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# Price Convergence to the EU: What Do the 1999 ICP Data Tell Us?

Martin Čihák and Tomáš Holub*

**Abstract**

The paper analyses the price convergence in the Czech Republic and other Central and Eastern European (CEE) countries towards the European Union (EU). Cross-country comparisons based on the International Comparison Project (ICP) 1999 are used. The authors conclude that in a benchmark convergence scenario, the equilibrium real exchange rate appreciation of the Czech koruna (CZK) should reach roughly 1.5–2.0 percent a year. They also warn, however, that there may be additional sources of real appreciation such as terms-of-trade changes or price deregulations, which may lead to a higher pace in the medium run. Studying a more detailed breakdown of commodities, the authors find that no clear distinction can be made between tradable and non-tradable goods, the "degree of non-tradability" varying between 10 and 85 percent. The implications of this for differences in the structures of relative prices in the CEE countries compared with the EU are analysed. The paper concludes that it may take about 15 years for the Czech relative price structure to converge to the least-developed EU countries.

**JEL Codes:** E31, E52, E58, F15, P22.
**Keywords:** Balassa–Samuelson model, inflation, relative prices.

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## Nontechnical summary

The authors, following up on their earlier research based on the International Comparison Program (ICP) conducted by the United Nations/OECD/Eurostat, use the ICP results from 1999 to discuss the price convergence process in the Czech Republic and other Central and Eastern European countries towards the European Union (EU).

At the beginning of the paper, the authors show that based on cross-country comparisons derived from the ICP, an economy that achieves a growth differential compared to the EU of 2 percentage points a year can be expected to show a real exchange rate appreciation of about 1.8 percentage points on average. This "benchmark convergence scenario" is close to the real exchange rate appreciation that some recent studies predict for transition economies. Such a rate of real appreciation would be consistent with fulfilling the Maastricht criteria on inflation and exchange rate stability, at least for those countries that do not use a currency board as their monetary arrangement, if the wide ±15% ERM II band is considered.

The benchmark scenario is analysed in more detail in the next part of the paper. In particular, the authors focus on the fact that the Czech Republic was still about 20 percentage points below the predicted level based on the cross-country comparison of price levels and gross domestic product (GDP) in 1999. A similar finding, derived from previous ICPs, has led some authors to suggest that there might be a risk of price jumps upon the Czech Republic's accession to the EU and/or the European Monetary Union. However, the present authors show that if cross-country differences in employment rates, non-tradable productivity, the size of the non-tradable sector, government policies and the structure of foreign trade are taken into account, the Czech price level in 1999 can be explained well. This finding makes the argument about future price jumps in the Czech Republic less compelling. In fact, the authors argue that after the recent exchange rate appreciation (to about 47 percent of the EU in 2001), the Czech koruna might have become "overvalued" in cross-country comparison, which could justify the central bank's policy steps to reduce the appreciation since mid-2001.

To better use the information contained in the 1999 ICP, the authors in the subsequent part of the paper focus on detailed data about the structures of relative prices. The detailed data show that the difference in the structures of relative prices between the accession countries – including the Czech Republic – and the advanced EU countries is still much bigger than the difference between the advanced EU countries and the least developed EU countries. The analysis confirms that countries with a lower price level tend to have a higher dispersion of relative prices. Also, the authors confirm their earlier results that one cannot strictly distinguish between the tradable and non-tradable commodity groups. Based on the 1999 ICP, they find that the "empirical degree of non-tradability" of the various commodity groups ranges from 10 percent (the most tradable) to 84 percent (the least tradable).

These more detailed findings are used by the authors to determine how prices in the individual groups are likely to develop depending on the rate of real economic convergence. The authors conclude that it may take about 15 years for the Czech aggregate price level to reach 60 percent of the EU price level, and it would take about 15 years to reach the same degree of relative price differences as observed in the least developed European countries. This corresponds to an average

rate of real effective exchange rate appreciation of 2.4–3.2 percent, which is above the "benchmark" estimate based on the aggregate cross-country regressions, but still below the trend observed in the Czech Republic over the past decade.

Another conclusion is that with inflation targets announced by the central bank, a GDP growth rate exceeding 4.7 percent and real exchange rate appreciation higher than 2-3 percent a year would force the prices of the "most tradable" commodity groups to decline, which may 'hit the constraint' of lower downward flexibility of prices. A growth rate of 4.7 percent does not appear to be a binding constraint, since the benchmark convergence scenarios include slower growth for the Czech Republic. However, if factors such as a significant exchange rate appreciation exceeding the benchmark scenario put additional downward pressure on inflation, prices of some commodities may start to decline.

The authors note that smooth price adjustment to the EU level would require companies and labour unions to adjust their behaviour to the circumstances of low inflation, nominal exchange rate appreciation and falling prices of tradable goods. Policymakers can contribute in this respect by communicating more actively to the private sector the implications of the announced inflation targets and the process of real economic convergence. This could help overcome the behavioural aspect of the downward rigidity of prices and its real macroeconomic costs which may stem from the companies' and labour unions' lack of experience with the low-inflation environment, the appreciating nominal exchange rate and the falling prices of tradable goods.

## 1. Introduction

This paper is a follow-up to our earlier work (Holub and Čihák, 2000 and 2001; Čihák and Holub, 2001 a,b) in which we used the International Comparison Program (ICP) data on relative price structures to assess whether sufficient room had been left in the Czech National Bank's (CNB) inflation targets for adjustments in the structures of relative prices. In the present paper, we return to this analysis, using the most recent data from the ICP 1999, published in late 2001.

The aim of the present paper is to identify the long-run growth rates of some important macroeconomic variables which are consistent both with the announced inflation targets and with smooth progress with real economic convergence. Understanding these equilibrium developments may facilitate monetary policy decisions thanks to better calibration of long-run trends (and current deviations from these trends) in the forecasting process. Good communication of these trends to the public may also make it easier to achieve the monetary policy targets if companies and labour unions take these into account in the wage-setting process, the exchange rate assumptions of their business plans, etc. In this case, it might reduce the problem of downward rigidities and thus the potential welfare costs of achieving the inflation targets.

The paper is organised as follows. After this introduction, the second section briefly describes the data used throughout the paper. The third section presents the empirical relationship between GDP and average price levels and its implications for real exchange rate appreciation during the convergence process. The fourth section focuses on relative price structures. We analyse past adjustments in relative prices and discuss their likely future path, including the implications for individual economic sectors. The fifth section concludes.

## 2. Data

The ICP is the main source of international data on relative prices and price levels. It is organised worldwide by the United Nations (see Kravis and others, 1982; Kurabayashi and Sakuma, 1990; Heston and Lipsey, 1999), and in Europe by OECD/Eurostat, in cooperation with the national statistical offices. The key results of the ICP include data on prices in individual countries, both on a highly aggregated level (GDP and its components) as well as on a commodity group level. The price data have the form of "comparable prices". The comparable price of a commodity $i$, denoted as $P_{ij}$, is the price of commodity $i$ in economy $j$ in terms of commodity $i$ in a reference economy (for instance, the price of bread in the Czech Republic in terms of bread in the EU-15). The ICP calculates the $P_{ij}$ s for various commodity groups. Private consumption, for instance, consists of thirty items.

The focus of the present paper is on the 1999 ICP data, released at end-2001. Our previous research, based mainly on the results of the 1996 ICP, led to a number of observations about the data. It appears appropriate to revisit these observations in the light of the 1999 ICP data.

While we will treat the ICP data as given in this paper, it is necessary to express some reservations at the beginning. Even though an ongoing effort has been made by the project's organisers to achieve the highest possible degree of comparability across countries and to adjust the prices for factors such as quality differences, considerable potential scope still exists for errors. For example, in the case of GDP in PPP, the margin of error is sometimes estimated at as much as 5 percentage points (see, for instance, Schreyer, Koechlin, 2002). This means, for example, that no meaningful distinction can be made between the Slovak per capita GDP, which stood at 48 percent of the EU average in 1999, and the Hungarian figure of 50 percent of the EU average. These measurement errors, of course, can cause unexplained residuals in the econometric relationships that we estimate in the subsequent sections, or even bias the estimates themselves. Moreover, as we go from the aggregated level to the disaggregated level, the probability of errors is likely to increase, as the errors may average out at the aggregated level.

Time-series comparisons of individual ICP projects need to be treated with caution, since both the national accounts methodologies and the ICP methodologies have been changing over time. For instance, before 1993, the transition economies were linked to the ICP through bilateral comparisons with Austria only, while since 1996 the comparisons have been multilateral, as is usual for advanced countries. Moreover, the comparisons have been influenced by the gradual introduction of the new system of national accounts, ESA95, which took place at different points of time in different countries (see Stapel, 2002). Finally, a number of smaller but important changes in the ICPs have taken place, such as the way the quality adjustments are calculated for different commodity groups or countries.[1]

The problems with time-series comparisons of the individual ICP projects were illustrated at the aggregate level for example in Schreyer and Koechlin (2002). They showed that the GDP in PPP of Portugal extrapolated from the 1990 ICP to 1999 differs from the actual 1999 ICP results by as much as 6 percentage points. For some transition economies, this difference may be even bigger. At a disaggregated level, we found large differences between the 1999 ICP data and the extrapolated prices of individual commodity groups from 1996 ICP to 1999 which we derived in our earlier research (see, for instance, Čihák and Holub, 2001).[2] These differences are largely attributable to the data collection and processing of the 1996 ICP vs. the 1999 ICP.

Given the disappointing performance of the extrapolations, we decided to reduce their use in this paper compared with our previous work to the minimum necessary for the current policy discussions. We still compare the results of the individual ICP projects, but these comparisons were informed by the above discussion of methodological changes in between the projects.

[1] Authors such as Filer and Hanousek (2000) argued that official measures of inflation in the Czech Republic have not captured well the effects of consumer substitution, outlet substitution, quality improvements and new goods, and have therefore been seriously biased upwards (by about 3 percentage points in 1996–97).

[2] When regressing the actual data on the extrapolated ones, we obtained a coefficient of determination equal to 0.55 only and a slope coefficient significantly below 1 (0.77), which is not what we had expected.

## 3. Price Levels

### 3.1 Basic Empirical Observations

Much of the existing literature focuses on the strikingly close relationship between the average price level and per capita GDP. Figure 1 and equation (1) show, for a pooled sample of European OECD countries[3] and all 13 EU accession countries in 1999, that in cross-country comparison, a strongly significant positive relationship exists between the price level and GDP per capita in purchasing power parities (standard errors are in parentheses):

$$\underset{}{\mu(GDP)} = \underset{(4.71)}{8.02} + \underset{(0.06)}{0.90}\ GDP_{PPP}\,, \qquad (1)$$

$$R^2=0.90,\ N=30,\ F=265.4,$$

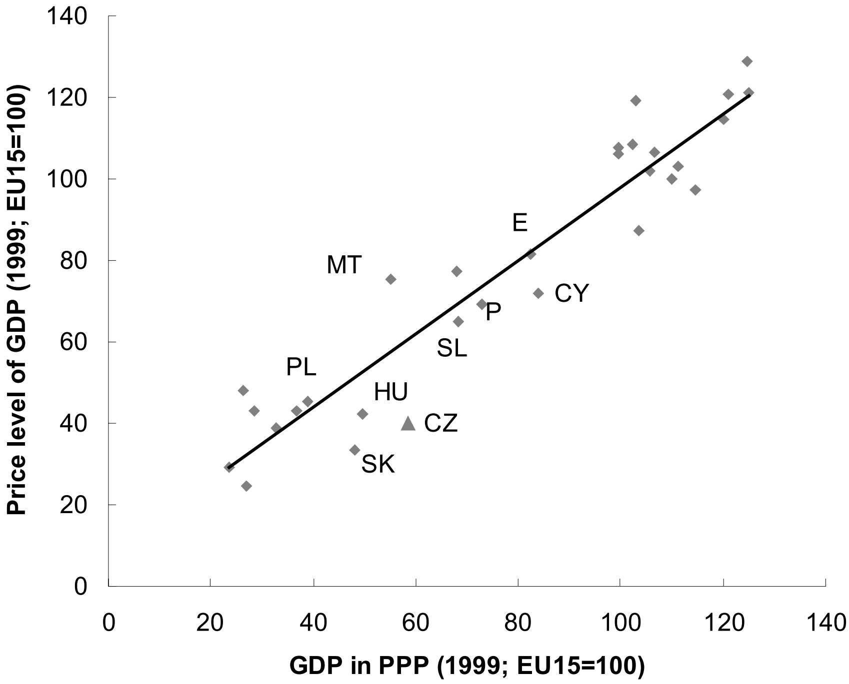
where $\mu(GDP)$ is the average price level of the GDP and $GDP_{PPP}$ is gross domestic product in purchasing power parities (in both cases, EU15=100). The regression, estimated using ordinary least squares (OLS), accounts for 90 percent of the cross-country variability in price levels. The null hypothesis of no correlation between the two variables can be rejected at the 1 percent significance level. The Jarque–Bera test for normality of residuals, as well as the White heteroscedasticity test, did not reveal any problems with using the OLS technique. The estimated slope coefficient is equal to 0.9. This means that an increase of 1 percentage point in GDP per capita in purchasing power parity (PPP) units relative to the EU average tends to be accompanied by an increase of 0.9 percentage points in the price level relative to the EU. As shown in Table 1, this is slightly less (but not significantly so) than the estimates for a narrower sample of 22 countries which we obtained in our earlier papers based on the 1996 ICP data (see, for instance, Holub and Čihák, 2000 and 2001). It is, however, similar to the estimates for an extended sample of 33 countries in 1996 which we used in Čihák and Holub (2001a,b).

As the relationship is linear, we cannot interpret the estimated slope as an elasticity, which is of more interest to us as we would like to see the relationship between real GDP growth and real exchange rate appreciation in percent. In Table 1, we thus provide two versions of the elasticity. The first one is for the Czech Republic's actual GDP and for the fitted price level. It thus shows what "hypothetical" elasticity one could expect if the Czech Republic lay on the estimated regression line and moved along it during its convergence process. The second one uses the Czech actual GDP as well, this time in combination with the actual Czech price level. It thus shows what

[3] Luxembourg was excluded as an outlier. One may rightly ask why we exclude Luxembourg but keep another outlier, i.e. the Czech Republic, in the sample and pay so much attention to explaining its residual. First, unlike the Czech Republic, the inclusion of Luxembourg leads to strong non-normality and heteroscedasticity of the residuals, both confirmed by standard econometric tests at the 1 percent probability level. Second, Luxembourg is not only an outlier, but also a strongly influential point. Including it in the regression changes the estimated slope to 0.75, which is more than two standard errors from the 0.90 estimated in equation (1). On the contrary, skipping the Czech Republic from the regression changes the estimated slope only marginally to about 0.89 (or 0.74 when Luxembourg is included). The special nature of Luxembourg can be explained by the fact that it has an extraordinarily high share of frontier workers, which biases the GDP per capita upwards as a measure of productivity. Arguably, a solution would be to use more appropriate proxies for productivity than GDP in the regression, rather than excluding Luxembourg. We do this in the following text, where we estimate more complex models.

"actual" elasticity there would be if the Czech Republic moved parallel with the estimated line, but with its actual residual in the regression. The two elasticities are 0.87 and 1.32 respectively, the latter one being substantially larger because the Czech Republic has a negative residual in the regression.

***Figure 1: Price Level vs. GDP in PPP (1999, 30 countries)***




***Source:*** *Czech Statistical Office; Eurostat.*
***Note:*** *CY=Cyprus; CZ=Czech Republic; E=Spain; HU=Hungary; MT=Malta; P=Portugal; PL=Poland; SK=Slovakia; SL=Slovenia.*

The results change only marginally when the price level of GDP is replaced in this regression by the price level of household consumption $\mu(cons)$, which is shown in equation (2). Compared with equation (1), the overall explanatory power of the regression is just marginally lower. The normality and heteroscedasticity tests do not show any econometric problems. The estimated slope coefficient is 0.94, while the corresponding "Czech" elasticities are 0.92 and 1.51. This finding is important, because it means that the issues that we are going to discuss below for the price level of GDP apply almost equally to the price level of household consumption, too. Table 1 below illustrates that the results of both (1) and (2) have been reasonably robust over time, the slope coefficient typically being in the interval of 0.85–1.0 and the "hypothetical Czech elasticity" being from 0.8 to 1.1.

$$\mu(cons) = \underset{(5.49)}{4.83} + \underset{(0.06)}{0.94}\ GDP_{PPP}, \qquad (2)$$

$R^2$=0.89, $N$=30, $F$=215.8,

***Table 1: Price Level vs. GDP in PPP, 1993–1999 (2001)***

| ICP year | No. of countries | Reference economy | Inter-cept | Slope coeff. | Standard error | $R^2$ | *Elast* 1) | *Elast* 2) | *CZ* 3) |
|---|---|---|---|---|---|---|---|---|---|
| ***GDP*** | | | | | | | | | |
| 1993 | 29 | Germany | 7.19 | 0.87 | 0.05 | 0.92 | 0.85 | 1.39 | -18 |
| 1996 | 21 | Germany | -3.49 6) | 0.97 | 0.08 | 0.88 | 1.06 | 1.82 | -23 |
| 1999 4) | 22 | Germany | -4.35 6) | 1.00 | 0.07 | 0.91 | 1.09 | 1.54 | -15 |
| 1996 | 33 | Germany | 4.30 6) | 0.88 | 0.05 | 0.91 | 0.92 | 1.68 | -26 |
| 1998 | 106 | Germany | 26.81 | 0.56 | 0.04 | 0.70 | 0.54 | 0.91 | -32 |
| 1998 | 103 | Germany | 23.93 | 0.62 | 0.04 | 0.79 | 0.59 | 1.00 | -24 |
| 1999 | 30 | EU15 | 8.02 | 0.90 | 0.06 | 0.90 | 0.87 | 1.32 | -21 |
| 2001 5) | 22 | EU15 | 13.24 | 0.86 | 0.08 | 0.86 | 0.79 | 1.05 | -16 |
| ***Consumption*** | | | | | | | | | |
| 1993 | 29 | Germany | 2.17 6) | 0.96 | 0.05 | 0.93 | 0.95 | 1.68 | -20 |
| 1996 | 22 | Germany | -1.05 6) | 0.98 | 0.09 | 0.86 | 1.02 | 1.71 | -23 |
| 1999 | 30 | EU15 | 4.83 6) | 0.94 | 0.06 | 0.89 | 0.92 | 1.51 | -23 |

***Note:*** *1) Implied elasticity of the price level with respect to GDP for the actual level of Czech GDP and fitted Czech price level;*

*2) Implied elasticity for the actual level of Czech GDP and the actual Czech price level;*

*3) Difference between the actual and predicted price level in the Czech Republic (in percentage points of the reference country's price level);*

*4) The regression was based on the 1996 ICP data, extrapolated to 1999;*

*5) OECD estimates (extrapolations from 1999);*

*6) NS = not significant at the 10 percent significance level.*

The robustness of the results presented in equation (1) and Figure 1 can be further assessed by increasing the size of the sample. In equation (3) and Figure 2, we increased the sample size from 30 to 106, using the World Bank's 1998 data (Germany=100). The slope coefficient is 0.56 in this case. This estimate must be treated with a good deal of caution, though. There are some clear outliers that are likely to bias the regression. Important problems with using the OLS technique in this case were also indicated by the tests of normality (the Jarque–Bera statistics were significant at the 1 percent level) and heteroscedasticity (the White test statistics were significant at the 10 percent level; the reported t-statistics are therefore White heteroscedasticity-consistent). A robust method of estimation might be thus appropriate to achieve a more reliable estimate. As a step in this direction, we decided to eliminate from the regression all countries with a residual exceeding 30 percentage points in absolute value (the standard error of regression being 14 percentage points), which meant three countries in this case.[4] The results are presented below as equation (3'). In this regression, the coefficient of determination increased to about 80 percent. We now cannot reject the hypotheses that the distribution of the residuals is normal. The problem with heteroscedasticity remains, though. We therefore report the White-consistent standard error for the slope coefficient. Its estimated size rose to 0.62. The implied "hypothetical" elasticity for the

[4] These countries were the Syrian Arab Republic (price level = 98% of Germany; GDP = 12% of Germany), Congo Republic (price level = 68%, GDP = 4%) and Luxembourg (price level = 86%, GDP = 167%).

Czech Republic is 0.54 and 0.59 for the two worldwide estimates, which is substantially below the interval estimated for the narrower (European) sample of countries.

$$\underset{}{\mu(GDP)} = \underset{(2.21)}{26.81} + \underset{(0.05)}{0.56}\ GDP_{PPP}\,, \tag{3}$$

$R^2$=0.70, $N$=106, $F$=240.6,

$$\mu(GDP) = \underset{(1.66)}{23.93} + \underset{(0.04)}{0.62}\ GDP_{PPP}\,, \tag{3'}$$

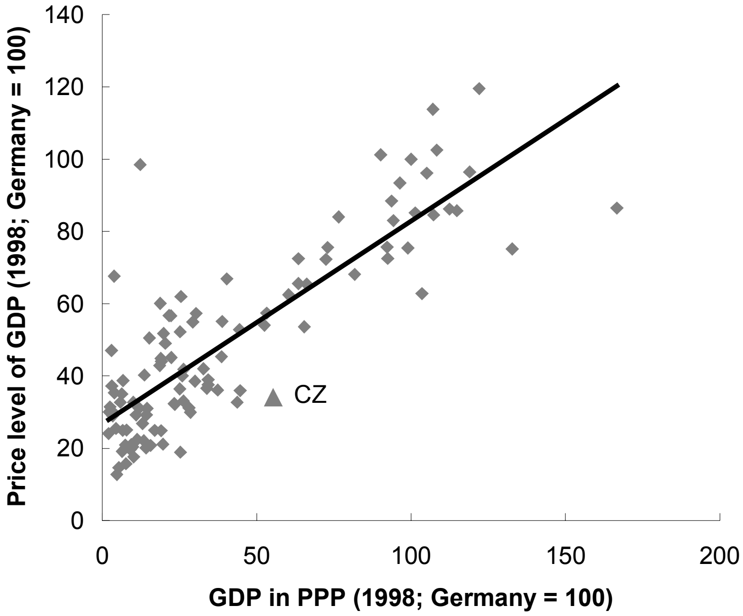
$R^2$=0.79, $N$=103, $F$=383.4,

This finding somewhat reduces the degree of confidence concerning the estimated elasticities, even though the statistical significance of the GDP vs. price level relationship remains very high. In the following text, though, we decided to work with the "European" estimate of the "hypothetical" elasticity, equal to about 0.9 for 1999. This region is more homogenous than the worldwide country sample, and is also more relevant for Czech domestic policy issues (such as the process of convergence towards the EU and the appreciation of the CZK/EUR exchange rate). Moreover, the figure 0.9 is close to the estimated "actual Czech elasticity" for the worldwide sample, which makes us believe that it is not overestimated for the Czech case.

The estimated elasticity can be used to predict the average pace of real exchange rate appreciation for an average catching-up economy starting with a GDP in PPP roughly at the current Czech level. If such an economy achieves a growth differential compared to the EU of, for example, 2 percentage points a year (which we think might be realistic for the Czech Republic),[5] its real exchange rate appreciation should be about 1.80 percentage points on average. Interestingly, this is close to the real exchange rate appreciation that some recent empirical studies predict for transition economies based on the Balassa–Samuelson effect (see, for instance, Halpern and Wyplosz, 2001; Begg et al., 2002; Deutsche Bundesbank, 2001; and the references therein). If we used the estimated "actual" elasticity for the Czech Republic in 1999 (1.3–1.5), the implied real exchange rate appreciation would go up to 2.6–3.0 percent, which is closer (but still below) the observed trend of CPI-based appreciation in the past decade.

[5] This calculation is only approximate, since it assumes a constant growth differential over time. At longer horizons, it would be necessary to take into account the fact that the growth rate in the converging country should decline during the convergence process. At the horizon of 15 years, the assumption of a 2 percentage point growth differential yields the same growth rate as if we had assumed that Czech GDP increases such that the gap vis-à-vis the EU declines by 2.5 percent every year.

***Figure 2: Price Level vs. GDP in PPP (1998, 106 countries)***




***Source:*** *World Bank.*

From the monetary policy point of view, it is important that this pace of real appreciation is slow enough to allow the EU accession countries to fulfil the Maastricht inflation and exchange rate criteria at the same time, if the wide ±15 percent fluctuation band of the ERM II mechanism is considered. If a country plans to spend the shortest required period of two years in the ERM II, the maximum Maastricht-consistent speed of real appreciation is 8–9 percent a year. This is way above any realistic estimate based on our cross-country comparisons, even if we use the "actual" elasticity for the Czech Republic. For countries with currency boards, however, even a speed of 1.8 percentage points might be a problem. In their case, the only channel of real appreciation is inflation, and the Maastricht limit is just 1.5 percentage points above the average inflation in the best three performers among the EU countries. Moreover, the currency board countries are among the CEECs with a relatively lower GDP per capita, and the average speed of their real convergence may thus easily exceed 2 percentage points a year.

When trying to draw some stronger policy conclusions, though, we should keep in mind that in spite of the simple cross-country regressions' good overall fit, there are still important residuals. Traditionally, the Czech Republic is among the countries that lie far below the regression line, i.e. it has a large negative residual in the cross-country comparison. In the regression based on the 1999 ICP data, the price level of the Czech GDP still appears about 20 percentage points lower than predicted by the simple regression (see Table 1). Ideally, we would like to find other explanatory variables for such residuals besides GDP.[6] This is important at least on three policy grounds. First, some authors argue that the large negative residual may signal a danger of price jumps – or at least some acceleration in inflation – after EU/EMU accession (see, for instance, Vintrová et al., 2002). If we can find a reasonable economic explanation for the negative residual,

[6] Remember also that we had to exclude Luxembourg from the regression as a strong influential point. In an extended regression, we would ideally like to include all the available observations.

this hypothesis may be unjustified. A second – and related – question is the interpretation of the recent sharp CZK appreciation. This appreciation brought the Czech price level to about 47 percent of the EU average in 2001. If one believes that the negative residual in 1999 was a sign of currency undervaluation, such a move could be viewed as an equilibrium phenomenon. Moreover, we show in Table 1 above that the Czech residual still remained rather high (roughly 15 percentage points) in 2001, which would mean that there may be scope for further strengthening of the CZK. On the other hand, if the 1999 Czech price level was in line with fundamentals, its recent appreciation might be considered excessive. Third, a better understanding of the residual may lead to a better estimate of the equilibrium real exchange rate path than the above "benchmark price convergence scenario" of 1.8 percent a year.

In the next two subsections, we thus extend the standard "GDP vs. price level" regression. In the first step, we include the variables that fit into the Balassa–Samuelson model (see Balassa, 1964; Samuelson, 1964). In the second step, we go beyond this simple model, testing the statistical significance of some additional factors.

### 3.2 Estimates Consistent with the Balassa–Samuelson Model

The Balassa–Samuelson model assumes that the law of one price holds for tradable commodities, but not for non-tradable ones. It also assumes perfect labour force mobility among sectors within an individual economy, but zero mobility of labour force among the economies. Under these assumptions, it can be shown easily that with a one-factor production function, the comparative price level of a country should be equal to (see Holub and Čihák, 2003)[7]:

$$\frac{P}{P^*} = \left(\frac{A_T}{A_T^*}\right)^{1-\gamma}\left(\frac{A_N^*}{A_N}\right)^{1-\gamma} = \left(\frac{GDP_{nom}}{GDP_{nom}^*}\right)^{1-\gamma}\left(\frac{A_N^*}{A_N}\right)^{1-\gamma} \tag{4}$$

where $P$ is the price level, $A_T$ and $A_N$ are the labour productivities in the tradable and non-tradable sectors respectively, $(1-\gamma)$ is the share of non-tradable goods in GDP and $GDP_{nom}$ is GDP per employee expressed using the nominal exchange rate (as opposed to GDP in PPP). Foreign variables are denoted with an asterisk. Furthermore, it can be shown that the same relationship also holds (with some minor modifications) for more advanced versions of the Balassa–Samuelson model with a two or three-factor production function, if we treat $A_N$ as the conventionally measured labour productivity in the non-tradable sector (see Holub and Čihák, 2003).

We can thus take equation (4) as a theoretically quite robust guide for our empirical analysis if we want to make it consistent with the Balassa–Samuelson model. Based on it, we propose the following changes to the simple empirical analysis of subsection 3.1:

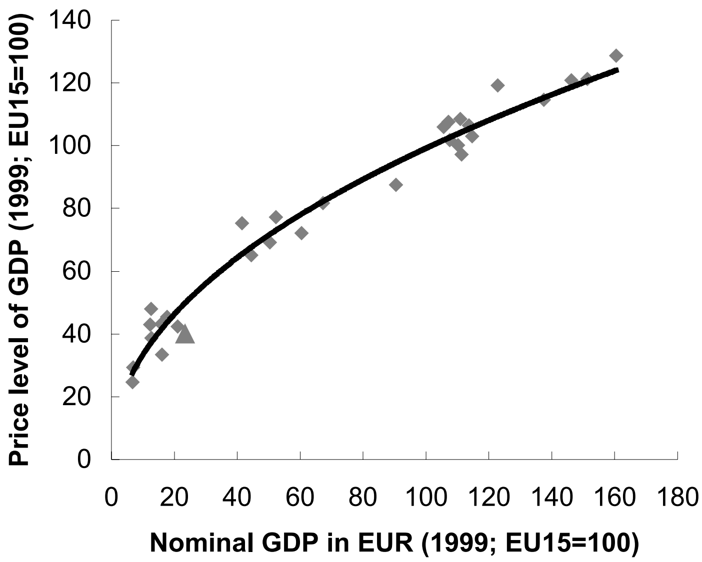
[7] This also assumes a geometric form of the price index. Such a form is optimal provided that we assume a Cobb-Douglas utility function with unitary elasticity of substitution between tradable and non-tradable commodities (see, for instance, Obstfeld and Rogoff, 1998). The traditional arithmetic average can be thought of as a log-linear approximation of this optimal price index (see, for instance, Obstfeld and Rogoff, 1998).

- Nominal GDP vs. GDP in PPP: Equation (4) shows that the relative price level depends on the ratio of nominal GDPs at home and abroad, which approximates labour productivity in the tradable sector under the model's assumptions, the elasticity being ($1-\gamma$), i.e. the share of non-tradables in GDP. A natural approach thus is to regress the relative price levels directly on the ratio of nominal GDPs. Such a regression is shown in equation (5), and also in Figure 3.[8] We can see that the relationship appears to be non-linear (which is in line with equation (4); but it results from the simplifying assumption about the price level index), and the estimated elasticity is 0.47. This means that an increase of 1 percent in nominal relative GDP should be accompanied by an increase of slightly less than 0.5 percent in the relative price level. This can be regarded as a fairly reasonable value. In practice, the share of services (the main group of mostly non-tradable commodities) in GDP varies roughly between 50 and 75 percent in the EU countries and accession economies. The estimated elasticity of 0.47 means that we need to take the lower bound of this interval as an estimate of the share of non-tradables in GDP. This is not unrealistic, since not all services are non-tradable, especially in the modern world.[9] We can thus conclude that the estimate using nominal GDP in equation (5) and Figure 3 is in line with the Balassa–Samuelson model, both qualitatively and quantitatively (see also Čihák and Holub, 2001a,b).

$$\underset{}{log[\mu(GDP)]} = \underset{(0.07)}{2.42} + \underset{(0.02)}{0.47}\, log[GDP_{nom}], \qquad (5)$$

$$R^2=0.96,\ N=30,\ F=668.0,$$

***Figure 3: Price level vs. Nominal GDP (1999, 30 countries)***




***Source:*** *Czech Statistical Office; Eurostat.*
***Note:*** *The black triangle corresponds to the observation for the Czech Republic.*

8 The reported standard error is the White heteroscedasticity-consistent value in this case.

9 Typical examples of tradable services can be found in the financial sector, advisory business and some other modern areas. On the other hand, it is also true that some branches of the secondary sector, such as some parts of the construction or supply network industries, are examples of non-tradable goods.

The use of GDP in nominal terms has certain advantages compared with the oft-used GDP in PPP. One reason is that using GDP in PPP may aggravate the measurement-error problem. If there is, for example, a downward error in measuring the prices of some goods, the measured price level goes down, but at the same time GDP in PPP goes up (because it is calculated as the ratio of nominal GDP to the relative price level). The measurement error thus influences both the dependent and explanatory variables, which biases the estimated slope coefficient. Another reason why GDP in PPP can be considered inferior is the fact that it also magnifies the possible errors due to omitting (or inappropriately approximating) productivity in the non-tradable sector in the regression.

- GDP per capita vs. GDP per employee: In the regressions of the price level on GDP, it is common to use GDP per capita. According to equation (4), nominal GDP can be used as a good proxy for productivity in the tradable sector. However, productivity is measured as GDP *per employee* (or, ideally, as GDP per hour worked) rather than per capita. This can create distortions, as there are important differences among countries in their labour participation rates within the working-age populations, as well as differences in the ratios of their working-age populations to total inhabitants. The ratio of total employment to population ranges from 35 to 55 percent in our sample of countries, and is positively correlated with GDP in PPP. For example in the Czech Republic, this ratio exceeds 45 percent, and exceeds the average of comparable economies (i.e. the Czech Republics lies above the regression line), which means that GDP per capita biases Czech productivity upwards.[10] To eliminate this distortion, we defined the variable *empl%* by dividing the ratio of employment to total population in each economy by the (unweighted) EU average (which is 45 percent). We then divided GDP by *empl%* in the more detailed regression (see below).[11]

- Productivity in the non-tradable sector: As shown in equation (4), by regressing the price level on GDP only, we omit the possible cross-country differences in the productivities of their non-tradable sectors. In particular, we ignore the possibility that a country has a relatively low/high price of non-tradables owing to a high/low productivity in this sector, which reduces/increases unit labour costs in the production of non-tradables, for the given wage level determined in the tradable sector. It is a rather common, textbook approach to assume away such differences, but in practice it might not be justified. We illustrate this in Table 2, which presents Eurostat's estimates of labour productivities in different sectors for the EU members and candidate countries. As we can see, there are significant differences in labour productivities between the candidate countries and the EU in services. In trade, transport, communications, financial and business services etc., the labour productivity of the candidate countries is just 65–70 percent of the EU average (compared to 41 percent in manufacturing).

[10] To illustrate this, let's take the comparison of the Czech Republic with Hungary. The Czech GDP in PPP *per capita* is 59 percent of the EU average, while the Hungarian figure is 50 percent of the EU average. In spite of that, Hungary has a slightly higher price level than the Czech Republic (42 percent, vs. 40 percent in the Czech Republic). When we consider GDP in PPP *per employee*, however, the Hungarian figure was about the same as the Czech one in 1998 (see Table 2), and in fact slightly exceeded the Czech level in 1999. From this point of view, the comparison of the Czech and Hungarian price levels comes as no surprise.

[11] While in theory, such a step should undoubtedly lead to an improvement in the overall regression fit, we must note that in practice it may also lead to some noise if the methodologies for measuring total employment differed among countries.

Moreover, there are examples of countries that have a lower labour productivity in some services branches than in manufacturing (this branch often being public services).

***Table 2: Productivity in Different Sectors (1998; EU15=100)***

| | Agricul-ture | Manu-facturing | Construct-ion | Trade, transport, comm. | Financial, business services | Public services | Total |
|---|---|---|---|---|---|---|---|
| Austria | NA | 115 | 141 | 117 | 140 | 122 | 107 |
| Belgium | 171 | 155 | 153 | 143 | 144 | 112 | 137 |
| Denmark | 164 | 94 | 99 | 104 | 117 | 91 | 97 |
| Germany | 103 | 107 | 107 | 102 | 162 | 106 | 116 |
| Greece | 63 | 55 | 105 | 88 | 130 | 77 | 72 |
| Finland | 122 | 123 | 90 | 116 | 127 | 87 | 105 |
| France | 146 | 120 | 105 | 114 | 122 | 102 | 113 |
| Italy | 106 | 87 | 102 | 119 | 118 | 83 | 98 |
| Luxembourg | 106 | 145 | 108 | 135 | 161 | 154 | 153 |
| Netherlands | 158 | 132 | 107 | 115 | 81 | 95 | 104 |
| Portugal | 54 | 50 | 52 | 61 | 76 | 65 | 55 |
| Spain | 112 | 94 | 96 | 113 | 114 | 91 | 95 |
| Sweden | 143 | 106 | 92 | 99 | 112 | 70 | 91 |
| Bulgaria | 37 | 20 | 29 | 27 | 69 | 21 | 25 |
| Czech Rep. | 88 | 53 | 72 | 63 | 54 | 54 | 58 |
| Estonia | 46 | 26 | 41 | 55 | 44 | 38 | 37 |
| Hungary | 77 | 49 | 54 | 67 | 99 | 52 | 58 |
| Latvia | 12 | 29 | 39 | 42 | 31 | 27 | 27 |
| Lithuania | 26 | 30 | 45 | 45 | 47 | 30 | 30 |
| Poland | 13 | 38 | 68 | 61 | 43 | 49 | 38 |
| Romania | 24 | 31 | 53 | 60 | 77 | 28 | 32 |
| Slovakia | 54 | 42 | 51 | 73 | 107 | 45 | 53 |
| Slovenia | 94 | 58 | 72 | 80 | 82 | 86 | 71 |
| Turkey | 32 | 49 | 54 | 94 | 101 | 54 | 44 |
| CC-13 | 28 | 41 | 58 | 69 | 66 | 47 | 41 |

***Source:*** *Eurostat.*

To take account of these differences, we calculated a proxy variable for productivity in the non-tradable sector (denoted *prodNT*) as a weighted average of the labour productivities reported in Table 2 for construction, trade, transport and communications, financial and business services, and public services. The weights corresponded to the share of these branches in gross value added.[12] We must note that this approach is not without drawbacks. First, the data were available for 23–24 countries only, which reduced the number of observations by almost one third, and thus

[12] In the Czech Republic, the value of *prodNT* was roughly 60 percent of the EU average, and was thus smaller than in some other transition countries such as Hungary (71 percent) and the Slovak Republic (74 percent). This contrasts with our earlier results (Čihák, Holub, 2001a,b), which suggested that Czech productivity in the non-tradable sector was relatively high compared to other transition economies. The difference in the results illustrates the possible scope for error in measuring non-tradable prices and productivities.

limited the degrees of freedom for our extended regression analysis. Second, the data were available only for 1998 and may thus not be fully consistent with the 1999 ICP. And finally, to compute the productivities in PPP, the nominal productivities had to be divided by the estimated price levels in each industrial sector. This – together with the general problem of measuring output in services – may lead to important measurement errors, influencing both the dependent variable and one of the explanatory variables, which could bias the estimates.[13] By including this explanatory variable in the regression, we thus run into the same kind of problem that led us to argue against using GDP in PPP as an explanatory variable instead of nominal GDP (see above).

- Share of non-tradables in GDP: In equation (4), we can see that for any given level of GDP and productivity in the non-tradable sector, the price level depends on the share of non-tradables in GDP, $(1\text{-}\gamma)$. For a country with low productivity in tradables and high productivity in non-tradables, a high share of non-traded goods tends to lower the relative price level and vice versa. If we again define non-tradables as construction, trade, transport and communications, financial and business services, and public services (which might be arguably too broad a definition), their share on GDP varies roughly between 55 and 85 percent in the EU and accession economies (with a statistically significant tendency of richer countries to have a higher share).[14] These are important differences that cannot be ignored.

Combining the three factors discussed above, and in line with equation (4), we estimated the following regression[15]

$$\underset{}{log[\mu(GDP)]} = \underset{(0.32)}{3.74} + \underset{(0.11)}{0.85}\,(1\text{-}\gamma)log[(GDP_{nom})/empl\%] - \underset{(0.19)}{0.61}\,(1\text{-}\gamma)\,log[prodNT], \qquad (6)$$

$$R^2=0.95,\ N=23,\ F=179.74,$$

Both slope coefficients have the expected signs. The fact that they are significantly less than one in absolute terms (while we would expect coefficients equal to one, based on equation (4)) suggests that on average the "true" share of non-tradables in GDP may be smaller than the proxy variable we have used here. This outcome might be realistic. More difficult to explain is the fact that the two estimated coefficients differ from each other (this difference being statistically significant at the 5 percent probability level according to the standard Wald restriction test), while equation (4) predicts them to be the same. The difference in the estimated slope coefficients is most probably a result of using imperfect proxies for the productivities – a problem that is

[13] To give just one example, we may ask if Slovak labour productivity in financial and business services is really twice as high as the Czech one, given the past common history and many similarities between the two countries.

[14] In the Czech Republic, $(1\text{-}\gamma)$ was roughly 65 percent, and was slightly below the regression line, reflecting the Czech industrial tradition. This factor thus can hardly explain the relatively low Czech price level.

[15] In the extended regressions we did not exclude Luxembourg as an outlier. The additional variables are to a large extent able to explain its relatively low price level.

difficult to circumvent.[16] Alternatively, this may signal that the basic Balassa–Samuelson assumptions do not hold (see section 3.3).[17]

The overall fit of the regression is relatively high (95 percent). Nonetheless, the Jarque–Bera test identified non-normality of the residuals at the 5 percent probability level. On closer inspection, we found that the problem was caused by Turkey. This country has a fairly high positive residual in all the previous regressions, as its price level is 48 percent of the EU average, while its GDP in PPP is only 26 percent – and its nominal GDP just 13 percent – of the EU average. Moreover, adding the other explanatory variables made the problem bigger rather than smaller. This may be partly due to the rather specific structure of the Turkish economy (high employment in agriculture, etc.), the recent quite turbulent economic developments, some data problems, and so on. In any case, we decided to present also an alternative version of equation (6) with a dummy for Turkey ($D\_Tur$), even though we had wanted to avoid dropping any countries from the extended regressions.

$$\underset{}{log[\mu(GDP)]} = \underset{(0.24)}{3.96} + \underset{(0.09)}{0.98}(1-\gamma)log[(GDP_{nom})/empl\%] \underset{(0.15)}{-0.80}(1-\gamma)\,log[prodNT] \underset{(0.09)}{+0.39}D\_Tur, \qquad (6')$$

$$R^2=0.97,\ N=23,\ F=230.32$$

This dummy is highly statistically significant, and improves the overall statistical fit of the regression. Even more importantly, the estimate has now passed the normality test, as well as the homoscedasticity test. The estimated slope coefficients are closer to one, which means that our proxy for the size of the non-tradable sector may be better than we had thought based on equation (6). The coefficients, though, remain significantly different from each other.[18] As discussed above, this may reflect problems with our proxy variables or departures of the reality from the simple Balassa–Samuelson model.

In one important respect, both versions of the estimate lead to the same result. If we use them to predict the price level for the Czech Republic, we get 47 percent of the EU average, compared with the actual figure of 40 percent (and 50 percent in the regression given in equation (5)). The

[16] We encountered the same kind of problem (with even greater intensity) in our earlier research (Čihák, Holub, 2001a,b), even though we used a different proxy for productivity in the non-tradable sector. Similar results were also obtained in a panel data analysis of the Balassa–Samuelson model in the Czech Republic by Flek et. al. (2002). When we estimated a restricted version of equation (6) in which we forced the absolute value of the slope coefficients to be the same, the slope was equal to 1.1, which is interesting in the light of the above discussion of the non-tradable sector's share in GDP. The coefficient of determination declined to 0.93 for this regression.

[17] Fischer (2002), for example, presents an extended model with an "investment-demand channel" besides the traditional B–S effect. If some part of the capital is non-tradable, an increase in total factor productivity, which leads to higher investment demand, puts upward pressure on the prices of non-tradables and thus on the real exchange rate. This effect is always positive. The appreciation is particularly strong if tradable productivity goes up and if this sector is capital-intensive. When non-tradable productivity rises, the B–S effect and the investment-demand channel work in the opposite direction, explaining the smaller slope coefficient on non-tradable productivity.

[18] In the restricted version of the equation, the slope coefficient was 1.18 and the coefficient of determination 0.96.

Czech Republic thus still represents an outlier in this analysis, its negative residual being close to (even though slightly below) two standard errors, but a smaller one than in the traditional regression of the relative price level on GDP in PPP. Put in a different way, while the Czech koruna appears to have been "undervalued" by as much as 50 percent in 1999 according to the simple regressions of section 3.1, according to the estimates consistent with the Balassa–Samuelson its "undervaluation" seems to have been around 17 percent only. Incidentally, this corresponds to the size of the real exchange rate appreciation from 1999 till mid-2002. As a result, we might conclude that the Czech koruna does not look "undervalued" any more based on the cross-country estimates of the Balassa–Samuelson model, and no "jump" corrections in the price level could be thus predicted for the future based on this model.

### 3.3 Additional Factors outside the Balassa–Samuelson Model

So far, we have concentrated only on the Balassa–Samuelson approach to explaining price convergence. Our conclusion was that this model performs reasonably well in cross-country comparison, and is broadly consistent with the observed facts both qualitatively and quantitatively.[19] Nevertheless, even such excellent performance may not be good enough for some countries and some policy issues. In particular, the residuals in the regression are still very large for some countries, which means that their price-level convergence process may deviate substantially from the benchmark scenario. As already noted above, the Czech Republic was such a case in 1999. In spite of the better performance in the extended estimates of section 3.2, the Czech Republic still had the largest negative residual, only slightly below two standard errors of regression. Even though this has partly changed since 1999, we are still not sure that the CZK's appreciation was an equilibrium – and thus permanent – phenomenon. This, of course, increases the degree of uncertainty in monetary policy debates.

It is therefore natural to ask if there are other factors determining price levels besides those available within the Balassa–Samuelson framework which would explain at least some portion of the remaining price level differences among countries. We outline some of these possibilities here:

- <u>Government interference:</u> Price levels can be distorted by various government actions, such as price regulation, taxes and subsidies. Perhaps the simplest explanation is the hypothesis that regulation of certain "socially sensitive" prices in the Czech Republic has been more meticulous than in other similar countries. In the next section we focus in more detail on regulated prices when we study price convergence on a more disaggregated level. In this section, we wanted to focus more on the role of taxes, subsidies, agricultural policy, etc. In order to approximate the fiscal influences, we included the share of general government revenues in GDP (*gov*) and/or the ratio of indirect tax revenues to GDP in the regression. These two variables, however, were statistically insignificant and we omitted them from the final regression.[20] As a result, the only proxy variable for government actions that eventually

[19] Except, perhaps, for the difference in the (absolute value of the) estimated slope coefficients for nominal GDP and labour productivity in the non-tradable sector (see section 3.2).

[20] The ratio of indirect tax revenues to GDP even had a negative estimated coefficient, contrary to economic intuition. We obtained a similar result concerning the two public-finance variables in our previous work – see, for instance, Čihák, Holub (2001a,b).

remained in the estimate is the size of agricultural employment in each country (*wagremp*). This variable should roughly approximate the political temptation to regulate/subsidise the agricultural sector. The expected sign of the regression coefficient is positive in this case.[21]

- <u>Imperfect mobility of labour across sectors, and sector-specific human capital:</u> While the Balassa–Samuelson model assumes perfect mobility of labour across sectors, a vast body of theoretical and empirical literature supports the notion of persisting wage differentials across industries and/or sectors (see, for instance, Dickens and Katz, 1987; Gibbons and Katz, 1989; Katz and Summers, 1989; Helwege, 1992; Fallick, 1993; Neal, 1995; and Strauss, 1998). The theoretical part of the literature explains wage differentials using efficiency wage models (based on morale, turnover or shirking), search theory, unionisation and segmented labour markets. The empirical studies show that there are large and persistent wage differentials among industries, even after controlling for a wide variety of worker and job characteristics. Katz and Summers (1989), for instance, present numerous empirical results concerning the presence and pattern of large inter-industry wage differentials and find that workers in high-wage industries earn substantial rents, which impedes labour mobility. Strauss (1998) finds that the assumptions of the Balassa–Samuelson model concerning wage equalisation and competitive labour markets are violated in most industrialised countries. Significant wage and unit labour cost differentials exist both across industries and between traded and nontraded sectors, and these differentials do not diminish over time. The author concludes that these findings are consistent with the presence of industry and/or sector-specific human capital. Helwege (1992) also finds significant evidence of industry-specific human capital and demonstrates that random wage shocks create industry wage variation that persists for substantial periods of time. In general, the literature suggests that workers are compensated for some skills specific to their industry or line of work, and that these skills may constitute an important component of the typical worker's human capital. Imperfect mobility of labour and sector-specific human capital may be a powerful explanation of the residuals in the Balassa–Samuelson model. But it would probably be very difficult to analyse these factors empirically. We did not find any suitable proxies, and thus leave this as a potential challenge for future research.

- <u>Terms-of-trade changes:</u> Finally, and perhaps most importantly, the law of one price may not hold even for tradable goods, contrary to the Balassa–Samuelson assumptions. Much of the literature examining the law of one price has shown that the law fails, particularly in the short to medium run (Mark, 1990, or Engel, 1993). If the law of one price does not hold, price convergence may be partly explained by changes in the terms of trade in tradable goods. In Čihák and Holub (2001a,b), we have suggested that the low price level in the Czech Republic might reflect the large share of higher-value-added products in Czech exports. For these products, perceived quality usually plays a more important role than price. If Czech exports were still perceived as "low-quality goods from the East", Czech producers would have to compensate for the lower perceived quality with lower prices. And the lower the price

[21] We also tried to use a dummy variable to capture the distortionary effect of the EU's agricultural policy (=1 for EU countries; =0 otherwise). Its coefficient was positive, but not statistically significant. We decided to drop this dummy from the regression, since most of the high-income countries in the sample are EU members, and including the dummy would be de facto equivalent to giving up all these important observations, which could substantially distort the explanatory power of the cross-country differences in GDP.

elasticity, the bigger the compensation has to be. In order to estimate the impact of this factor, we included in the regression the share of exports of SITC groups 6, 7 and 8 in the total exports of the individual economies (*exp6_8*).[22] The relationship between this variable and the price level cannot be expected to be monotonic, though, since underdeveloped economies must undercut their prices, while "luxury" products from advanced countries can enjoy a monopolistic premium. Therefore, we also included in the regression the product of this export share and the logarithmic nominal GDP per employee of the individual countries ($exp6\_8*log[GDP_{nom}/empl\%]$).

***Table 3: Regression Including Factors outside Balassa–Samuelson Framework***

| **Dependent variable:** $log[\mu(GDP)]$ | | | |
|---|---|---|---|
| **Number of countries included: 23** | | | |
| **Variable** | **Coefficient** | **Standard error** | **P-value** |
| *Constant* | 4.19 | 0.30 | 0.00 |
| $(1-\gamma)log[GDP_{nom}/empl\%]$ | 0.55 | 0.18 | 0.01 |
| $(1-\gamma)\ log[prodNT]$ | -0.48 | 0.16 | 0.01 |
| *wagremp* | 0.84 | 0.18 | 0.00 |
| *exp6_8* | -1.70 | 0.59 | 0.01 |
| $exp6\_8*log[GDP_{nom}/empl\%]$ | 0.41 | 0.13 | 0.01 |
| R2 | 0.98 | | |
| R2 adjusted | 0.98 | | |
| F-statistics | 194.34 | | 0.00 |

The resulting regression is presented in Table 3. Overall, the explanatory power of this regression is high, reaching 98 percent. All the included explanatory variables are highly statistically significant and their coefficients have the expected signs. The residuals of the regression have passed the normality and homoscedasticity tests.[23]

The foreign trade variables are highly statistically significant. The estimate suggests that the relationship between *exp6_8* and the price level is non-monotonic: for economies with nominal GDP below 65 percent of the EU average a larger share of industrial exports leads to a lower average price level, and vice versa.[24] This supports the hypothesis that less-developed countries have to cope with less favourable prices, if they are to export higher-value-added products. The fact that the structure of foreign trade has an important influence on the price level speaks against the Balassa–Samuelson simplifying assumption that the law of one price holds perfectly for tradable goods. As a result, it is necessary to take into account terms-of-trade changes as part of the price convergence process. This is in line with time-series evidence for the Czech Republic,

[22] In Čihák and Holub (2001a,b) we used the share of machinery exports only (i.e. SITC 7). This proxy variable also performed reasonably well here, but the variable that we chose eventually proved to give slightly better results.

[23] Note that Turkey was included in this regression, but did not cause the same problem with normality of residuals as in section 3.2. It did, however, have one of the largest positive residuals.

[24] In Čihák, Holub (2001a,b) we found that this threshold was around 80 percent of the EU average, but this was in terms of GDP in PPP. This is roughly equivalent to nominal GDP of about 65 percent of the EU average.

for example the appreciating trend of the CZK's real effective exchange rate on a PPI basis[25], improved terms of trade in some commodity groups[26], etc.

The estimated coefficients on nominal GDP per employee and non-tradable productivity per employee in Table 3 are now smaller than in equations (6) and (6'), and are close to each other, which is in line with the Balassa–Samuelson model. Their size, if interpreted directly, would suggest that only around 50–55 percent of services are non-tradable, implying that a 1 percent increase in relative nominal GDP would tend to increase the relative price level on average by 0.4 percent only (i.e. by less than in equation (5)). However, an increase in nominal GDP per employee also raises the variable ($exp6\_8*log[GDP_{nom}/empl\%]$). As a result, a 1 percent increase in relative nominal GDP tends to increase – ceteris paribus – the relative price level by 0.68 percent, which is above what equations (5) and (6) would imply, but very close to the implications of equation (6'). The major difference here is that the regression of Table 3 tends to indirectly ascribe part (about two fifths) of this increase to changes in the terms-of-trade conditions rather than to the pure Balassa–Samuelson productivity effect. All these findings highlight the importance of focusing on all channels of the equilibrium real appreciation in monetary policy discussions, and not concentrating on the Balassa–Samuelson effect alone. The Czech experience, for example, has been instructive in this respect, as the long-run trend of real effective exchange rate appreciation (around 4% a year on a CPI basis) has substantially exceeded the estimated size of the Balassa–Samuelson effect (typically 1–2% at most; see, for instance, Flek et al., 2002).

The extended regression in Table 3 can explain the Czech price level reasonably well. The negative residual of the Czech Republic is just 1.3 percentage points, and lies well within two standard errors of regression.[27] This favourable result stems from the fact that the extended estimate: (i) takes into account the lower Czech per-employee GDP in comparison with per-capita GDP, (ii) uses nominal GDP rather than GDP in PPP to approximate Czech productivity in the tradable sector, which leads to a less favourable picture, (iii) indirectly explains the relatively low Czech food prices (see section 4) with a smaller share of agricultural employment, leading to a less interventionist agricultural policy, and (iv) takes into account the relatively large share of Czech industrial exports, which may necessitate price undercutting if the country's products are still partly perceived as inferior. This finding presents a rather strong challenge to those economists who warn against a price jump in the Czech Republic after the EU/EMU accession due to the large Czech residual in the simple regression of the price level on GDP in PPP. In fact, given the rather strong real exchange rate appreciation between 1999 and 2001 (to about 47% of the EU level), the CZK might have become "overvalued" in cross-country comparison, justifying the CNB's policy steps to reduce the exchange rate appreciation since late 2001.

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[25] Compared with 1996, the PPI-based real effective exchange rate appreciated by 17–18 percent.

[26] For example in the SITC 7 group (machinery and transport equipment), the Czech terms of trade improved by almost 20 percent between 1996 and 1999, contributing substantially to a reduced trade deficit (and a surplus at present) in this category of goods.

[27] Luxembourg, another outlier to the downside, was also within the two-standard-error confidence band (its residual being –7.5 percentage points). The biggest negative outlier in the regression was now Italy, with a negative residual of about –12 percentage points (while the biggest positive residuals were obtained for Germany, Estonia and Turkey).

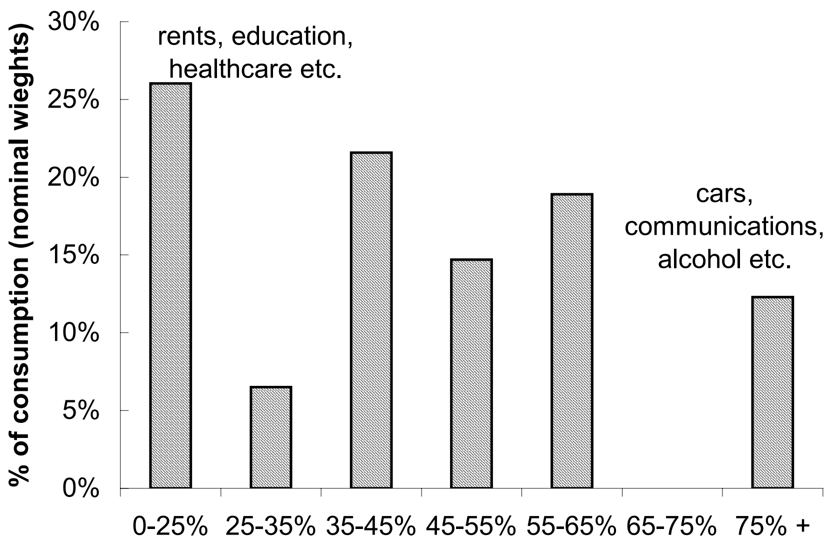
# 4. Relative Prices

## 4.1 Differences in Relative Prices According to the 1999 ICP

The international price differences do not concern average price levels only, but also – and perhaps even more importantly – relative prices. The differences in the structures of relative prices in the Czech Republic and the EU countries according to the 1999 ICP are illustrated in Figure 4. Figure 4 is a histogram showing the distribution of comparable prices of individual commodity groups in the Czech Republic in relation to the average of the 15 EU countries. In the Czech Republic, rents, schooling, health care and transport services are at less than 20 percent of the EU price level, while communications, cars and alcoholic beverages have virtually the same price as in the EU.

The wide dispersion of the *comparable* prices in Figure 4 confirms our conclusions based on earlier ICPs that *relative* prices (such as the price of cars in terms of rents) are significantly different in the Czech Republic and in the EU. Comparable prices, $P_i$, do not allow for a direct computation of relative prices. However, the choice of a reference country (such as the EU average in Figure 4) can help us to compare to what extent the structure of relative prices of a given country is similar to the system of relative prices in the reference country.

***Figure 4: Distribution of Comparable Prices, Czech Republic, 1999***




***Source:*** *Czech Statistical Office.*

## 4.2 Changes in Structures of Relative Prices Over Time

Table 4 shows how the structure of comparable prices according to the 1999 ICP differs from the price structures identified in the previous ICP projects. In this particular table, Germany is chosen as the benchmark country (we will discuss later how the conclusions change if other countries are used as benchmarks). The table shows that while the comparable price of household consumption has remained virtually unchanged at about 35 percent of the German level between 1996 and 1999, prices of most commodity groups have undergone substantial changes. The most remarkable developments have taken place in prices of communications, which increased from 26 to 83 percent of German prices from 1996 to 1999. The comparable price of alcoholic beverages and tobacco, on the other hand, has decreased as a percentage of German prices in the same period. While recognising the methodological changes in ICP projects across time, we might conclude that substantial changes in the structures of relative prices have taken place.

***Table 4: Differences in Prices; the Czech Republic vs. Germany (Germany=100)***

| | **1993** | **1996** | **1999** |
|---|---|---|---|
| **HOUSEHOLD CONSUMPTION** | 25.8 | 34.7 | 34.6 |
| *of which:* Foodstuffs, non-alcoholic beverages* | *36.6* | *46.8* | 48.3 |
| Alcoholic beverages, tobacco* | *46.2* | *63.0* | 47.2 |
| Garments and shoes | 36.0 | 48.3 | 54.5 |
| Rent, fuel, energy | 11.4 | 18.8 | 20.0 |
| Housing equipment and maintenance | 38.5 | 58.5 | 53.3 |
| Health and medical care | 18.5 | 14.0 | 22.7 |
| Transportation* | *42.0* | *62.4* | 58.8 |
| Communications | 31.4 | 26.4 | 83.3 |
| Recreation and culture* | *32.0* | *41.6* | 41.9 |
| Education | 13.0 | 10.9 | 15.0 |
| Restaurants, cafes, hotels | 26.9 | 31.1 | 47.2 |
| Other goods and services | 27.0 | 36.2 | 35.6 |
| **GDP TOTAL** | 25.6 | 31.7 | 37.6 |

***Sources:*** *Czech Statistical Office, OECD; *) authors' computations for 1993 and 1996.*

In order to quantify the changes in the structures of relative prices over time and the differences across countries, it is useful to summarise the extent of (and changes in) relative price differences in a single number. We have therefore defined a *coefficient of relative price differences* (Holub and Čihák (2000; 2001) and Čihák and Holub (2001a,b)), which captures the degree of deviation of relative prices in a given country vis-à-vis the system of relative prices in the reference country. This coefficient is calculated as a weighted standard deviation of the comparable price levels of individual goods in the given country relative to the average comparable price level, that is,

$$\rho = \frac{1}{\mu}\sqrt{\sum_i w_i (P_i - \mu\ )^2}\ , \qquad (7)$$

where $w_i$ are the weights of the individual commodities in the consumption basket and $\mu$ is the average price level of consumption. If, for instance, the structures of relative prices in the given country and the reference country were identical, all comparable prices would be the same (that is, equal to $\mu$), and the coefficient of relative price differences would be at its minimum value of zero. The higher the differences in relative prices, that is, the higher the dispersion of comparable prices around their average, the higher the value of $\rho$. Theoretically, the coefficient of relative price differences is not limited from above, but, as shown in Tables 5–6, the empirical values for European countries tend to be well below 1.[28]

There are two basic ways of calculating $\rho$, depending on the choice of weights of individual commodities in the consumption basket. First, we can use the actual structure of nominal consumption expenditures by households ("nominal weights"). The nominal weights might appear a natural way of calculating the weights, but they tend to underestimate the importance of items with artificially low (regulated) prices, such as rents, thereby biasing downward the coefficient of relative price differences in transition economies. Second, we can calculate the "real weights" of commodities by imputing their internationally comparable prices. The calculation of "real weights" is somewhat artificial, since it implicitly assumes that the real structure of consumption in transition countries would not change with changes in relative prices. It is, however, likely that if the relative price of a commodity increases, people will try to substitute its consumption with other commodities, so that the structure of consumption will tend to shift towards those commodities whose prices grow the least.29 Real weights are thus likely to overestimate the actual extent of relative price distortions. In this paper, we calculate the results both for real weights and for nominal weights, which allows us to assess a range of likely values and scenarios.

Tables 5 and 6 below summarise our calculations of the coefficients of relative price differences for a sample of 31 countries in 1980 through 1999, with Germany used as the reference country. The calculations are based on the standard breakdown to 29 (30 for 1999) commodity groups of private consumption used in the ICP. The results with real weights are not qualitatively different from those for nominal weights. In general, the coefficients of relative price differences in the "core" EU countries are well below 0.20, and even in the least developed EU countries they are usually not higher than 0.35. The coefficients in transition countries are in most cases much higher, typically above 0.50.

The coefficients of relative price differences might be expected to decline over time in the transition countries as these gradually converge to the EU in terms of both their real GDP and overall price level. We used this assumption in our previous research (see Holub and Čihák, 2000; 2001, and Čihák and Holub, 2001a,b) to estimate the potential consequences of the relative price convergence process for the Czech price convergence. It could be argued that countries such as Spain and Portugal, which joined the EU in the 1980s, have gone through a substantial price adjustment since their accession, and their coefficients of relative price dispersion are much lower than in all the transition countries.

As shown in Tables 5 and 6, though, we cannot observe such a general tendency in the accession countries over the 1993–99 period. Actually, the non-weighted average of the coefficient with real weights for the transition countries declined just marginally from 0.65 to 0.64 between 1993 and 1999, and the average coefficient with nominal weights in fact increased from 0.53 to 0.56 over the same period. For both types of coefficient, an increase was recorded in 1996 that was then (partly or fully) corrected in 1999, even though the pattern was not the same in all countries. For

[28] For a discussion of the advantages and disadvantages of the coefficient from the theoretical perspective, see Holub and Čihák (2003).

[29] At the same time, though, there will also be income effects: the growth of wealth is likely to increase the share of consumption of luxury items.

example, the Czech Republic's coefficients declined in 1996 but went up in 1999. At the same time, the coefficient has also increased in the EU countries. The average coefficient of relative price differences for the current EU15 countries (excluding Germany, which serves as the benchmark) went up in 1996, and this increase was then only partially corrected in 1999.

***Table 5: Coefficients of Relative Price Differences; Real Weights (Germany=benchmark)***

| | **1980** | **1985** | **1990** | **1993** | **1996** | **1999** |
|---|---|---|---|---|---|---|
| Austria | 0.23 | 0.16 | 0.16 | 0.13 | 0.19 | 0.15 |
| Belgium | 0.18 | 0.21 | 0.16 | 0.13 | 0.12 | 0.15 |
| Bulgaria | NA | NA | NA | 0.55 | 0.76 | 0.77 |
| Cyprus | NA | NA | NA | NA | NA | 0.35 |
| Czech Rep. (Czechoslovakia) | NA | NA | 0.51 | 0.59 | 0.57 | 0.64 |
| Denmark | 0.25 | 0.22 | 0.22 | 0.17 | 0.30 | 0.21 |
| Estonia | NA | NA | NA | 0.67 | 0.81 | 0.64 |
| Finland | NA | 0.34 | 0.31 | 0.24 | 0.25 | 0.21 |
| France | 0.17 | 0.16 | 0.19 | 0.16 | 0.12 | 0.15 |
| Greece | 0.49 | 0.35 | 0.37 | 0.31 | 0.35 | 0.33 |
| Hungary | 0.63 | 0.63 | 0.56 | 0.45 | 0.59 | 0.60 |
| Ireland | 0.34 | 0.40 | 0.29 | 0.26 | 0.32 | 0.24 |
| Iceland | NA | NA | 0.37 | 0.36 | 0.37 | 0.32 |
| Italy | 0.33 | 0.30 | 0.27 | 0.23 | 0.32 | 0.29 |
| Latvia | NA | NA | NA | 0.72 | 1.02 | 0.94 |
| Lithuania | NA | NA | NA | 1.02 | 0.90 | 0.78 |
| Luxembourg | 0.31 | 0.20 | 0.18 | 0.17 | 0.16 | 0.15 |
| Malta | NA | NA | NA | NA | NA | 0.38 |
| Netherlands | 0.20 | 0.15 | 0.15 | 0.12 | 0.12 | 0.18 |
| Norway | 0.37 | 0.33 | 0.29 | 0.26 | 0.30 | 0.30 |
| Poland | 0.63 | 0.77 | 0.60 | 0.68 | 0.83 | 0.57 |
| Portugal | 0.40 | 0.67 | 0.46 | 0.39 | 0.38 | 0.42 |
| Romania | NA | NA | 0.52 | 0.72 | 0.91 | 0.56 |
| Russia | NA | NA | 0.69 | 0.89 | 0.81 | NA |
| Slovakia | NA | NA | NA | 0.73 | 0.86 | 0.68 |
| Slovenia | NA | NA | NA | 0.38 | 0.42 | 0.35 |
| Spain | 0.15 | 0.44 | 0.29 | 0.21 | 0.25 | 0.23 |
| Sweden | NA | 0.29 | 0.25 | 0.15 | 0.19 | 0.22 |
| Switzerland | NA | NA | 0.19 | 0.14 | 0.14 | 0.17 |
| Turkey | NA | 0.48 | 0.51 | 0.44 | 0.57 | 0.58 |
| United Kingdom | 0.23 | 0.24 | 0.22 | 0.18 | 0.27 | 0.21 |
| Transition countries average | NA | NA | NA | 0.65 | 0.77 | 0.65 |
| EU15 average (excl. Germany) | NA | 0.30 | 0.25 | 0.20 | 0.24 | 0.22 |

***Sources:*** *Czech Statistical Office, OECD, own computations.*

***Table 6: Coefficients of Relative Price Differences; Nominal Weights (Germany=benchmark)***

| | 1980 | 1985 | 1990 | 1993 | 1996 | 1999 |
|---|---|---|---|---|---|---|
| Austria | 0.22 | 0.16 | 0.16 | 0.13 | 0.17 | 0.15 |
| Belgium | 0.18 | 0.20 | 0.15 | 0.13 | 0.12 | 0.15 |
| Bulgaria | NA | NA | NA | 0.52 | 0.79 | 0.72 |
| Cyprus | NA | NA | NA | NA | NA | 0.35 |
| Czech Rep. (Czechoslovakia) | NA | NA | 0.53 | 0.53 | 0.48 | 0.52 |
| Denmark | 0.24 | 0.23 | 0.23 | 0.18 | 0.28 | 0.21 |
| Estonia | NA | NA | NA | 0.59 | 0.63 | 0.56 |
| Finland | NA | 0.39 | 0.40 | 0.27 | 0.27 | 0.22 |
| France | 0.17 | 0.16 | 0.18 | 0.16 | 0.12 | 0.14 |
| Greece | 0.42 | 0.34 | 0.40 | 0.31 | 0.31 | 0.30 |
| Hungary | 0.48 | 0.48 | 0.52 | 0.42 | 0.49 | 0.53 |
| Ireland | 0.30 | 0.36 | 0.28 | 0.25 | 0.35 | 0.24 |
| Iceland | NA | NA | 0.43 | 0.38 | 0.36 | 0.32 |
| Italy | 0.30 | 0.25 | 0.23 | 0.20 | 0.27 | 0.24 |
| Latvia | NA | NA | NA | 0.48 | 0.63 | 0.72 |
| Lithuania | NA | NA | NA | 0.66 | 0.66 | 0.64 |
| Luxembourg | 0.38 | 0.18 | 0.17 | 0.15 | 0.15 | 0.14 |
| Malta | NA | NA | NA | NA | NA | 0.39 |
| Netherlands | 0.21 | 0.15 | 0.14 | 0.12 | 0.12 | 0.19 |
| Norway | 0.32 | 0.38 | 0.34 | 0.28 | 0.32 | 0.30 |
| Poland | 0.51 | 0.64 | 0.62 | 0.58 | 0.64 | 0.51 |
| Portugal | 0.49 | 0.41 | 0.44 | 0.35 | 0.34 | 0.33 |
| Romania | NA | NA | 0.74 | 0.53 | 0.64 | 0.57 |
| Russia (USSR) | NA | NA | 0.84 | 0.66 | 0.57 | NA |
| Slovakia | NA | NA | NA | 0.58 | 0.46 | 0.52 |
| Slovenia | NA | NA | NA | 0.38 | 0.35 | 0.33 |
| Spain | 0.14 | 0.35 | 0.27 | 0.20 | 0.23 | 0.20 |
| Sweden | NA | 0.34 | 0.29 | 0.15 | 0.22 | 0.24 |
| Switzerland | NA | NA | 0.20 | 0.14 | 0.13 | 0.16 |
| Turkey | NA | 0.47 | 0.65 | 0.48 | 0.46 | 0.51 |
| United Kingdom | 0.24 | 0.24 | 0.23 | 0.18 | 0.29 | 0.22 |
| Transition countries average | NA | NA | NA | 0.53 | 0.58 | 0.56 |
| EU15 average (excl. Germany) | NA | 0.27 | 0.26 | 0.20 | 0.23 | 0.21 |

***Sources:*** *Czech Statistical Office, OECD, own computations.*

There are several explanations of why the coefficient of relative price differences in transition countries, as well as in the EU, did not decline monotonously in 1993–99:

- The methodology of the ICP changes over time, which makes time-series comparisons not fully appropriate.
- All the countries are being compared with Germany, which at the beginning of 1993 was subject to a significant shock, namely the unification with East Germany, which partially influenced the structure of relative prices in Germany as a whole. However, the results do not

change substantially if another "core" EU country, such as France, is chosen as the benchmark instead of Germany. Even with France as the benchmark, the conclusion still remains valid that the coefficient of relative price differences for the Czech Republic declined between 1993 and 1996, but increased between 1996 and 1999.[30]

- Between 1993 and 1996, prices in the EU countries were influenced by the impact of the EMS crisis and the forced devaluation of several currencies (Italy and the United Kingdom).
- If the downward flexibility of prices in individual countries differs, the economic slowdown in 1996 might have led to an increase in the coefficient of relative price differences compared to Germany.

The price system in the new EU member countries (Austria and Sweden) might have been temporarily disturbed by the preparations for EU accession and by accession itself. A similar development was observed before EU accession in Spain and Portugal, too, where the coefficient of relative price differences temporarily increased in 1985 (in Portugal this was true only for the real-weights-based coefficient). This development – somewhat paradoxical at first sight – can be explained: the acceptance of common EU policies (such as the common agricultural policy) or tax harmonisation can increase some groups of prices in less advanced countries, thereby distorting the system of relative prices temporarily; it is only afterwards that relative prices begin to "settle down" and that the coefficient of relative price differences starts to decrease again.

### 4.3 Relative Prices vs. Price Levels Across Countries

The 1999 ICP data are consistent with our earlier hypothesis that lower differences in relative prices are typically related to a higher aggregate price level. The negative relationship between relative price differences and price levels can be shown in Figure 5, which plots price levels in individual countries against the coefficient of relative price differences. The data in Figure 5 can be fitted by a standard linear regression. For Figure 5a (real weights), we obtain the following result:

$$\underset{}{\mu} = \underset{(0.06)}{1.24} - \underset{(0.13)}{1.26}\rho, \qquad (8a)$$

$$R^2=0.77,\ N=29,\ F=87.96\ .$$

For Figure 5b (nominal weights), we get:

$$\mu = \underset{(0.06)}{1.30} - \underset{(0.17)}{1.58}\rho, \qquad (8b)$$

$$R^2=0.76,\ N=29,\ F=85.27,$$

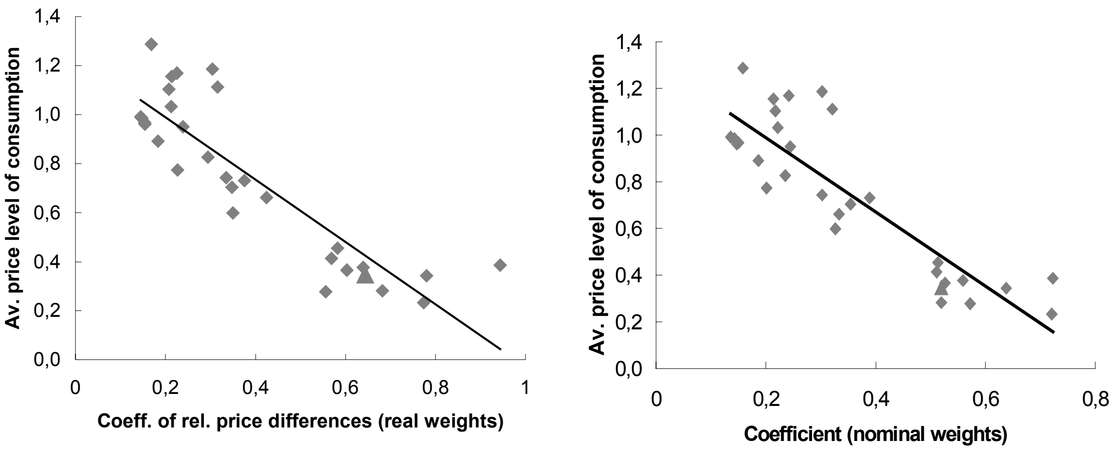
where $\mu$ is the average price level of consumption and $\rho$ is the coefficient of relative price differences, defined in (31). This shows that the coefficient of relative price differences is

[30] For the sake of brevity, the detailed results for France as the benchmark economy are not reported here, but are readily available from the authors.

negatively related to the overall price level. All the coefficients and both regressions are significant at the 1 percent significance level.[31]

***Figure 5: Coefficient of Relative Price Differences vs. Price Level***

(a) Real Weights (b) Nominal Weights




***Sources**:ICP and authors' calculations.*

Table 7 compares the above estimates to those that we obtained previously for ICP 1993 and 1996 and earlier comparison projects since the early 1980s (see Čihák and Holub, 2001). The comparison suggests that these estimates are reasonably stable over time. This stability is quite remarkable, considering that the first half of the observations covers the period before the start of the economic transition in CEE countries, and the second part includes three observations corresponding to three different stages of the transition process.

The stability of these cross-country estimates suggests that the increases in relative price differences found in the previous sub-section may be only temporary. The results in Table 7 mean that transition countries with similar relative price structures to Germany tend to have price levels closer to Germany (i.e. higher). The stability of this relationship gives some credibility to the hypothesis that in the longer term, with output convergence, relative price differences can be expected to decline, while aggregate price levels would increase. We quantify these relationships in greater detail in the next subsection.

[31] Figure 5 suggests that the relationship between $\mu$ and $\rho$ might also be well fitted by a hyperbola rather than a straight line. Indeed, the relationship between $\mu$ and $\rho$ can be estimated as $\mu = 0.27\rho^{-0.80}$ (for real weights) and $\mu = 0.24\rho^{-0.87}$ (for nominal weights), with a corresponding $R^2$ of 0.79 and 0.76 respectively, i.e. similar or slightly higher determination coefficients compared with the linear function. However, given that the "hyperbolic" function does not add much new insight or precision and the notation becomes more cumbersome, we hold on to the assumption of a linear function.

***Table 7: Regression of the Price Level on the Coefficient of Relative Price Differences***

| | Intercept | Slope | St. error | $R^2$ | $p$-value |
|---|---|---|---|---|---|
| **Nominal weights** | | | | | |
| 1980 | 1.17 | -1.35 | 0.34 | 0.56 | 0.001 |
| 1985 | 1.08 | -0.75 | 0.43 | 0.17 | 0.099 |
| 1990 | 1.22 | -1.12 | 0.34 | 0.37 | 0.004 |
| 1993 | 1.21 | -1.61 | 0.17 | 0.81 | 0.000 |
| 1996 | 1.26 | -1.57 | 0.26 | 0.64 | 0.000 |
| 1999 | 1.30 | -1.58 | 0.17 | 0.76 | 0.000 |
| **Real weights** | | | | | |
| 1980 | 1.13 | -1.13 | 0.25 | 0.62 | 0.001 |
| 1985 | 1.15 | -0.88 | 0.25 | 0.45 | 0.003 |
| 1990 | 1.34 | -1.54 | 0.34 | 0.54 | 0.000 |
| 1993 | 1.13 | -1.23 | 0.13 | 0.81 | 0.000 |
| 1996 | 1.19 | -1.17 | 0.16 | 0.73 | 0.000 |
| 1999 | 1.24 | -1.26 | 0.13 | 0.77 | 0.000 |

## 4.4 Macroeconomic Developments and Relative Price Adjustments

We have shown that the structures of relative prices in Central and Eastern European countries are very different from those in the European Union, and that according to the 1999 ICP, the differences even increased between 1996 and 1999. In our previous research (see Holub and Čihák, 2000 and 2001; Čihák and Holub, 2001a,b), we argued that the future adjustment of the structures of relative prices may push towards a higher inflation rate in CEE economies compared with the EU, assuming that the prices are asymmetrically downward-sticky. We used this to discuss the implications for the appropriate choice of inflation targets in the Czech Republic. At present, though, the inflation targets are publicly announced in the Czech Republic as well as in many other transition countries, thus clearly outlining the disinflation process before EMU accession. It is unlikely that these targets could be revised (at least not upwards), and we would not advocate such a step as it would hurt the hard-gained credibility of the new inflation-targeting regimes in Central Europe. Therefore, it is reasonable to turn our earlier questions around and ask what the existing inflation targets mean in combination with the real convergence process for individual prices. And if some of the prices are likely to be forced to decline in nominal terms over the convergence process, the relevant question now is how the potential downward-stickiness problem could be reduced, e.g. by an appropriate wage bargaining mechanism, so as to minimise the negative consequences for the real economy.

In order to answer these questions, we need to establish empirically the sensitivity of individual prices with respect to GDP. This requires running regressions between prices and GDP similar to those shown in Figures 1 and 2 and in the estimates (1), (2) and (3), but separately for each commodity group. The explanatory variable is still GDP in PPP or nominal GDP, but the dependent variable is not the general price level, but the price of a specific commodity group. For a tradable commodity, one can expect that the slope coefficient will not be significantly different from zero, while for a nontradable commodity, it will be.[32]

***Table 8: Price vs. GDP Regressions for 30 Commodity Groups***

| | GDP in PPP | | | | | | GDP in EUR* | |
|---|---|---|---|---|---|---|---|---|
| Commodity group | Slope '99 | t-stat. '99 | R2 '99 | CR '99 | Slope '96 | CR '96 | Slope '99 | R2 '99 |
| Bread and cereals | 0.97 | 10.26 | 0.79 | -27 | 0.98 | -31 | 0.49 | 0.85 |
| Meat | 0.93 | 9.40 | 0.76 | -27 | 1.09 | -18 | 0.38 | 0.79 |
| Fish | 0.47 | 5.77 | 0.54 | -28 | 0.69 | -20 | 0.19 | 0.58 |
| Milk, cheese and eggs | 0.65 | 6.80 | 0.62 | -31 | 0.99 | -28 | 0.26 | 0.67 |
| Oils and fats | 0.54 | 5.50 | 0.52 | -23 | 0.70 | -18 | 0.22 | 0.59 |
| Fruits, vegetables, potatoes | 0.86 | 8.89 | 0.74 | -22 | 0.87 | -15 | 0.36 | 0.84 |
| Other food | 0.64 | 6.94 | 0.63 | -28 | 0.57 | -30 | 0.24 | 0.70 |
| Non-alcoholic beverages | 0.59 | 6.74 | 0.62 | -21 | 0.83 | -27 | 0.23 | 0.70 |
| Alcoholic beverages | 0.58 | 2.57 | 0.19 | -20 | 0.95 | -45 | 0.20 | 0.27 |
| Tobacco | 1.18 | 7.78 | 0.68 | -18 | 1.29 | -16 | 0.58 | 0.87 |
| Clothing including repairs | 0.61 | 7.56 | 0.67 | -14 | 0.72 | -25 | 0.33 | 0.72 |
| Footwear including repairs | 0.96 | 11.22 | 0.82 | -6 | 0.51 | 0 | 0.42 | 0.77 |
| Rentals for housing | 1.09 | 5.92 | 0.56 | -38 | 1.16 | -17 | 0.76 | 0.75 |
| Maintenance, household services | 1.23 | 8.17 | 0.70 | -28 | … | … | 0.54 | 0.74 |
| Electricity, gas and other fuels | 0.80 | 7.89 | 0.69 | +6 | 0.90 | -25 | 0.48 | 0.76 |
| Furniture, floor coverings, textiles | 0.59 | 9.04 | 0.75 | +1 | 0.37 | -3 | 0.31 | 0.82 |
| Household appliances and repairs | 0.96 | 11.69 | 0.83 | -30 | 0.50 | -13 | 0.44 | 0.86 |
| Other household goods and services | 0.82 | 9.64 | 0.77 | -16 | 0.81 | -21 | 0.38 | 0.84 |
| Medical products and equipment | 0.59 | 4.66 | 0.44 | -31 | … | … | 0.23 | 0.40 |
| Medical services | 1.21 | 12.26 | 0.84 | -31 | 1.19 | -34 | 0.80 | 0.88 |
| Personal transport equipment | 0.31 | 2.79 | 0.22 | -16 | 0.21 | -20 | 0.10 | 0.24 |
| Operation of transport equipment | 0.53 | 7.84 | 0.69 | -5 | 0.94 | -12 | 0.23 | 0.73 |
| Purchased transport services | 1.20 | 9.19 | 0.75 | -36 | 1.04 | -18 | 0.69 | 0.84 |
| Communication | 0.57 | 2.71 | 0.21 | +9 | 0.57 | -30 | 0.35 | 0.35 |
| Recreational equipment and repairs | 0.66 | 9.15 | 0.75 | -14 | 0.41 | -18 | 0.30 | 0.77 |
| Recreational and cultural services | 0.97 | 13.96 | 0.87 | -23 | 1.35 | -30 | 0.57 | 0.93 |
| Newspapers, books and stationery | 0.92 | 5.59 | 0.53 | -44 | 1.23 | -33 | 0.39 | 0.60 |
| Education | 1.19 | 14.91 | 0.89 | -24 | 1.08 | -26 | 0.84 | 0.97 |
| Restaurants and hotels | 0.86 | 6.95 | 0.63 | -29 | 1.12 | -39 | 0.39 | 0.78 |
| Miscellaneous goods and services | 0.93 | 13.21 | 0.86 | -22 | 0.92 | -27 | 0.50 | 0.92 |

***Note:*** ** Logarithmic form of regression for nominal GDP in EUR; CR denotes the Czech residual (in percentage points). The labels '99 and '96 refer respectively to our current results based on the 1999 ICP and to the 1996 ICP reported in Čihák and Holub (2001a).*

[32] In theory, tradability could be analysed more directly, by investigating whether foreign competition participates in a particular market. In practice, such an analysis is impaired by data constraints and by the difficulty of defining precisely "foreign competition in a particular market". This leads us to the indirect method described above.

The results of these regressions for 30 commodity groups in 30 European OECD or EU-accession countries are summarised in Table 8. The results seem to confirm that there are many different "degrees of non-tradability" of the various commodity groups, as measured by the slope coefficient in the regression between their prices and GDP in EUR, ranging from 0.10 (the most tradable) to 0.84 (the least tradable).[33] None of the commodity groups included in the ICP can be characterised as purely "tradable," since the relationship between the price and GDP both in PPP and in EUR was positive and statistically significant (and in most cases highly so) in all groups. This finding can be explained by the fact that retail prices of even the most "tradable" commodities include nontradable elements, such as transportation costs, wholesale margins and retail margins. Also, there may be systematic differences in the terms of trade in tradable goods between more and less developed countries (see section 3.3).

Table 8 also compares the results of the commodity group regressions with results derived previously for the 1996 ICP (Čihák and Holub, 2001a). The comparison with the 1996 ICP shows that the structure of the slope coefficients has been reasonably stable, as there have not been many changes in the ordering of the groups according to the slope coefficients, though there have been many changes in the actual estimated values. Also the residual of the Czech Republic has been changing in both directions for different commodity groups; however, in general, there was a trend of decline in the Czech residuals. Importantly, the negative residuals in prices of energy and communications vanished (and actually became slightly positive), which is mostly a reflection of increases in regulated prices between 1996 and 1999 (and in the case of communications also a result of a decline in advanced countries). In the same period, the residual of the Czech Republic in rentals for housing has become more negative, reflecting the persistence of rigid price controls in this area.

The estimated regressions in Table 8 can be used to determine how prices in these individual groups are likely to develop in the future depending on the speed of real economic convergence. If $\pi$ is the domestic inflation rate, $\varepsilon_i$ is the estimated elasticity for commodity group $i$, and $\bar{\varepsilon}$ is the average elasticity for the overall consumption basket, then the domestic price of commodity group $i$ (denoted $P_i^D$) should follow the following expression:

$$\frac{\Delta P_i^D}{P_i^D} = \pi + \left(\varepsilon_i - \bar{\varepsilon}\right)\frac{\Delta GDP_{PPP}}{GDP_{PPP}}, \qquad (9)$$

As an example, let us consider the price of personal transport equipment, which has the lowest elasticity in the regression on GDP in PPP (around 0.18 for a country with the Czech GDP in PPP). The average elasticity for the overall consumption basket was estimated in equation (2) and Table 1 as $\bar{\varepsilon}$=0.92.[34] This means that for personal transport equipment, the term in the brackets on the right-hand side of equation (9) approximately equals –0.74. Assuming that inflation is on

[33] Recall from equation (4) that the slope coefficient in the regression of price levels (individual prices in this case) on nominal GDP is equal to the share of nontradables in GDP (in the individual goods). Theoretically, the slope coefficient would be 0 for a perfectly tradable commodity and 1 for a perfectly nontradable commodity.

[34] We consider the estimated "theoretical" elasticities here, which means that we ignore the Czech residuals for the moment. We discuss what would happen if we took these into account in the next paragraph (see also footnote 36).

the CNB's target of about 3.5 percent,[35] prices of personal transport equipment will stagnate in nominal terms in the Czech Republic if domestic GDP growth is about 4.7 percent a year. For a GDP growth rate higher than 4.7 percent, prices in this commodity group would be forced to decline, which may 'hit the constraint' of lower downward flexibility of prices.

A growth rate of 4.7 percent does not appear to be a binding constraint, as the benchmark convergence scenarios include a slower growth rate for the Czech Republic (see section 3.1). However, the above result needs to be treated with a degree of caution. First, at a more disaggregated level, it might be possible to find commodities with a smaller elasticity than the 0.18 for personal transport equipment, which would decrease the figure for the GDP growth that does not force some prices to decline. Second, the above calculations count too much on the average estimated relationships in the simplest versions of the estimates. As we have shown in section 3, there are additional factors besides GDP growth which may influence the average price level, and thus also the real exchange rate appreciation and changes in relative prices. Note that the GDP growth rate of 4.7 percent should on average be associated with a real appreciation of about 2-2.5 percentage points, while the Czech speed of real appreciation has been historically above that level with a GDP growth rate well below 4.7 percent. It is thus quite possible that in the Czech case, some prices may be forced to decline in nominal terms with a growth rate smaller than 4.7 percent.[36] Finally, the actual exchange rate, economic growth and inflation developments may deviate substantially from their long-run trends in the short or medium run. This may force a decline in some nominal prices in these periods even though there is no need for them to decline in the long run. The Czech Republic, for example, has been experiencing such a situation since late 2001.

***Table 9: Inflation Rate and the "Binding" GDP Growth Rate (percent)***

| Inflation rate | GDP rate "forcing" declines in some prices |
|---|---|
| 0.0 | 0.0 |
| 0.5 | 0.7 |
| 1.0 | 1.4 |
| 1.5 | 2.0 |
| 2.0 | 2.7 |
| 2.5 | 3.4 |
| 3.0 | 4.1 |
| 3.5 | 4.7 |

Table 9 illustrates these reservations by showing that if the Czech inflation rate is – for any of the above reasons – below the 3.5 percent target, GDP growth can become "binding" (i.e. some prices may have to decline) at much lower levels than 4.7 percent. The more likely candidates for price

[35] The targets are defined as a continuous declining band starting at 3–5 percent in January 2002 and ending at 2–4 percent in December 2005. We take 3.5 percent as the average value for the whole 4-year period.

[36] When we use the estimated "empirical" elasticities for the Czech Republic, reflecting the Czech residuals in the simple relationships and thus indirectly the other factors of price convergence besides the GDP growth, the critical pace of the GDP growth turns out to be 2.7 % only, i.e. below the realistic potential.

declines are those commodity groups for which the price-GDP slope coefficients, as reported in Table 8, are below average. Besides personal transport equipment, this includes commodities such as fish, furniture, floor coverings, textiles, medical products, communications and recreational equipment.

***Table 10: Speed of Adjustment, Czech Republic vs. EU15***
***(Number of years, unless otherwise noted)***

| Scenario No. | 1 | 2 | 3 | 4 | 5 | 6 |
|---|---|---|---|---|---|---|
| Percent of the GDP gap closed every year: | 2.5 | 2.5 | 2.5 | 1.5 | 1.5 | 1.5 |
| Percent of the "price vs. GDP" residuals closed every year [1)]: | 2.5 | 5.0 | 10.0 | 2.5 | 5.0 | 10.0 |
| Real effective exchange rate (% change per year, 5 yrs) [2)] | 3.2 | 4.2 | 5.7 | 2.5 | 3.4 | 5.1 |
| Real effective exchange rate (% change per year, 15 yrs) [2)] | 2.5 | 3.1 | 3.6 | 2.0 | 2.6 | 3.2 |
| $\mu$ reaches 60% of EU | 17 | 13 | 9 | 23 | 16 | 11 |
| $\mu$ reaches 80% of EU | 46 | 36 | 30 | 64 | 52 | 48 |
| $\rho$ reaches 40% [3)] | 7 | 5 | 4 | 9 | 7 | 4 |
| $\rho$ reaches 30% [3)] | 19 | 14 | 10 | 25 | 18 | 13 |
| Weight of prices declining, $\pi$=3.5% [4)] | 4 | 20 | 27 | 0 | 10 | 27 |
| Weight of prices declining, $\pi$=2.0% [4)] | 28 | 29 | 31 | 20 | 29 | 31 |

***Notes:*** *[1)] Residuals of the Czech Republic in the regressions between prices in the commodity groups and GDP, based on the 1999 ICP data.*

*[2)] Positive change means appreciation; yearly average over the specified period.*

*[3)] $\rho$ stands for the coefficient of relative price differences based on nominal weights. The coefficient for real weights would be about 6–12 percentage points higher (see Tables 6–9), with the adjustment process running along a similar path as for the nominal weights.*

*[4)] Nominal weight in percent of falling prices in consumption with overall inflation equal to 3.5 % or 2.0 %, respectively.*

With convergence in output and appreciation of the real effective exchange rate, the degree of differences in relative prices is likely to decline in the Czech Republic, as illustrated in Table 10. The table is based on the assumption that the rate of GDP convergence to the EU will be 1.5 and 2.5 percent per year respectively, i.e. that 1.5 or 2.5 percent of the output gap with respect to the EU average will be closed each year.[37] We also assume that the factors influencing the residuals in the price-GDP regression (such as cross-country differences in terms of trade, in sector productivities or in the degree of government interference in various economic sectors – see section 3) adjust towards their EU levels. In particular, it is assumed that the residuals in the price-GDP regressions decline over time at a rate of 2.5, 5.0 and 10.0 percent per year. Based on the regression estimates presented in Table 8, the columns in Table 10 show six different convergence

[37] As noted earlier, the 2.5 percent convergence rate is consistent with most cross-country studies of economic growth and convergence (see, for instance, Barro, 1991). The assumption of a 1.5 percent convergence rate is added to illustrate the sensitivity of the results with respect to the assumed convergence rate.

scenarios corresponding to different combinations of GDP convergence rates and the rates at which the price residuals are adjusted.[38]

As shown in Table 10, assuming that the price adjustment follows the GDP adjustment and that the adjustment of the price residual does not proceed in jumps, it may take about 15 years for the aggregate price level to reach 60 percent of the EU price level (and it may take several decades to reach 80 percent of the EU price level). It would also take about the same time to reach the degree of relative price differences in the least developed European countries. The adjustment of price levels appears to be more sensitive to the overall GDP growth, while the adjustment of the structures of relative prices is more sensitive to the factors that influence the speed with which the residuals in the simple "price vs. GDP" regression are eliminated. Changes in the speed of GDP convergence influence both the aggregate price levels and the structures of relative prices, but the impact on relative prices is relatively marginal.

The rate of real exchange rate appreciation in Table 10 can be compared with the "benchmark convergence scenario" of section 3. For the 2.5 percent GDP convergence rate and the 15-year horizon (which are the values consistent with the benchmark scenario),[39] Table 10 suggests an average rate of real effective exchange rate appreciation of 2.5–3.6 percent. This is above the "benchmark" estimate of 1.8 percent in section 3, but close to the estimate based on the "actual" elasticity for the Czech Republic (which was 2.6–3.0 percent – see also section 3), but still somewhat below the observed trend of CPI-based appreciation in the past decade.

Table 10 also presents the nominal weight in consumption of commodity groups the prices of which might be forced to decline over time for each convergence scenario. The computations were done both for the overall inflation being on the target, as well as for a consistent target undershooting, the inflation rate reaching 2 percent points only. This is a more detailed approach compared to Table 9, as it takes into account not only the GDP growth, but also the existing residuals of the Czech Republic and different scenarios of their future shrinkage. Table 10 confirms that for the realistic scenarios with relatively slow growth and modest real exchange rate appreciation, no or very few prices would need to decline. The number of falling prices, though, increases with faster GDP growth and/or real exchange rate appreciation. The calculations also confirm the basic conclusion from Table 9 that with inflation-target undershooting, an increasing number of prices may need to decline for realistic GDP growth rates. In fact, they show that the weight of falling prices may be 20-30 pervent if the inflation rate is 2 percent only. This shows the importance of targeting a slightly higher inflation rate than in the EU, and the potential costs of systematic target undershooting.

Is the possible downward rigidity of prices an important practical problem? According to economic theory, there are reasons why prices may be more downward-sticky in the short or

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[38] The factors influencing the adjustment of relative prices can be analysed more explicitly by replacing the simple regressions for commodity prices on GDP with extended regressions such as the one in Table 3. However, it is not clear how quick might be the adjustment of some of the explanatory factors (such as the share of agricultural employment in a country), if they adjust at all. The price adjustment may then be less speedy or less smooth than suggested in Table 10.

[39] In the horizon of 15 years, the 2.5 percent convergence rate implies an average 2 percentage point differential in output growth, which, as indicated in the discussion of the "benchmark", could be realistic for the Czech Republic.

medium run. Yet there is no theory which would argue that prices may be downward-sticky in the longer run. The mainstream economic theory maintains that in the long run, prices adjust flexibly to their equilibrium levels as companies are able to reset their costs and output prices in line with macroeconomic fundamentals.

What does this mean in practice? We will focus on wage growth, which represents the most important cost factor at the macroeconomic level. The anecdotal evidence on how the Czech labour unions negotiate for wage increases suggests that the unions require some percentage share of the increase in real labour productivity plus compensation for headline CPI inflation. The problem is, however, that in the "tradable" sectors, in which there is the biggest scope for productivity increases, prices of final products are not likely to grow in line with the overall CPI; their changes may in fact be negative. The above wage-negotiating formula is then misguided and only coincidentally may lead to an outcome sustainable for the company (i.e. only if the labour force's share of productivity growth is set at a low level). The only reasonable formula which would stabilise the ratio of labour costs to the value added of the company would be to give the labour force an increase equivalent to 100 percent of the real productivity growth plus the change in price of the company's value added.[40]

Given that we cannot exclude the possibility that some prices of the "most tradable" commodities might be forced to decline in some periods or even over the longer run, what is the appropriate response from policy makers and economic agents? It is the companies' and labour unions' responsibility to realise the above fact and to adjust their behaviour to the circumstances of low inflation, nominal exchange rate appreciation and falling prices of tradable goods. The only thing that policy makers can possibly do in this respect is to communicate more actively to the private sector the implications of the announced inflation targets and real convergence process. This may help overcome the behavioural aspect of the downward rigidity of prices and its real macroeconomic costs, which may stem from the companies' and labour unions' lack of experience with a low-inflation environment, an appreciating nominal exchange rate and falling prices of tradable goods.

## 5. Conclusion

The results of the calculations in this paper, derived from the 1999 ICP data, are largely in line with our earlier results derived from previous ICP projects. The new calculations confirm that there is a significantly positive cross-country relationship between aggregate price levels and outputs. Given its relative stability over time, this relationship can provide reasonable predictions of the average pace of real exchange rate appreciation in transition economies, as shown in this paper.

[40] As an example, assume that there is a company in the "tradable" sector in which labour productivity grows at 10 percent a year, but the prices of its value added are forced to decline by 3 percent owing to nominal exchange rate appreciation. Assume that the inflation rate is 3.5 percent. The appropriate wage growth in this company would be 7.0 percent, but the unions' negotiating formula would require an increase of 13.5 percent with a 100 percent share of the labour force in productivity growth, 10.5 percent with a 70 percent share, and 8.5 percent with a 50 percent share. Only if the share were by chance set at 35 percent would the negotiating formula lead to the appropriate increase of 7.0 percent.

Similarly to our previous research, we find that the Czech Republic still appears to have a lower price level than could be expected given its output per capita. However, we show that if we consider GDP per employee rather than GDP per capita and if we allow for differences in productivity in the non-tradable sector, the shares of non-tradables in GDP, government policies and terms-of-trade changes, the lower price level can be explained to a large extent.

Importantly, the 1999 ICP data are also consistent with the conclusion from our earlier research that a strong negative relationship exists between price levels and the degree of difference in the structures of relative prices in individual countries. Using the 1999 ICP data, we show how the prices of individual commodity groups, and therefore also the structures of relative prices as well as the aggregate price levels, are likely to adjust as the Czech output converges to EU countries. We find that the structure of relative prices is quite sensitive with respect to the above-mentioned additional factors that can explain the large residuals of the Czech Republic in the "price vs. GDP" regressions, such as cross-country differences in productivity in the non-tradable sector, or terms-of-trade changes. The impact of GDP on the structures of relative prices, on the other hand, seems rather minor relative to its impact on the aggregate price level.

Using the cross-country estimates based on the 1999 ICP, we assessed whether the convergence of output can force prices of some commodity groups to decline, given the existing inflation targets. This is essentially a restatement of a question analysed in our previous research: whether the process of adjustment in the structures of relative prices can lead to inflation pressures, given that some prices may not be very flexible downwards. We found that if inflation is around the target rate, the process of economic adjustment itself is not likely to lead to price declines in many commodity groups. However, if factors such as a significant exchange rate appreciation exceeding the benchmark scenario speed can put additional downward pressure on inflation, the prices of some commodities may start to decline.

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