backward-looking agents can lead to downward-biased estimates. Fukač (2003) extends the model of Vašíček and Fukač (2000) by incorporating forward-looking inflation expectations, which makes it closer to our methodology. Similarly, Benes and N'Diaye (2004) use the Kalman filter to estimate potential output and the NAIRU. However, their methodology reflects some *a priori* assumptions about the transmission mechanism of monetary policy, whereas in our approach this information is extracted from data. Besides that, Benes and N'Diaye (2004) are more oriented on comparing the Kalman filter with Hodrick-Prescott filter results rather than on precise estimations of the NAIRU.

Another analysis aimed at estimating the Czech NAIRU is provided by Hájek and Bezděk (2001). It explores a Hodrick-Prescott statistical filter, without any economic restrictions. Finally, Bezděk, Dybczak and Krejdl (2003)[4] follow the Elmeskov (1993) methodology, which differs quite substantially from ours. It must be said that the Elmeskov methodology is disputable from a theoretical viewpoint – in the first step it assumes the existence of a constant NAIRU and only subsequently is the time-varying NAIRU estimated.

Our focus on possible time movements of the NAIRU instead of on its fixed value implies the use of the time-varying NAIRU methodology, following, for example, Gordon (1997) or Greenslade *et al*. (2003). This methodology uses Gaussian maximum likelihood methods, as described, for

[2] See, among others, also (Elmeskov, 1993), (Richardson *et al*., 2000), and (Turner *et al*., 2001).

[3] See (Hurník – Navrátil, 2005) for a sensitivity analysis of the potential output growth (based on a production function approach) to NAIRU movements in the Czech Republic.

[4] This study reflects the framework used by the Czech Ministry of Finance for determining the structural and cyclical component of the budget deficit.

example, by Hamilton (1994),[5] and is frequently used for estimating unobservable variables. Its advantage is that it allows us to estimate simultaneously the NAIRU and the relationship between inflation and the unemployment gap (i.e. the deviation of actual unemployment from the NAIRU).

The Phillips curve (Phillips, 1958) has become a generic term for any relationship between the rate of change in nominal prices (or wages) and the behaviour of real indicators of demand intensity, such as the unemployment rate or output. Gordon (1997) uses the following backward-looking form of the model:

$$p_t = \alpha_B(L) \cdot p_{t-1} + \beta(L) \cdot d_t + \eta(L) \cdot h_t + e_t \tag{1}$$

Where $p_t$ denotes the first differences of the logarithms of the price level (quarterly inflation), $d_t$ stands for the logarithm of the excess demand index (on product or the labour market) normalised at zero, $h_t$ is the vector of supply-shock variables, $L$ is the polynomial in the lag operator and $e_t$ is a serially uncorrelated error term. As we are interested in estimating the time-varying NAIRU, equation (1) can be rewritten as follows:

$$p_t = \alpha_B(L) \cdot p_{t-1} + \beta(L) \cdot (u_t - u_t^*) + \eta(L) \cdot h_t + e_t \tag{2}$$

The term $(u_t - u_t^*)$ represents the unemployment gap, i.e. the difference between the actual unemployment rate and the time-varying NAIRU. We introduce several changes in this approach in order to capture more properly the character of the Czech economy as a small open economy with the central bank operating in a forward-looking, inflation targeting regime. Namely, we extend the analysis for forward-looking element and for exchange rate pass-through.

In general, there are two strategies for catching supply shocks and filtering out their influence on the NAIRU and inflation. The first is to use an explicit set of variables describing such shocks. The alternative strategy is based on removing the influence of supply shocks on inflation using core inflation ($P_t^{core}$) as a key inflation variable. It represents headline inflation ($P_t^{cpi}$) excluding regulated and energy price movements. We prefer to follow the second approach as the regulated and energy prices were the source of the main inflationary shocks in the Czech Republic.

In line with Driver *et al.* (2003), we extend the analysis by incorporating the forward-looking element of inflation expectations into equation (2). This form of the Phillips curve is in accordance with the New Keynesian paradigm and is derived from the following microfoundations:[6] Firms set prices to maximise profit, but – due to adjustment costs such as menu costs and long-term contracts – they are not able to adjust prices every period. Therefore, they set prices to be optimal not only for the current period but also for

[5] See also (Blake, 2002), (Basdevant, 2003), or (Harvey, 1989). This method for estimating the NAIRU represents a reduced-form approach combining structural and purely statistical methods.

[6] For the derivation of the New Keynesian Phillips curve see, for example, (Fuhrer – Moore, 1995), (Christiano – Eichenbaum – Evans, 2001), (Galí – Gertler, 1999), or (Calvo, 1983).

the forthcoming periods when they cannot change prices. It follows that firms do not set prices solely according to the current economic conditions but also according to those expected to prevail in the future.

Therefore, we add a new variable $E_t p_{t+1}$ into equation (2) which denotes a forward-looking expectation formed at time $t$ about price level developments in the next period, $t + 1$. In this specification of the Phillips curve, inflation is determined by lagged and expected inflation and by the unemployment gap. The lagged values of inflation combine the backward-looking part of inflation expectations and the inherent sluggishness in inflation. If a forward-looking element is significant, then – in the case of inflation decreasing over time – the estimated NAIRU path without incorporating the forward-looking component would be downward-biased, as economic agents base their inflation expectations not only on previous (higher) inflation.

Inflation expectations are not directly measurable.[7] The first possible way of obtaining inflation expectations is to extract them from the yield curve. This strategy comes from the Fisher equation and assumes that the real interest rate is stable and movements in nominal interest rates are given by changes in inflation expectations. But, as for example Kotlán (2002) shows, there are also other factors influencing the yield curve and real interest rate movements. Therefore, it is hard to extract inflation expectations solely from the yield curve without having any knowledge of the whole economic model.

A second way is to use inflation expectation surveys. In the Czech Republic, however, inflation expectation surveys (of households, firms or the money market) are available only from May 1999, which constrains the data sample substantially.[8]

The third method assumes that forward-looking expectations can be represented by perfect foresight, i.e. $E_t p_{t+1} = p_{t+1}$. This implies that the forward-looking agents are not only able to foresee perfectly future inflation, but also the shocks that will hit the economy. Perfect foresight can be expected to result in an upward bias in the estimated impact of expected inflation. Having in mind all pros and cons, including data availability and the resulting problems with using monthly data for analysis, we follow this methodology in order to avoid the bias that would be certainly present in the results of pure backward-looking alternative. At the same time, we are aware of the fact that an optimum method of formulating forward-looking expectations still represents a future research challenge.

In an open economy, real economic activity is also affected by external developments.[9] Thus the unemployment gap may not catch all the inflation pressures resulting from excess economic activity. In order to reflect these external influences on prices (through an "indirect" channel *via* real eco-

[7] As Figure 2 documents, the recent disinflation period was not smooth and gradual; there were also periods with increasing inflation. Therefore, a simple linear decreasing trend cannot properly approximate the inflation expectations.

[8] Despite this constraint, the results based on inflation expectations survey data are briefly discussed in the empirical part of our paper.

[9] See, for example, (Obstfeld – Rogoff, 1996), (McCallum – Nelson, 2001), (Barro – Tenreyro, 2000), or (Galí – Monacelli, 2002) for more detail.

nomic activity), we add into equation (2) the real exchange rate gap ($z_t$), i.e. the deviations of the real exchange rate from its "equilibrium" value (see later in the text for a definition). In addition to that channel, there is a "direct" price channel, as prices of imported goods directly influence the aggregate price index. To capture this effect, equation (2) is extended to include import price inflation ($p_t^{imp}$), or alternatively the change in the nominal exchange rate.

Thus we deal with both the real and nominal determinants of inflation. The former group is represented by the unemployment gap and the real exchange rate gap. In the latter group, there are lagged inflation, inflation expectations and import price inflation. The sum of the coefficients of the nominal inflation determinants must in practice equal unity to ensure the existence of the natural rate of unemployment consistent with the constant rate of inflation (the homogeneity condition), i.e. the existence of a vertical Phillips curve in the long run. Although there are some counter-arguments to the existence of long-run Phillips curve verticality, raised especially by Akerlof *et al.* (2000), we follow the mainstream approach and impose $1 - \alpha_1 - \alpha_2 = 0$.[10] Our final estimated equation is as follows:

$$p_t^{core} = \alpha_1 \cdot E_t p_{t+1}^{core} + \alpha_2 \cdot p_{t-1}^{core} + (1 - \alpha_1 - \alpha_2) \cdot p_t^{imp} + \\ + \beta \cdot (u_{t-1} - u_{t-1}^{*}) + \gamma \cdot z_{t-1} + e_t \tag{3}$$

## 3. The Data

The data used for the analysis cover the period 1Q1994–1Q2004.[11] As we are strictly limited by the available data, we do not follow Gordon (1997), who uses a large lag structure. In particular, we are not able to adopt the polynomial lag distribution for $p_t$ and for $(u_t - u_t^{*})$. We use instead only one lag for inflation and for the unemployment gap. The data are seasonally adjusted using the X11 procedure.

In our basic model, we use the ILO definition for measuring unemployment, as published in the Czech Labour Force Surveys. This indicator appears to measure "true" unemployment more adequately than the Ministry of Labour and Social Affairs (MLSA) method.[12] It is released quarterly and this determines the use of quarterly data in the present analysis. *Figure 1* demonstrates the differences. Unemployment according to the MLSA definition is used as a robustness check of our estimates in an alternative unemployment model.

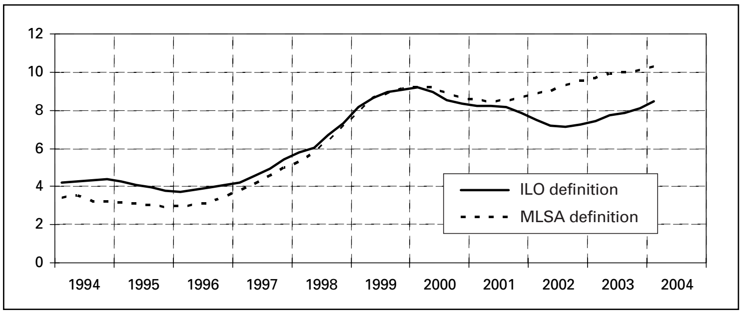
The price dynamics are approximated by core inflation ($p_t^{core}$), which ex-

[10] See, among others, also (Christiano – Eichenbaum – Evans, 2001). For further criticism of the mainstream approach, see (Franz, 2003).

[11] We exclude the year 1993 because of the split-up of the Czech and Slovak Federation and the introduction of VAT. We suspect that both factors would have substantially affected the variables in question.

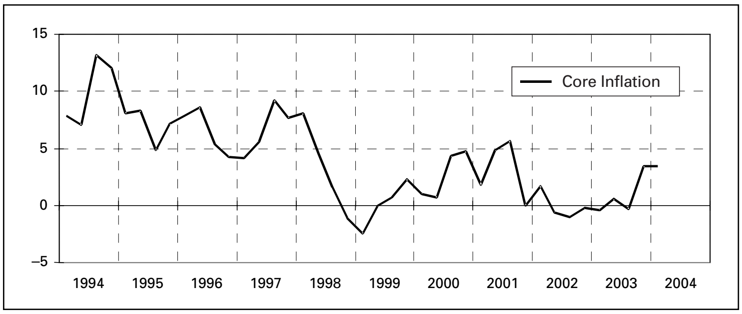
[12] According to the ILO definition, the unemployment measure involves all persons aged 15+ who satisfy all of the following three conditions: (i) are without a job; (ii) are actively seeking a job; (iii) are immediately available for a job. By contrast, unemployment measured by the Ministry of Labour and Social Affairs simply counts those registered as "unemployed".

FIGURE 1 Unemployment Rate (in %, seasonally adjusted)




*Sources*: Czech Statistical Office; Ministry of Labour and Social Affairs

FIGURE 2 Core Inflation (in %, annualised, seasonally adjusted)




*Source:* Czech Statistical Office; Czech National Bank

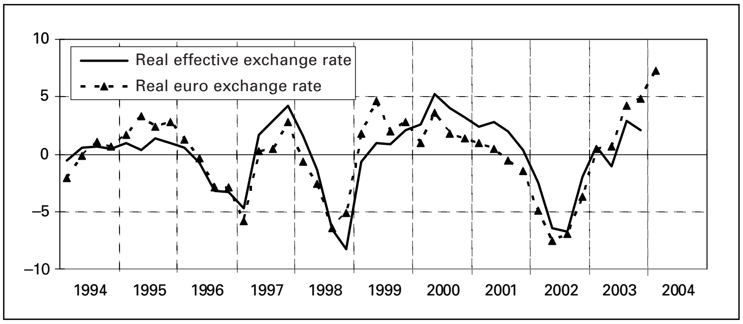
cludes the direct influences of changes in regulated and energy prices. *Figure 2* shows the quarterly annualised core inflation rates.

To approximate the real exchange rate gap in *Figure 3*, we use the real effective exchange rate (the nominal effective exchange rate deflated by the Czech and effective foreign inflation).[13] The real effective exchange rate trend is then obtained by using the Hodrick-Prescott filter with the coefficient $\lambda = 1600$. This approximates in our basic model the equilibrium real exchange rate.[14] Positive values in Figure 3 mean an undervalued exchange rate *vis-à-vis* the trend, and *vice versa*. To check the robustness of the estimation, we also use a detrended real exchange rate against the euro as

[13] *REER = NEER*(CPId / CPIf)*. Effective foreign inflation (*CPIf*) is a weighted inflation of trade partners of the Czech Republic. The shares in total Czech foreign trade determine the respective weights.

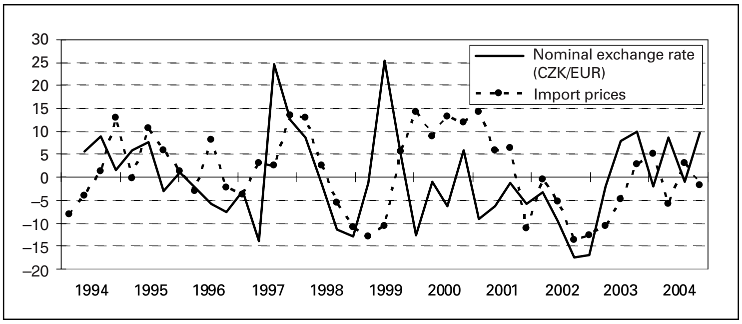
[14] We are aware that our trend approximation is not the only approach to defining and computing the "equilibrium" real exchange rate (note, among others, the FEER or NATREX methodology). However, given that this research is primarily focused on labour market performance, at this stage we do not go deeper into examining the sensitivity of the NAIRU to alternative exchange rate equilibrium concepts.

FIGURE 3 Real Exchange Rate Gap (in %)




*Sources:* Czech National Bank; Czech Statistical Office; Federal Statistical Office of Germany; International Monetary Fund

FIGURE 4 Change in the Nominal Exchange Rate (in %, annualised) and Import Price Inflation (in %, annualised, seasonally adjusted)




*Sources:* Czech National Bank; Czech Statistical Office; Federal Statistical Office of Germany; International Monetary Fund

the EU is the main trading partner of the Czech Republic, accounting for approximately 60 % of total Czech imports. Note also that we prefer to deflate the nominal exchange rate by the CPI instead of some labour cost measure, as in this analysis we are predominantly interested in the information incorporated into CPI inflation.

To catch the "direct" foreign channel, we can use the change in the nominal exchange rate (CZK/EUR). Indeed, we do so in our alternative import inflation model. However, this relies on a precondition that the law of one price holds. If pricing-to-market prevails, then the change in import prices is a better indicator of the direct price pressures stemming from other countries – see, for example, (Betts – Devereux, 1996) for a more detailed explanation.[15] Thus we use import price inflation in our basic model. *Figure 4* shows both indicators.

[15] Our definition of import prices excludes energy price movements.

TABLE 1 Unit Root Test

| | KPSS |
|---|---|
| Core inflation | 0.100 |
| Real euro exchange rate gap | 0.057 |
| Real effective exchange rate gap | 0.054 |
| Nominal exchange rate gap (1st difference) | 0.140 |
| Import prices (1st difference) | 0.115 |

*Notes:* Test includes constant (critical value at 10% level is 0.347) or constant and trend in case of core inflation (critical value at 10% level is 0.119).

For unit root testing we use the Kwiatkowski-Phillips-Schmidt-Shin (1992) test (KPSS). The KPSS null hypothesis is (trend) stationarity. As Amano and Norden (1992) show, the KPSS test can reduce the incorrect conclusions of the traditional ADF test, especially in the case of a small sample. As indicated in *Table 1*, for all the variables involved we cannot reject the null (stationarity) hypothesis.

## 4. Estimating the NAIRU

The use of Gaussian maximum likelihood methods for estimating the NAIRU combines inflation equation (3) with equation (4), which describes the explicit path of the NAIRU. Equation (3) is a "signal" equation and equation (4) is a "state" equation.

$$\begin{gathered} u_t^* = u_{t-1}^* + \varepsilon_t \\ e_t \sim N(0,\sigma_e^2);\ \varepsilon_t \sim N(0,\sigma_\varepsilon^2) \\ \operatorname{cov}(e_j,\ \varepsilon_k) = 0 \end{gathered} \tag{4}$$

The error term $\varepsilon_t$ in state equation (4) is expected to be white noise with standard deviation $\sigma_\varepsilon$ · If $\sigma_\varepsilon = 0$, then the NAIRU is constrained to be constant and the estimation is quite simple. But if $\sigma_\varepsilon \neq 0$, then the NAIRU is changing over time and the estimation is more complicated. Specification (4) implies that the NAIRU follows a random walk and changes in the NAIRU are driven by $\sigma_\varepsilon$. The disturbance vectors $e_t$ (from (3)) and $\varepsilon_t$ are assumed to be uncorrelated with each other in all time periods.

If there is no constraint imposed on the standard deviation $\sigma_\varepsilon$, we shall see high volatility in the estimated NAIRU series which appear hard to interpret from an economic point of view. To obtain the results in less volatile form it is common to smooth the original estimated trajectory of the time-varying NAIRU. Gordon (1997) recommends *a priori* setting of the standard deviation $\sigma_\varepsilon$, and in such a way limiting the possible fluctuations of the NAIRU directly. This means constraining the standard deviation $\sigma_\varepsilon$ within the two-equation system, i.e. setting explicitly the signal-to-noise ratio ($\sigma_\varepsilon^2 / \sigma_e^2$). The second approach is to use external smoothing techniques such as the Hodrick-Prescott filter with a well-known $\lambda$ coefficient at a low level.

We rely on another smoothing technique. We compute smoothed series of the NAIRU by backward recursion, as proposed, for example, by Harvey (1989, 1993) or Hamilton (1994). In such a way we obtain the conditional forecast and its standard error, which minimises the mean squared error. First, we compute the conditional forecasts $u^*_{t+1|t}$, $u^*_{t|t-1}$, and mean squared errors $P_{t+1|t}$, $P_{t|t-1}$, for $t = 1,2,\ldots,T$; see (5) and (6). The smoothed series and mean squared errors are then computed by backward recursion:

$$u^*_{t|T} = u^*_{t|t} + (u^*_{t+1|T} - u^*_{t+1|t}) \tag{5}$$

$$P_{t|T} = P_{t|t} + (P_{t+1|T} - P_{t+1|t}) \tag{6}$$

$$\text{for } t = T - 1,\ T - 2,\ldots 1$$

This method enables us to avoid the problem with the *a priori* (arbitrary) choice of standard deviation in the Gordon fashion, or of the $\lambda$ coefficient in the Hodrick-Prescott filter. This estimation methodology, however, rests on the assumption that the disturbances $\varepsilon_t$ and $e_t$ are normally distributed. If this assumption does not hold, then the maximum likelihood methodology cannot give the conditional mean of the state vector. Then the estimated values of the NAIRU could be biased.[16]

## 5. The Results

We estimate a two-equation system where the signal equation is given by (3) and the state equation by (4). Our basic model (A) involves the ILO definition of unemployment, the real effective exchange rate gap and import inflation. We present the results of estimating this system with incorporated forward-looking expectations, as well as pure backward-looking expectations. We also present estimates with an alternative measurement of unemployment, the real exchange rate gap and import inflation. Because of the short data sample we cannot test whether or not our estimates are time-invariant.[17] *Table 2* contains the estimated coefficients and their $t$-statistics in brackets.

From our basic model (A) we can derive the proportion of backward- and forward-looking agents in the economy $(\alpha_1/(\alpha_1 + \alpha_2))$. The results imply that this ratio is approximately 50:50. The above share of forward-looking agents might seem unexpectedly high, as compared to Driver *et al.* (2003), who estimate 30 % for the UK and 10 % for the USA.[18]

[16] Despite this potential bias the presented estimation methodology remains the optimum estimator in the sense that it minimises the mean square error. To test the distribution of the disturbances terms ($\varepsilon_t$ and $e_t$), we need to know the NAIRU. Therefore, it is impossible to test ex--ante the "true" distribution of the disturbance terms.

[17] For instance, this means a situation where a change in monetary policy regime (such as introducing inflation targeting) influences transmission from the real economy to prices, or where the share of forward-looking agents is changing over time.

[18] Therefore, we must interpret our results with caution. But Christiano *et al.* (2001) or Beneš and Vávra (2003) derive the New Keynesian Phillips curve from microfoundations using backward indexation and structural inflation persistence, and also demonstrate that the share of forward-looking agents might be close to 50 %.

TABLE 2 Estimation Results

| | Expected inflation ($\alpha_1$) | Lagged inflation ($\alpha_2$) | Unemployment gap ($\beta$) | Exchange rate gap ($\gamma$) | S.E. |
|---|---|---|---|---|---|
| (A) Basic model | 0.486*** (4.090) | 0.460*** (4.344) | –0.629*** (–4.451) | 0.162 (1.140) | 1.782 |
| (B) Pure backward--looking model | – | 0.852*** (17.752) | –0.496*** (–4.832) | 0.045 (0.222) | 2.252 |
| (C) Alternative unemployment model | 0.491*** (3.898) | 0.444*** (3.850) | -0.707*** (–3.015) | 0.127 (0.837) | 1.739 |
| (D) Alternative real exchange rate gap model | 0.456*** (3.835) | 0.478*** (4.252) | –0.508*** (-7.011) | 0.083 (0.545) | 1.970 |
| (E) Alternative import inflation model | 0.494*** (3.335) | 0.468*** (3.766) | –0.506*** (–7.110) | 0.266* (1.962) | 1.929 |

*Notes:* Model (A) uses the core inflation, inflation expectations, import prices, real effective exchange rate and ILO unemployment rate. Model (B) then uses the same variables as model (A), but the coefficient $\alpha_1$ is fixed at zero. In comparison, model (C) uses also the same variables as model (A); only the used rate of unemployment follows the MLSA definition instead of the ILO definition. Finally, in model (D) the real exchange rate gap is derived from the exchange rate against the euro instead of the real effective exchange rate and in model (E) the change in the euro nominal exchange rate is used to catch the "direct" foreign price channel instead of import prices.

*, **, *** denote the significance at the 10%, 5% and 1% level, respectively; *t*-statistic in brackets.

The estimated coefficients $\alpha_1$ and $\alpha_2$ imply small values of the coefficient of import inflation ($1 - \alpha_1 - \alpha_2 = 0.05$). However, this alone does not indicate that the prices of import goods really have such a small impact on inflation, since import inflation also works through inflation expectations.

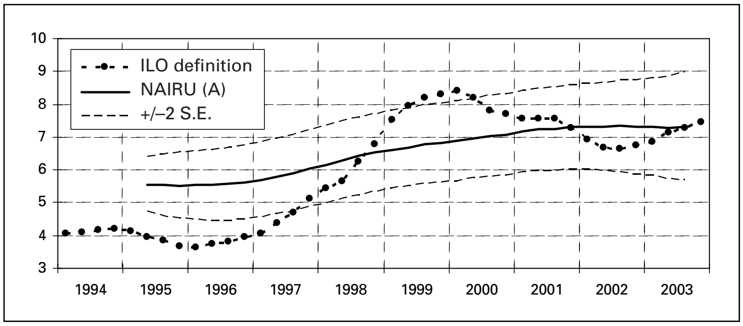
Coefficient $\beta$ measures the elasticity of inflation to the unemployment gap. In our basic model (A), it has an expected negative sign and a reasonable value of –0.6. The influence of the exchange rate gap (coefficient $\gamma$) on inflation is characterised by an elasticity of 0.16. It is, however, insignificant (see the value of the *t*-statistic in Table 2).

*Figure 5* shows the Czech NAIRU (in the definition of our basic model (A) specification and within the band of $\pm$ double the standard error) and the actual unemployment rate. The band represents the 95% confidence interval.

After a period of stability in 1995–1996, the Czech NAIRU started increasing, from 5.5 % in 1996 to approximately 7.5 % in 2003. More precisely, the critical period in which we observe the rise in the NAIRU starts in the last quarter of 1996 and ends in the first quarter of 1999. Since then the NAIRU has remained roughly stable. Conditional on using different methodology and data sets, we can compare our results with other estimates.

Our results are comparable with previous research. Vašíček and Fukač (2000), as well as Fukač (2003) or Hájek and Bezděk (2001), use registered unemployment. Their results suggest increases in NAIRU from 3–4.5 % in 1996 to 8.5–9 % in 2000. Bezděk, Dybczak and Krejdl (2003), as well as Beneš and N'Diaye (2004) adopt the ILO definition of unemployment and show that a 95% confidence interval NAIRU moved from 4 % in 1994 to 7.5 % in 2002. All the estimates at our disposal show the NAIRU growing since 1996.

FIGURE 5 Unemployment Rate and the NAIRU (in %)




In general, the reasons for the increasing NAIRU arise from the situation on the labour market. We should mention labour market regulations, the activity of trade unions, a rising minimum wage, the level of social benefits and the share of labour in GDP. All this supports the idea about worsening labour market performance for institutional reasons, which is expressed by the rising NAIRU.

The increase in the NAIRU implies a proportional slowdown in potential output growth. From a business-cycle point of view, we can identify two periods. First, from 1995 to 1998 actual unemployment was below its equilibrium level, thus there were inflation pressures arising from the labour market. High domestic demand and an undervalued exchange rate had not yet forced firms to restructure and reduce over-employment. Wages grew faster than labour productivity and this was reflected – through private consumption – in pressures on consumer prices.[19]

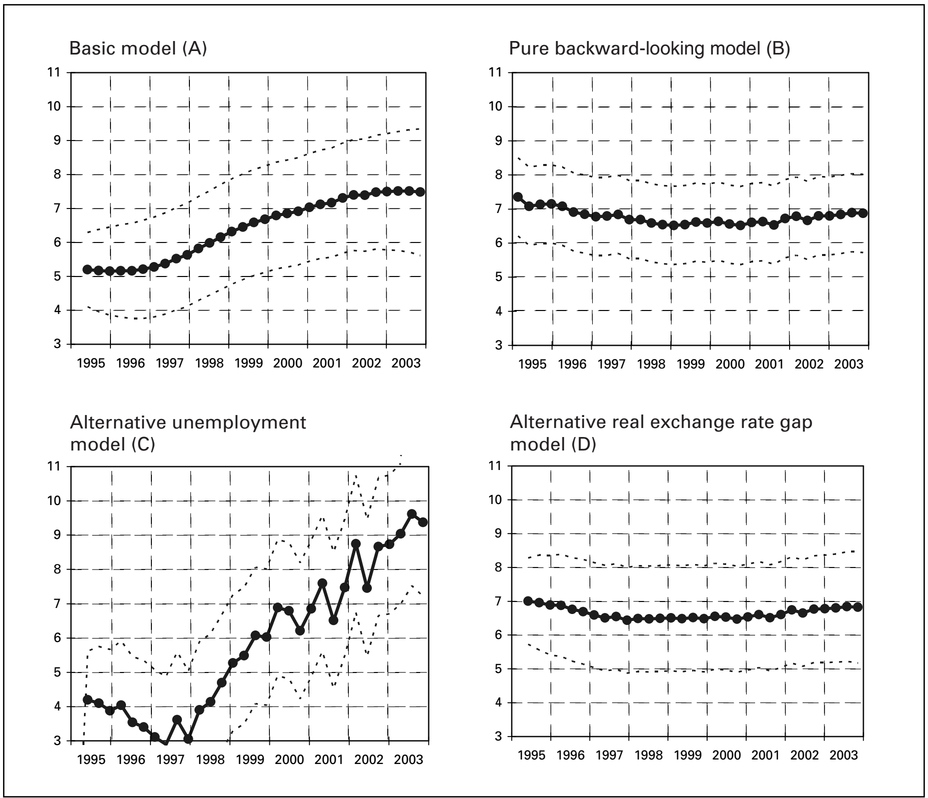
By contrast, the period 1999–2000 was characterised by a positive unemployment gap, i.e. with actual unemployment above the NAIRU. The economy was in recession after an exchange rate shock in 1997, followed by tightening fiscal and monetary stances.

Toward the end of the sample, the level of unemployment is close to the NAIRU. This implies relatively stable conditions on the labour market, meaning that the economy has adjusted to a relatively high level of "equilibrium" unemployment. So there is probably no excessive room for countercyclical policies, but rather for structural policies aimed at reducing the NAIRU.

During the period of a rising NAIRU (1996–1999), the Czech economy experienced both pro- and anti-inflationary shocks. Nominal exchange rate appreciation in the second half of 1996 was followed by a speculative attack on the currency in May 1997. The currency crisis resulted in an exchange rate depreciation and also by the liberalisation of regulated prices. As a result, inflation rose dramatically in the second half of 1997. Then, during 1998 the economy was hit by an opposite shock, namely an external disin-

[19] As Hurník and Navrátil (2002) note, however, this imbalance was reflected also in growing external disequilibrium.

FIGURE 6 Basic and Alternative Estimates of the NAIRU




flationary shock caused mainly by oil and food prices. Although we have tried to eliminate the possible impact of such shocks, our results could still be biased, as the estimates depend critically on the relationship between the unemployment rate and inflation. The high degree of uncertainty surrounding the NAIRU estimates is explicitly demonstrated by a rather wide confidence band, the observed increase in the NAIRU is statistically insignificant at the conventional 5% level. Alternative models also express a relatively high uncertainty with regard to the level of the NAIRU.

The assumption that a specification with pure backward-looking expectations can be downward-biased is confirmed only partially; even though the differences between the NAIRU using the backward-looking and forward-looking specification are primarily negative since 2000, their magnitude is not fundamental (see the differences between the NAIRU based on the basic model (A) and the NAIRU based on the pure backward-looking model (B) in *Figure 6*). This means that the pure backward-looking specification would shift the NAIRU estimates only slightly downward at the end of the sample. On the other hand the pure backward-looking specification shifts upwards the results for the second half of the nineties. However, the reason for such a bias is not certain.

Alternative estimate using the MLSA definition of unemployment (model (C)) gives a very similar result in comparison to the basic model (A). Indeed, as seen in Figure 6, the implied changes in the NAIRU are very close to our basic model. By contrast, the implied NAIRU in the alternative definitions of the real exchange rate gap and import inflation (models (D) and (E)) is more stable, fluctuating in a narrow band around 7 %. One must admit, however, that the statistical significance of these two alternative models is worse than of the basic model (see the standard error statistics in Table 2).

In order to check the robustness of inflation expectations used in our model, we also estimated an alternative model with financial market expectations based on the CNB inflation expectations survey data.[20] This NAIRU estimate is higher and approaches a 10% level whereas the rate of forward-lookingness is lower (20%). However, these results face several problems. First, the time series of survey expectations are available with monthly frequency starting in May 1999. Consequently, there is a need to use the monthly unemployment according to the MLSA data instead of ILO unemployment that is released only quarterly. The rate of unemployment in the MLSA definition is higher than the ILO unemployment over the whole sample, shifting the results upwards. Second, as a high frequency of monthly data requires the use of a richer lag structure, we face the multicolinearity problem. Third, the survey expectations cover year-to-year inflation one year ahead, while in our baseline model we use quarter-to-quarter inflation *one quarter* ahead. Even so, bearing in mind all the above-mentioned problems, the alternative estimation still lies within the confidence interval of our baseline estimation (see Figure 5).

## 6. Conclusion

The NAIRU estimates are based on the relationship between price movements and the position of the economy in the business cycle. Assuming that such a relationship exists in the short run and using a methodology for estimating unobservable variables, we obtain estimates of the NAIRU that indicate that – after stabilizing at around 6 % during 1995–1996 – the NAIRU increased between 1996–1999 by approximately 1.5 percentage points and stabilized thereafter.

The estimated movements of the NAIRU raise the question as to what factors caused these movements. Generally, the reasons for a rising NAIRU appear to reflect the situation on the labour market. In particular we argue that the structure of the labour market had changed substantially in the second half of the 1990s, as the new system of social benefits and the new Labour Act were introduced. Moreover, regular minimum wage increases were adopted also in that period. All these institutional changes could have resulted in a rising NAIRU.

In such a case we could conclude with an indication of worsening aggre-

[20] See (CNB, 2004) for data. We do not report these alternative estimation results here in full, but they are available from the authors upon request.

gate labour market performance and decreasing efficiency. However, the story does not end here, as the critical period of the rise in the NAIRU is connected with several offsetting shocks that hit the economy. As a result, there still is a fair amount of uncertainty surrounding the NAIRU estimates. This opens room for future research.

## REFERENCES

AKERLOF, G. A. – DICKENS, W. T. – PERRY, G. L. (2000): Near-Rational Wage and Price Setting and the Long-Run Phillips Curve. *Brookings Papers on Economic Activity*, 2000, no. 1, pp. 1–60.

AMANO, R. A. – NORDEN, S. (1992): Unit-Root Tests and the Burden of Proof. *Bank of Canada, Working Paper*.

BARRO, R. – TENREYRO, S. (2000): Closed and Open Economy Models of Business Cycles with Marked and Sticky Prices. *NBER Working Paper*, 2000, no. 8043.

BASDEVANT, O. (2003): On Applications of State-Space Modelling in Macroeconomics. *Reserve Bank of New Zealand*, *Discussion Paper*, 2003, no. 2.

BENEŠ, J. – N'DIAYE, P. (2004): A Multivariate Filter for Measuring Potential Output and the NAIRU: Application to The Czech Republic. *IMF Working Paper*, no. 04/45.

BENEŠ, J. – VÁVRA, D. (2003): Nominal Inertia in the Business Cycle: Backward Indexation and Structural Inflation Persistence. Prague, Czech National Bank – mimeo.

BETTS, C. – DEVEREUX, M. (1996): The Exchange Rate in a Model of Pricing to Market. *European Economic Review*, vol. 40, 1996, pp. 1007–1021.

BEZDĚK, V. – DYBCZAK, K. – KREJDL, A. (2003): Czech Fiscal Policy: Introductory Analysis. *Czech National Bank*, *Working Paper*, 2003, no. 7.

BLAKE, A. P. (2002): State-Space Models and the Kalman Filter: Application, Formulation and Estimating. Bank of England, CCBS – mimeo.

BLANCHARD, O. (1999): *European Unemployment: The Role of Shocks and Institutions.* Roma, Lezioni Paolo Baffi, 1999.

CALVO, G. A. (1983): Staggered Prices in a Utility-Maximizing Framework. *Journal of Monetary Economics*, vol. 12, 1983, no. 3, pp. 983–998.

CNB (2004): *Financial Market Inflation Expectations 1999–2004*. Prague, Czech National Bank, http://www.cnb.cz.

DRIVER, R. L. – GREENSLADE, J. V. – PIERSE, R. G. (2003): The Role of Expectations in Estimates of the NAIRU in the United States and the United Kingdom. *Bank of England*, *Working Paper*, no. 180.

ELMESKOV, J. (1993): High and Persistent Unemployment: Assessment of the Problem and Its Causes. Paris, *OECD*, *Working Paper*, no. 132.

FRANZ, W. (2003): Will the (German) NAIRU Please Stand up. *ZEW Discussion Paper*, 2003, no. 35.

FUHRER, J. C. – MOORE, G. R. (1995): Inflation Persistence. *Quarterly Journal of Economics*, vol. 110, 1995, no. 1, pp. 127–159.

FUKAČ, M. (2003): *Non-Accelerating Inflation Rate of Unemployment*. Mimeo.

GALÍ, J. – GERTLER, M. (1999): Inflation Dynamics: A Structural Econometric Analysis. *Journal of Monetary Economics*, 1999, no. 44, pp. 195–222.

GALÍ, J. – MONACELLI, T. (2002): Monetary Policy and Exchange Rate Volatility in a Small Open Economy. *NBER Working Paper*, no. 8905.

GORDON, R. J. (1997): The Time-Varying NAIRU and its Implications for Economic Policy. *Journal of Economic Perspectives*, vol. 11, 1997, pp. 11–32.

GREENSLADE, J. V. – PIERSE, R. G. – SALEHEEN, J. (2003): A Kalman Filter Approach to Estimating the UK NAIRU. *Bank of England*, *Working Paper*, no. 179.

HÁJEK, M. – BEZDĚK, V. (2000): Odhad potenciálního produktu a produkční mezery v ČR. (Estimate of Potential Output and the Output Gap in the Czech Republic) *Czech National Bank*, *Working Paper*, 2000, no. 26.

HAMILTON, J. D. (1994): *Time Series Analysis*. Princeton, Princeton University Press, 1994.

HARVEY, A. C. (1989): *Forecasting, Structural Time Series Models and the Kalman Filter*. Cambridge University Press, 1989.

HARVEY, A. C. (1993): *Time Series Models*. 2nd edition. The MIT Press, 1993.

HURNÍK, J. – NAVRÁTIL, D. (2002): Transmise poptávkových tlaků do inflace a vnější nerovnováhy. (Transmission of Excess Demand into Inflation and External Imbalance – in Czech) In: Varadzim, F. (ed.): *Růstové a cyklické souvislosti vyvoje transformujících se ekonomik*. Ostrava, VŠB-TU, 2002.

HURNÍK, J. – NAVRÁTIL, D. (2005): Potential Output Estimation: A Production Function Approach. *Prague Economic Papers* (forthcoming).

CHRISTIANO, L. J. – EICHENBAUM, M. – EVANS, C. L. (2001): *Nominal Rigidities and the Dynamic Effects of a Shock to Monetary Policy*. www.faculty.econ.nwu.edu/faculty/christiano/research/modeling/papermay10-2001.pdf.

KOTLÁN, V. (2002): Monetary Policy and the Term Spread in a Macro Model of a Small Open Economy. *Czech National Bank*, *Working Paper*, 2002, no. 1.

KRUGMAN, P. (1994): Past and Prospective Causes of High Unemployment. *Federal Reserve Bank of Kansas City Economic Review*, 1994, $4^{th}$ Q, pp. 23–43.

KWIATKOWSKI, D. – PHILLIPS, P. C. B. – SCHMIDT, P. – SHIN, Y. (1992): Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root. *Journal of Econometrics*, 1992, no. 54, pp. 159–178.

McCALLUM, B. T. – NELSON, E. (2001): Monetary Policy for an Open Economy: An Alternative Framework with Optimizing Agents and Sticky Prices. *NBER Working Paper*, no. 8175.

OBSTFELD, M. – ROGOFF, K. (1996): *Foundations of International Macroeconomics*. The MIT Press, 1996.

PHILLIPS, A. W. (1958): The Relation between Unemployment and the Rate of Change of Money Wage Rates in the United Kingdom, 1861–1957. *Economica*, 1958, November, pp. 283–299.

RICHARDSON, P. – BOONE, L. – GIORNO, C. – MEACCI, M. – RAE, D. – TURNER, D. (2000): The Concept, Policy and Measurement of Structural Unemployment: Estimating a Time Varying NAIRU across 21 OECD Countries. *OECD Working Paper*, 2000, no. 250.

STAIGER, D. – STOCK, J. H. – WATSON, M. H. (1996): How Precise are Estimates of the Natural Rate of Unemployment? In: Romer, C. D. – Romer, D. H. (eds.): *Reducing Inflation: Motivation and Strategy*. University of Chicago Press, 1996.

TURNER, D. – BOONE, L. – GIORNO, C. – MEACCI, M. – RAE, D. – RICHARDSON, P. (2001): Estimating the Structural Rate of Unemployment for the OECD Countries. *OECD Economic Studies,* 2001, no. 33.

VAŠÍČEK, O. – FUKAČ, M. (2000): Quantitative Analysis of the Non-Accelerating Inflation Rate of Unemployment and Non-Accelerating Inflation Product. Application for the Czech Republic. *Masaryk University Brno*, *Working Paper*, 2000.

## SUMMARY

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# Labor-Market Performance and Macroeconomic Policy: Time-Varying NAIRU in the Czech Republic

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During the second half of the 1990s, the Czech economy experienced a sharp increase in the unemployment rate. We attempt to determine whether this was caused by structural changes, worsening labour-market performance, or by the changing business-cycle position. This has direct implications for both monetary and fiscal policy. We use NAIRU (non-accelerating inflation rate of unemployment) estimates using time-varying NAIRU. The estimates indicate that the NAIRU increased between 1996 and 2002 by approximately 1.5 percent. Estimated increases in the NAIRU can be associated with the worsening of labour-market efficiency.