The Time Varying NAIRU in the Czech Republic

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Abstrakt

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 estimate NAIRU using time-varying NAIRU. The estimates indicate that the NAIRU increased between 1996 and 2002 by approximately 1.5%. Estimated increases in the NAIRU can be associated with the worsening of labour-market efficiency.

Klíčová slova

non-accelerating inflation rate of unemployment – time varying NAIRU – maximum likelihood methods – forward-looking expectations

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.

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]

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

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

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) 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 example, by Hamilton (1994),[4] 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

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[3] See (Hurník – Navrátil, 2004) for a sensitivity analysis of the potential output growth (based on a production function approach) to NAIRU movements in the Czech Republic.

[4] 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.

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:

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

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:[5] 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 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. 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.

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[5] 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).

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.[6]

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.[7] 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 economic 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 $1 - \alpha_1 - \alpha_2 = 0$.[8] Our final estimated equation is as follows:

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

3. The Data

The data used for the analysis cover the period 1Q1994-Q2004. 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.

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

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

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. It is released quarterly and this determines the use of quarterly data in the present analysis. 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 excludes the direct influences of changes in regulated and energy prices. To approximate the real exchange rate gap, we use the real effective exchange rate (the nominal effective exchange rate deflated by the Czech and effective foreign inflation). 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. To check the robustness of the estimation, we also use a detrended real exchange rate against the euro as 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. Thus we use import price inflation in our basic model.

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. 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.

$$u_t^* = u_{t-1}^* + \varepsilon_t$$

$$e_t \sim \mathrm{N}(0, \sigma_e^2) \,; \varepsilon_t \sim \mathrm{N}(0, \sigma_\varepsilon^2) \tag{4}$$

$$\mathrm{cov}(e_j, \varepsilon_k) = 0$$

The error term ε_t in state equation (4) is expected to be white noise with standard deviation σ_ε. 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 σ_ε. The disturbance vectors e_t (from (3)) and ε_t are assumed to be uncorrelated with each other in all time periods.

If there is no constraint imposed on the standard deviation σ_ε, we shall see high volatility in the estimated NAIRU series, which appear hard to interpret from an economic point of

5. mezinárodní konference Finanční řízení podniku a finančních institucí
VŠB-TU Ostrava, Ekonomická fakulta, katedra Financí

Ostrava
7.-8. září 2005

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 σ_ε, and in such a way limiting the possible fluctuations of the NAIRU directly. This means constraining the standard deviation σ_ε 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 λ 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 $t = 1,2,...,T$; see (5) and (6). The smoothed series and mean squared errors are then computed by backward recursion:

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

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

for $t = T - 1, T - 2,...1$

This method enables us to avoid the problem with the *a priori* (arbitrary) choice of standard deviation in the Gordon fashion, or of the λ coefficient in the Hodrick-Prescott filter. This estimation methodology, however, rests on the assumption that the disturbances ε_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.[9]

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 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.[10] Table 1 contains the estimated coefficients and their *t*-statistics in brackets.

TABLE 1 Estimation Results

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[9] 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 (ε_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.

[10] 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.

	Expected inflation (α_1)	Lagged inflation (α_2)	Unemploy-ment gap (β)	Exchange rate gap (γ)	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 α_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.

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.[11]

The estimated coefficients α_1 and α_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 β 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 γ) on inflation is characterised by an elasticity of 0.16. It is, however, insignificant (see the value of the *t*-statistic in Table 1 above).

Figure 1 shows the Czech NAIRU (in the definition of our basic model (A) specification and within the band of ± double the standard error) and the actual unemployment rate. The band represents the 95% confidence interval.

After a period of stability in 1995-996, 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 3ůů4.5 % in 1996 to 8.5ůů9 % in 2000. Bezděk, Dybczak and Krejdl (2003), as well as Beneš and N´Diaye (2004) adopt the ILO definition of

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[11] 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 %.

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 1 Unemployment Rate and the NAIRU (in %)
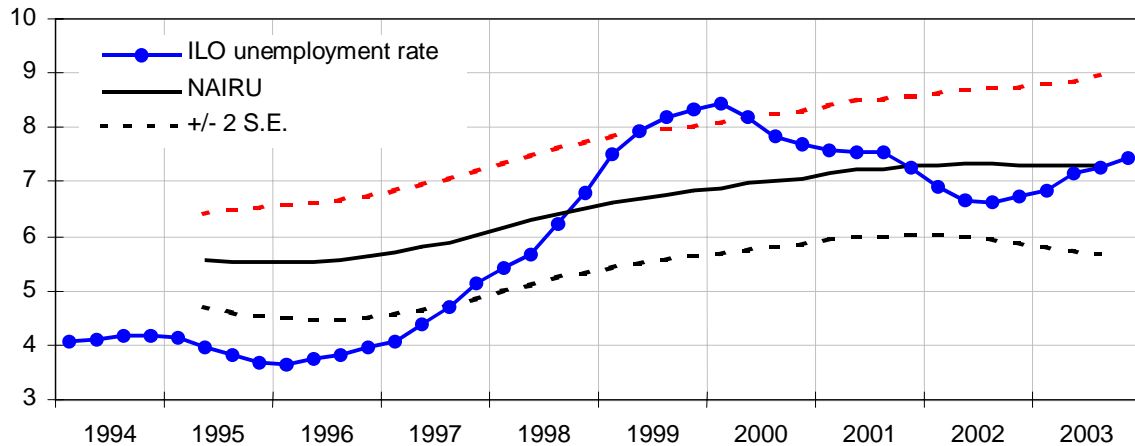
In general, the reasons for the increasing NAIRU arise from the situation on the labour
market. We should mention labour market regulations, the activity of trade unions, a rising
minimum wage, the level of social benefits and the share of labour in GDP. All this supports
the idea about worsening labour 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.[12]

By contrast, the period 1999ůů2000 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 (1996ůů1999), 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 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 disinflationary 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

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

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 *2*). 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 2, 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 1).

FIGURE 2 Basic and alternative estimates of the NAIRU
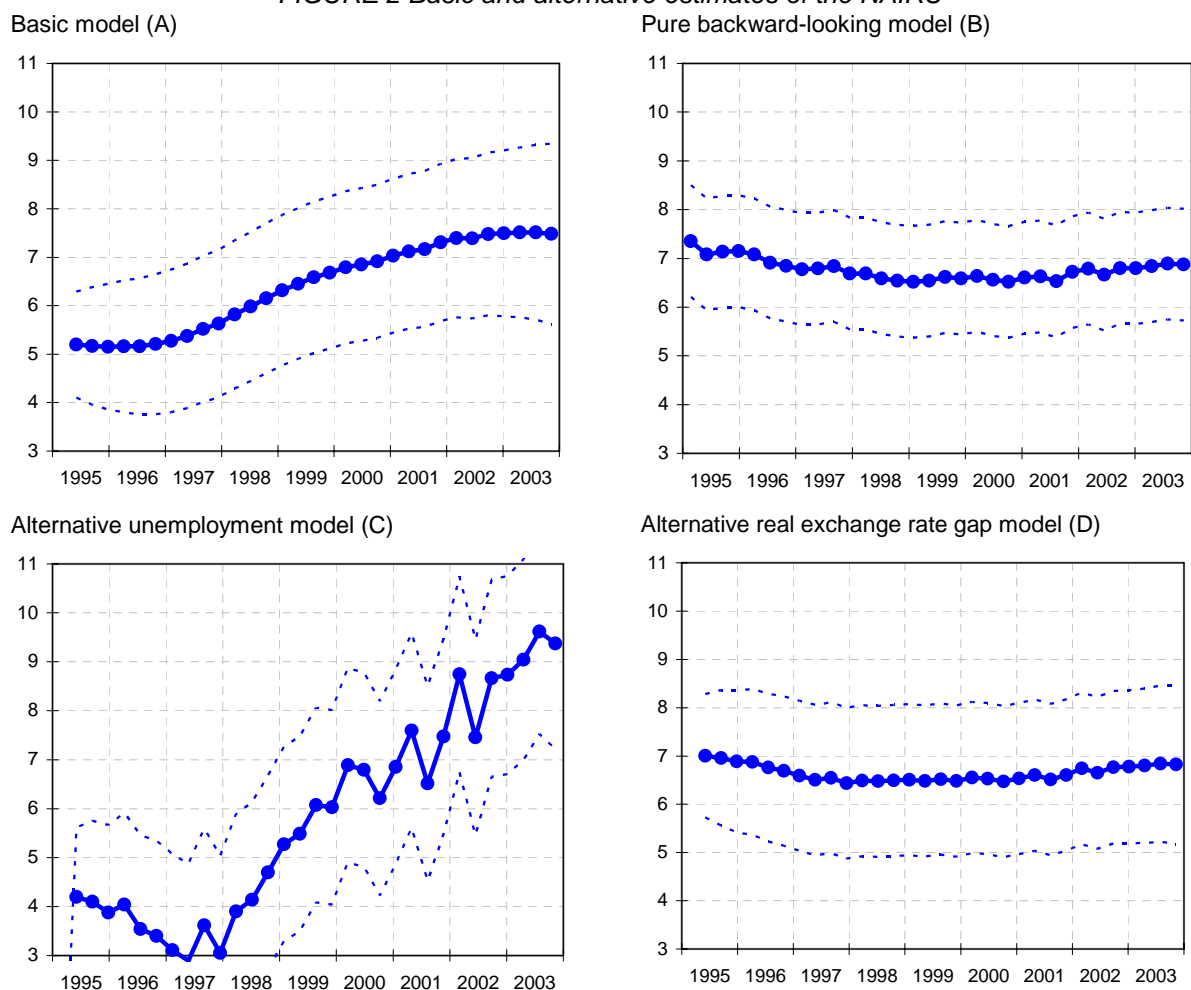
In order to check the robustness of inflation expectations used in our model, we also estimated an alternative model with financial market expectations based on the CNB inflation expectations survey data.[13] This NAIRU estimate is higher and approaches a 10% level whereas the rate of forward-lookingness 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 1).

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 1995ůů1996 – 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 aggregate 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.

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[13] 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.

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Summary

Trh práce v České republice: v čase se měnící NAIRU

Druhá polovina 90. let je spojena s prudkým nárůstem nezaměstnanosti v ČR. Naším cíle bylo ukázat, zda byl tento vývoj způsoben strukturálními změnami, zhoršením fungování trhu práce nebo zda šlo pouze o cyklický vývoj. Toto rozlišení je klíčové pro nastavení fiskální a měnové politiky. Výsledkem je odhad v čase se měnící NAIRU. Podle našich odhadů vzrostla NAIRU v letech 1996 až 2002 cca o 1,5%. Tento nárůst dáváme do souvislosti se zhoršením efektivity trhu práce.