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Sectoral Productivity and Real Exchange Rate Appreciation: Much Ado about Nothing?*

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1. Introduction

If the candidate countries are to catch up with the EU economic level, they need to achieve substantial productivity gains. Is this process consistent with maintaining moderate inflation and exchange rate stability? This dilemma is frequently discussed within the theoretical framework of the Balassa-Samuelson Effect (BSEF).[1]

Sectoral productivity improvements are likely to be associated with rises in sectoral wages. A "productivity-related" wage increase in the traded sector, however, could spill over into rising wages in the nontraded sector with lower productivity growth than in the traded sector. If wages tend to equalize across sectors in spite of productivity differentials, the nontraded sector would have to allow for higher price increases, since it cannot accommodate the rising wage level of the traded sector. For such movements in relative prices, we use – in accordance with the literature – the term "dual inflation".

By this logic, higher productivity growth in the traded sector causes higher price inflation in the nontraded sector. This translates either into a rising domestic CPI level (which causes cross-country inflation differentials to emerge) or into nominal exchange rate appreciation (or some combination of the two). As a result, in a country with higher productivity growth

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[1] (Balassa, 1964) and (Samuelson, 1964).

of tradables relative to nontradables than abroad, there are appreciation pressures on the real exchange rate.

If such a mechanism really works, then the EU candidate countries could face, as a by-product of the catching-up process, a trend of real exchange rate appreciation, and the process of joining the EMU (ERM2) could be adversely affected by incompatibility of the "real" catching-up process with the "nominal" convergence criteria.

In candidate countries with fixed exchange rates, the existence of the BSEF would probably imply a strengthening of monetary restriction (and consequently a slowdown in the real catching-up process), if inflation is to be kept below the Maastricht criterion.

In the case of floating exchange rates, the dilemma between the nominal convergence criteria and real catching-up seems less pronounced.[2] But one has to note that, in the presence of the BSEF, a floating exchange rate regime imposes a trade-off between the inflation target and exchange rate stability. Consequently, one cannot rule out the possibility of the BSEF having a negative impact on real catching-up under floating exchange rates as well (for example, in the form of a worsening trade balance amid rapid exchange rate appreciation).

Yet for the EU candidate countries (and their central banks) with floating exchange rates, such as the Czech Republic, estimating the magnitude of the BSEF is important – the stronger the BSEF, the bigger part of the ongoing real exchange rate appreciation could be explained by "structural" or "equilibrium" forces, with robust implications for the central banks' view on the optimum real exchange rate evolution before, as well as after, entering the ERM2.

Researchers are currently studying these issues in Western Europe too, despite the absence of the catching-up problem there. Alberola and Tyrvainen (1998, p. 7) explain this interest in studying the BSEF by the fact that "[...] problems may arise if inflation differentials persist in EMU [...] From the point of view of the ECB, the less uniform the inflation in the EMU countries, the less straightforward is the choice of the appropriate stance of the common monetary policy. From the point of view of a member country, higher inflation leads to a change in relative prices in a manner which is equivalent to an appreciation of the 'real exchange rate'."

Research in western, as well as in developing, countries has generated varying results. Analyses conducted in the late 1960s, as well as some more recent findings, seem to confirm the presence of the BSEF.[3]

In contrast, Alberola and Tyrvainen (1998) demonstrate that for developed European countries this effect does not hold without allowing wages to enter the regressions (i.e., the authors relax the condition of wage equality of both sectors). Kohler's (1999) results include some developing (non-European) countries and these results are also rather mixed.

[2] Some authors argue that, in fact, the BSEF in the candidate countries would have to be very large in future to exceed the ERM2 fluctuation band of ±15 %. Therefore, the effect should be (relatively easily) masked by nominal exchange rate fluctuations within the band, with actually no impact on inflation. See, for example, (Jonáš, 2001) for a discussion.

[3] See, for example, (Edgren – Faxen – Odhner, 1969), (Canzoneri – Cumby – Diba – Eudey, 1998) or (Weidmann, 2002).

Empirical research on the EU candidate countries seems to signal almost uniformly the presence of the BSEF.[4] To be more precise, the empirical literature at our disposal questioning the relevance of the BSEF for the EU candidate countries is rather scarce.[5] At the same time, however, some authors question the adequacy of the whole approach for evaluating the impact of the real catching-up process on the nominal convergence criteria.[6]

In order to contribute some additional arguments to the ongoing discussion, we find it useful in this paper to study the developments in the Czech Republic in combination with selected EU countries.[7]

We first identify the presence of productivity differentials between the traded and nontraded sectors and then test their link to relative price developments. Furthermore, we demonstrate how dual inflation stemming from sectorally unbalanced productivity growth translates into real exchange rate appreciation (or, in our case, in cross-country inflation differentials, assuming constant nominal exchange rates).[8]

To achieve the above-mentioned goals, this paper is organized in the following manner: In Section 2 we define a simple model to formulate the mechanism of the BSEF in a way that enables subsequent empirical testing. Section 2 also includes a description of data sources. Sections 3 and 4 both deal with quantitative analysis. In Section 3 we estimate the impact of sectoral productivity differentials on dual inflation. In Section 4 we calculate (simulate) the impact on the real exchange rate (cross-country CPI-inflation differentials) stemming from the BSEF, and in Section 5 we set forth our conclusions. In Appendix we summarize briefly the theoretical formulation of the BSEF and its link to the analytical framework used in this paper.

2. The Model and Data Sources

2.1 The Model

Practically all BSEF-related literature starts with the formulation of a two-sector model involving Cobb-Douglas production functions. Then, the determinants of factor prices under perfect competition and factor mobility are derived for both sectors, and finally the BSEF hypothesis is formulated in the form of an equation for domestic relative prices. This

[4] See, among others, (Begg et al., 2001) and (Kovács, 2002) for the most recent empirical results supporting the presence of the BSEF and a broader overview of the literature.

[5] See (Egert, 2002) and (Mihaljek, 2002).

[6] For example, Nuti (2001, p. 13) points out the difficulty of separating tradables and non-tradables accurately. Moreover, he argues that "after all, tradables are both inputs in non-tradable goods, and substitutes for non-tradables".

[7] The majority of studies on this topic deals with (quite heterogeneous) panels of transition countries or with the EU countries only. In contrast, we include the Czech Republic in the panel of highly developed EU countries and attempt to compare our results with the above noted approaches.

[8] We follow the approach of Kohler (1999) and others, who also assume constant nominal exchange rates and demonstrate the BSEF in the light of cross-country inflation differentials. For more discussion see Section 4.

yields the following "testable" version of dual inflation (see the Appendix for more detail):

$$\ln P_{i,t}^{rel} = \alpha_i + \beta \ln \left(\frac{LP_{i,t}^{tr}}{LP_{i,t}^{ntr}} \right) + \varepsilon_t \qquad (1)$$

$$[\beta = \beta_1 = \beta_2 \text{ and } \beta_1 \ln (LP_{i,t}^{tr}) - \beta_2 \ln (LP_{i,t}^{ntr}) = \beta \ln (LP_{i,t}^{tr} / LP_{i,t}^{ntr})]$$

$LP_{i,t}^{tr} / LP_{i,t}^{ntr}$ relative labour productivity (value added per employee) in country *i* and time period *t*,
$P_{i,t}^{rel} = P_{i,t}^{ntr} / P_{i,t}^{tr}$ relative prices (sectoral value-added deflators).

The equation of dual inflation (1) adopted, *inter alia*, by Canzoneri, Cumby, Diba and Eudey (1998), Kohler (1999), Egert (2002), Mihaljek (2002) and Weidmann (2002), formalises the relationship between sectoral price and productivity movements. In other words, it enables us to estimate, under the given simplifications and constraints, the impact of a one percent change in relative labor productivity (traded to nontraded sector) on the sectoral relative price ratio (nontraded to traded).[9]

Model specification (1) is based solely on the sectoral indicators of each economy. This is why it is sometimes called the "domestic" or "internal" version of the BSEF, since the model at this stage determines relative prices only.[10]

Under the condition of equal wages in both sectors, the BSEF would predict coefficient β to be positive and equal to one. Consequently, the lower the empirical value of β, the more likely is the violation of the wage spillover condition (or profit-maximization conditions, as summarized in equations i-iv in the Appendix). Coefficient β is quantified in Section 3 using various panel estimations. In such a way, our results can be discussed within the context of the related literature.

The above-mentioned authors assume, in line with the BSEF theory, that the marginal effects of productivity on relative prices are equal in the traded and nontraded sectors ($\beta = \beta_1 = \beta_2$). Strictly speaking though, this theoretical assumption of equal marginal effects should not be taken for granted in the real world, so we make it subject to empirical testing. If we allow for different marginal effects of productivity in the traded and nontraded

[9] In the logic of the BSEF theory, the law of one price holds in the traded sector. This means that *P_tr* is determined by the world market and purchasing power parity holds such that *P_tr* = = *P*_tr E*, where *P*_tr* denotes the world price and *E* the nominal exchange rate. At the same time, *P_ntr* is determined merely by the "domestic" market. In contrast, the "testable" version of the dual inflation equation simply deals with sectoral value-added deflators (output prices), as available in the statistics. Analogously, total factor productivity is replaced by the available labour productivity indicators. See, for example, (Kohler, 1999) for a more detailed theoretical background and for the simplifications under which the model was developed.

[10] Therefore, empirical verification of model specification (1) alone does not necessarily express the magnitude of the BSEF in terms of consequences for real exchange rate evolution (cross-country inflation differentials under constant nominal exchange rates). Further steps are necessary to document the link between dual productivity and the real exchange rate.

sectors on relative prices, we would have to deal with a more general version of equation (1):

$$\ln P^{rel}_{i,t} = \alpha_i + \beta_1 \ln(LP^{tr}_{i,t}) - \beta_2 \ln(LP^{ntr}_{i,t}) + \varepsilon_t \quad (2)$$
$$[\beta_1 \neq \beta_2]$$

When performing the empirical analysis in Section 3, we would have to discriminate between models (1) and (2), depending on the test for equality of marginal effects. This is our intended conceptual contribution to the BSEF-related literature.

By estimating the common slope coefficients β within a panel regression framework, we are able to determine the extent to which dual productivity influences dual inflation, which is common for all the countries included in the panel. Subsequently, assuming constant nominal exchange rates, and taking into account the country-specific weights of tradables in consumption (value added), we can calculate for each country the "implied" (or "domestic") inflation which is attributable to dual productivity. Under the assumption of equality of βs, we obtain:

$$\Delta \ln CPI_{i,t} = \delta_{i,t} \beta \Delta \ln \left(\frac{LP^{tr}_{i,t}}{LP^{ntr}_{i,t}} \right) \quad (3.1)$$

CPI consumer price index,
δ share of nontradables in the CPI (approximated by the share of nontradables in value added),
β coefficient estimated in equation (1),

while for the case of two βs:

$$\Delta \ln CPI_{i,t} = \delta_{i,t} \, [\beta_1 \ln(LP^{tr}_{i,t}) - \beta_2 \ln(LP^{ntr}_{i,t})] \quad (3.2)$$

Finally, by comparing the values of "implied" (or "domestic") inflation internationally with a "benchmark" country (or group of countries), we can calculate (simulate) cross-country inflation differentials stemming from the BSEF. The interpretation of the results is (under constant nominal exchange rate E against the "benchmark" country B) analogous to real exchange rate change due to dual productivity differential:[11]

$$\Delta \ln RER_{i,t} = \Delta \ln CPI_{i,t} - \Delta \ln CPI_{B,t} - \Delta \ln E_{i,t} =$$
$$= \beta \left[\delta_{i,t} \Delta \ln \left(\frac{LP^{tr}_{i,t}}{LP^{ntr}_{i,t}} \right) - \delta_{B,t} \Delta \ln \left(\frac{LP^{tr}_{B,t}}{LP^{ntr}_{B,t}} \right) \right] \quad (4.1)$$
$$[\beta = \beta_1 = \beta_2; \; \Delta \ln E_{it} = 0]$$

[11] Note that $\Delta \ln E > 0$ means nominal exchange rate depreciation. Contrary to our model specifications (3) and (4), Mihaljek (2002) uses in his country regression framework the difference between CPI inflation in the Central European country and in the euro area as the dependent variable, while the respective productivity differentials at home vis-à-vis the euro area and nominal exchange rate changes stand as explanatory variables. Egert (2002) also uses an analogous model specification. We discuss the implications of this in more detail in Section 4.

When using two βs, we yield:

$$\begin{aligned}\Delta \ln RER_{i,t} &= \Delta \ln CPI_{i,t} - \Delta \ln CPI_{B,t} - \Delta \ln E_{i,t} = \\ &= \delta_{i,t}\,[\beta_1 \ln(LP^{tr}_{i,t}) - \beta_2 \ln(LP^{ntr}_{i,t})] - \delta_{B,t}\,[\beta_1 \ln(LP^{tr}_{B,t}) - \beta_2 \ln(LP^{ntr}_{B,t})]\end{aligned} \tag{4.2}$$

$$[\beta_1 \neq \beta_2;\ \Delta \ln E_{it} = 0]$$

Indeed, with common coefficient(s) β and identical (or similar) shares of nontradables in consumption (δ) across the investigated countries, the BSEF-related real exchange rate appreciation (i.e., in our case, the positive BSEF-implied cross-country inflation differential under a stable nominal exchange rate) will be determined merely by a higher productivity differential at home than abroad.

2.2 The data

The data sources used in the analysis are the OECD International Sectoral Data Base (1970–1995/7), Eurostat New Cronos (1996–1999) and the Bulletins of the Czech Statistical Office (1995–2001). We test the following countries for the presence of the BSEF: Belgium, Denmark, Finland, France, Italy, the Netherlands, the United Kingdom, (West) Germany, and the Czech Republic. Manufacturing and agriculture represent the traded sector (*tr*), while construction and transport represent the nontraded sector (*ntr*).

While the representatives of the traded sector are selected here in accordance with the prevailing convention, in the case of the nontraded sector we restrict ourselves, somewhat unconventionally, to construction and transport (unbalanced panel I) or even to construction only (unbalanced panels II and III). Given the existing nontrivial methodological difficulties in separating tradables and nontradables accurately, such an approach is predetermined merely by data availability restrictions.

The bulk of the studies dealing with the BSEF in the EU countries explore the OECD International Sectoral Data Base, possibly in combination with country-specific national accounts. In this way, however, it is difficult to include data for the second half of the 1990s in the analysis. In order to cope with this problem, we merge the OECD sources with Eurostat New Cronos, where data for the second half of the 1990s is available, albeit in a different structure in many cases (gross output instead of value added; producer price indexes instead of value-added deflators; missing data for certain sectors and periods).[12]

In order to integrate the two data bases without violating substantially the consistency of the data, we have to be extremely selective in choosing

[12] The International Sectoral Data Base has been specially designed to facilitate the calculation of indices of productivity at a detailed industry level. It provides annual time series data covering the period 1970–1995/7 for 14 OECD member countries. Detailed information is available at http://www.oecd.org//std/isdbsw.pdf. See also http://www.europa.eu.int/newcronos for the Eurostat data used for the second half of the 1990s.

the representatives of the nontraded sector, even in the case of the EU countries. In fact, only the data for construction and transport are of use for such an exercise. With regard to the Czech Republic, the situation is even more complicated, since there are no time series available which contain value-added indicators (and value-added deflators) in the required sectoral breakdown. Instead we have to use gross-output-per-worker indicators (and the corresponding producer price indexes) to incorporate the Czech Republic into the analysis.

As with the recent EU data, the structure of the Czech data permits the inclusion of construction and transport only. In the case of the latter, moreover, substantial difficulties arise with aggregating the data for the various branches of this sector (which includes the state-owned railways as well as foreign-owned mobile telephone operators). Because of this, when investigating the BSEF we have to rely predominantly on construction as the sole representative of the nontraded sector.

The above-mentioned problems with data create serious limitations for interpreting our empirical results.[13] These should be understood rather in terms of experiments which may or may not bring about statistically significant results and signal the potential importance of the BSEF in such a way.

On the other hand, to our knowledge it is highly questionable whether there are more-promising approaches available for analyzing internationally the recent evolution of the BSEF which would permit the inclusion of the Czech Republic or any other transition country.[14]

Having made the above reservations, we can now describe the structure of the data (*Tables 1–3*). The structure of the OECD data is as follows: value added per employee at constant prices and the value-added deflator for each sector. The Eurostat data includes gross value added per employee in M (manufacturing), C (construction) and T (transport and communications), deflated by the industrial producer price index in M, the output price index in C and by a price index derived from the price level and evolution in T. In sector A (agriculture), the Eurostat data used is as fol-

[13] There is also a problem with using producer (output) prices instead of value-added deflators for the most recent period in all three panels. This could bias the reported relative price and productivity developments as a whole and influence the estimates of β_1 and β_2. Ideally, one would have to check for existing quantitative differences between output prices and value--added deflators and see whether such potential errors due to data constraints are qualitatively important. We thank Tomáš Holub for this comment. We also neglect the existence of regulated prices in some segments and many other, still nonstandard features of the Czech economy.

[14] The literature investigating the BSEF in transition countries is, with a few exceptions, less explicit with respect to indicating data sources, and we suspect that nontrivial problems with data availability/reliability are inherent for all BSEF-related literature. Mihaljek (2002, p. 6) illustrates these problems in specific terms: "[...] most studies [...] try to compensate for the short time series by pooling data from different transition economies [...] from advanced EU accession candidates in Central Europe to relatively underdeveloped Central Asian CIS economies." He adds that the traded sector includes "[...] often also construction as well as electricity, gas and water supply, industries whose output is only to a small extent traded. The traded sector is in some studies the residual (i.e., GDP less industry). In others, it covers all services irrespective of their traded content. Some studies do not even consider nontradables, assuming that productivity growth in the sector is zero or equal across countries."

TABLE 1 Unbalanced Panel I – Four Sectors

Country	Period	Sectors	Source
Belgium	1970–1997	*M+A*,[a] *C+T*[b]	1970–1995 OECD; 1996–1997 Eurostat
Denmark	1970–1996	*M+A*,[a] *C+T*[b]	1970–1995 OECD; 1996 Eurostat
Finland	1970–1997	*M+A*,[a] *C+T*[b]	1970–1995 OECD; 1996–1997 Eurostat
France	1970–1997	*M+A*,[a] *C+T*[b]	1970–1997 OECD
Germany[c]	1970–1994	*M+A*,[a] *C+T*[b]	1970–1994 OECD
Italy	1970–1997	*M+A*,[a] *C+T*[b]	1970–1997 OECD
Netherlands	1970–1996	*M+A*,[a] *C+T*[b]	1970–1995 OECD; 1996 Eurostat
United Kingdom	1970–1993	*M+A*,[a] *C+T*[b]	1970–1993 OECD

Notes: [a] Manufacturing (*M*) and Agriculture (*A*) = traded sector
[b] Construction (*C*) and Transport (*T*) = nontraded sector
[c] West Germany only

Sources: http://www.oecd.org//std/isdbsw.pdf; http://www.europa.eu.int/newcronos

TABLE 2 Unbalanced Panel II – Two Sectors

Country	Period	Sectors	Source
Belgium	1970–1999	*M+C*	1970–1995 OECD; 1996–1999 Eurostat
Denmark	1970–1999	*M+C*	1970–1995 OECD; 1998–1999 Eurostat
Finland	1970–1999	*M+C*	1970–1997 OECD; 1998–1999 Eurostat
France	1970–1999	*M+C*	1970–1997 OECD; 1998–1999 Eurostat
Germany	1970–1994	*M+C*	1970–1994 OECD
Italy	1970–1999	*M+C*	1970–1997 OECD; 1998–1999 Eurostat
Netherlands	1970–1999	*M+C*	1970–1995 OECD; 1996;1998–1999 Eurostat
United Kingdom	1970–1993	*M+C*	1970–1993 OECD

Note: See Table 1 for definitions of sectors and for data sources.

TABLE 3 Unbalanced Panel III – Two Sectors

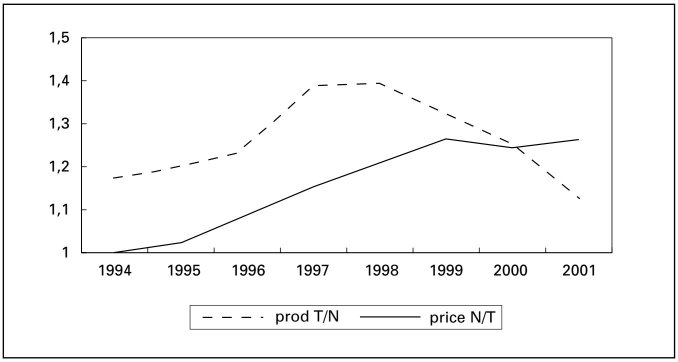
Country	Period	Sectors	Source
Belgium	1992–1999	*M+C*	1992–1995 OECD; 1996–1999 Eurostat
Denmark	1991–1999	*M+C*	1991–1996 OECD; 1998–1999 Eurostat
Finland	1992–1999	*M+C*	1992–1995 OECD; 1996–1999 Eurostat
France	1992–1999	*M+C*	1992–1997 OECD; 1998–1999 Eurostat
Germany	1987–1994	*M+C*	1987–1994 OECD
Italy	1992–1999	*M+C*	1992–1997 OECD; 1998–1999 Eurostat
Netherlands	1990–1999	*M+C*	1990–1995 OECD; 1996; 1998–9 Eurostat
United Kingdom	1986–1993	*M+C*	1986–1993 OECD
Czech Republic	1994–2001	*M+C*	1994–2001 CSO

Notes and Sources: See Table 1.

lows: (gross value added at constant prices / employment in agriculture, hunting, forestry and fishing), deflated by the index of producer prices of agricultural products.

The two data bases have been integrated, with 1990 as the base period. The Czech data includes annual labor productivity indicators and the cor-

FIGURE 1 Sectoral Productivity Ratio (*T/N*) and Relative Prices (*N/T*) in the Czech Republic



Source: Bulletins of the CSO, 1994–2001

responding price indices, as officially published by the Czech Statistical Office, with 1994 as the base period.

We are aware that the BSEF assumes the existence of long-run time series, a condition difficult to achieve in the Czech Republic because of its relatively short history of a functional market economy. Therefore, when interpreting our empirical results, not only the theoretical and methodological reservations, but also the limited data availability, should be noted.

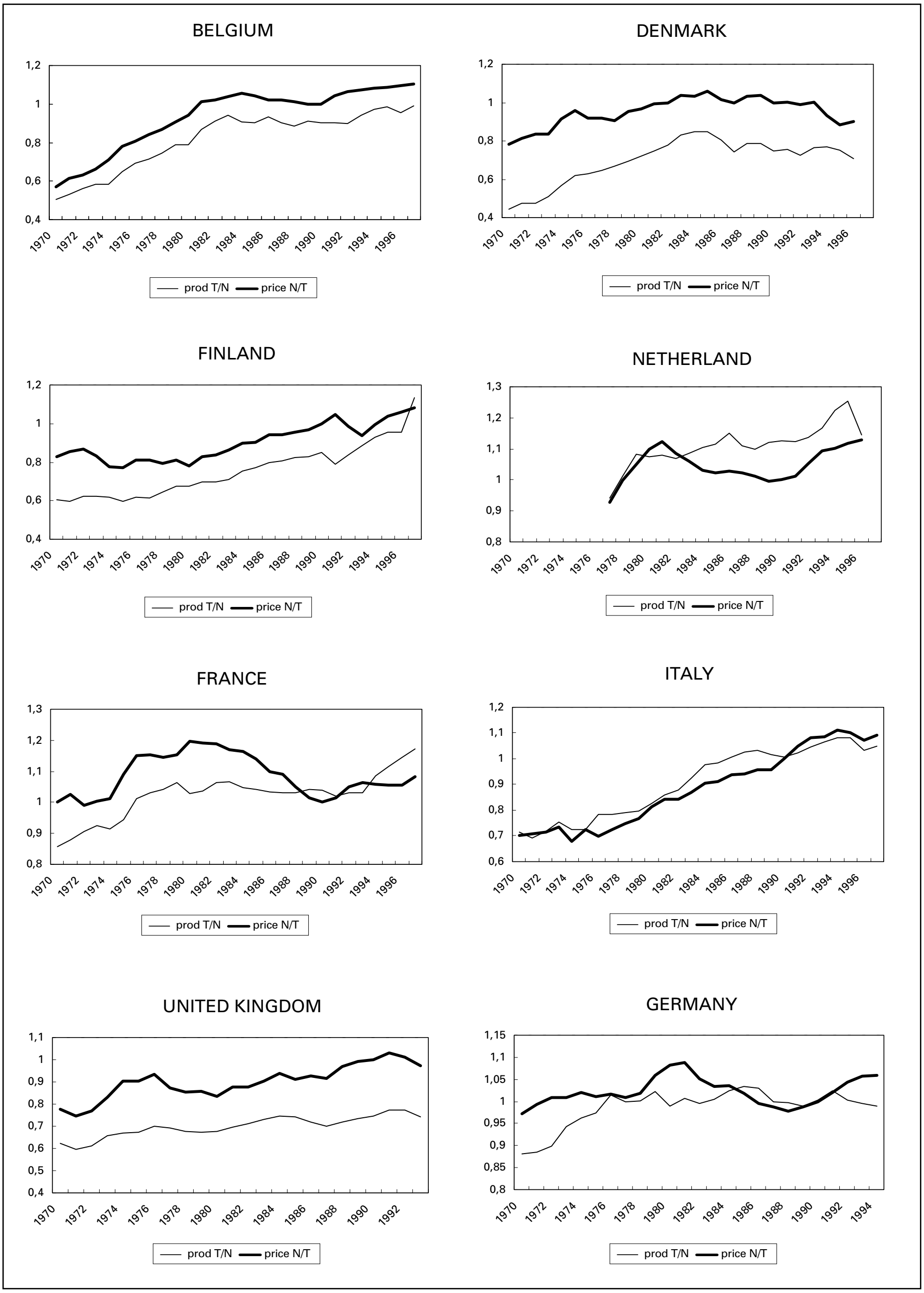
3. Testing for the Presence of Dual Inflation

In this section, we intend to analyze in more detail the phenomena of dual inflation, both in selected EU countries and in the Czech Republic. Before employing the methods of econometric analysis, we first look at the results of descriptive statistics. *Figure 1* shows a sharp break in relative productivity developments since 2000. Consequently, any straightforward interpretation of the Czech data is rather difficult.

With regard to the developments in the selected EU countries, there are three basic tendencies (*Figures 2* and *3*):

1. There is a trend of faster productivity growth in the traded sector than in the nontraded sector (see the prevailing upward slope of the T/N lines, representing the ratio of sectoral productivity levels at national constant prices). This makes further analysis plausible, because the basic condition exists from which the entire causal mechanism of the BSEF starts.
2. Figures 2 and 3 display in most cases a remarkable correlation between the sectoral productivity ratios (T/N) and relative price developments (N/T). These results stand for the different representatives of the traded and nontraded sectors in both figures. This could be interpreted as *prima facie* evidence of the presence of dual inflation and further justifies more advanced analysis along the lines of the approach developed in Section 2.

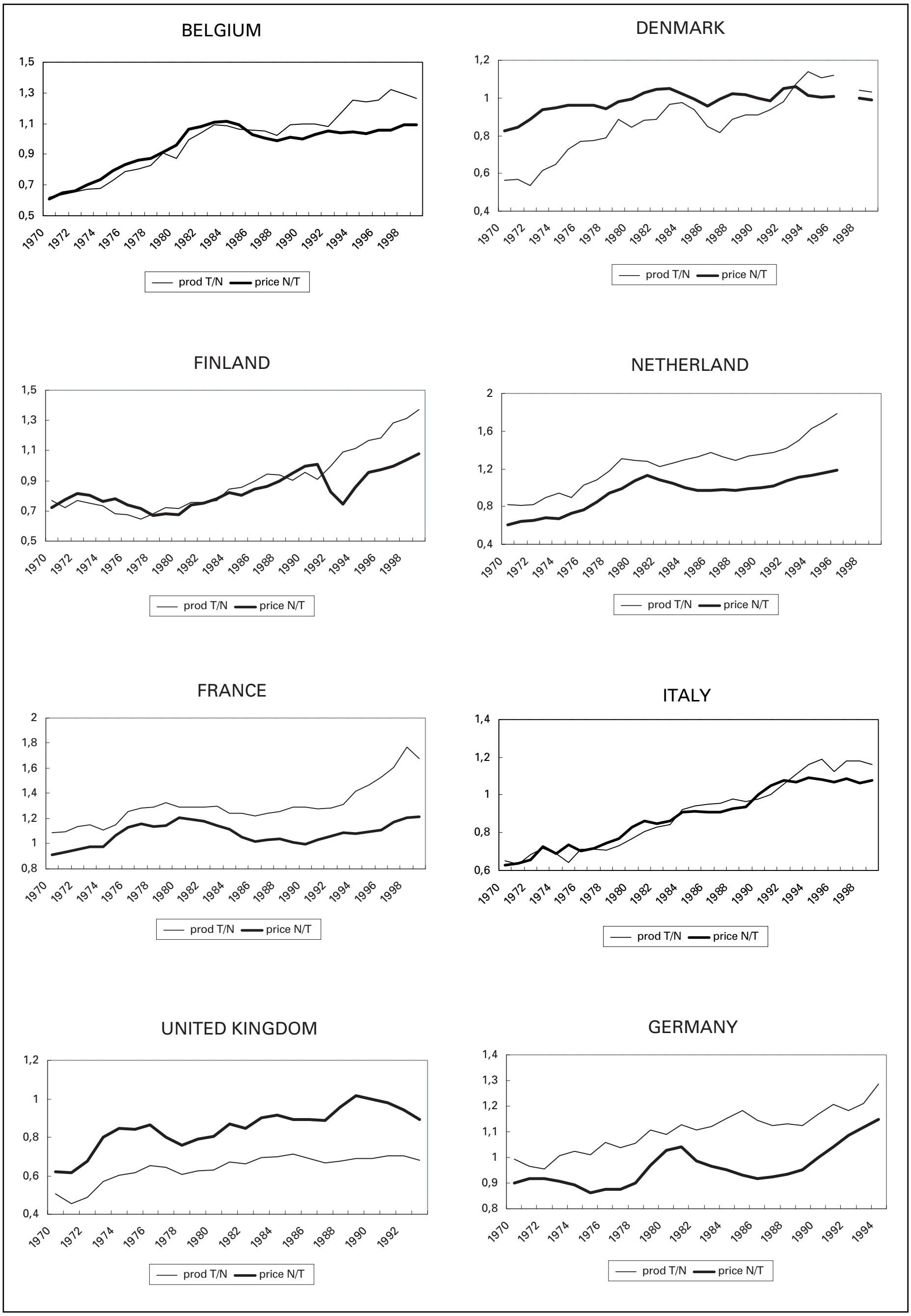
FIGURE 2 Sectoral Productivity Ratio (*T/N*) and Relative Prices (*N/T*) – Four Sectors



Notes: *T* = manufacturing + agriculture; *N* = construction + transport. For each year, sectoral productivity levels in national currencies are used for calculating the prod *T/N* ratio, while in the case of the price *N/T* we use sectoral basic price indexes (1990 = 100).

Source: Eurostat, OECD

FIGURE 3 Sectoral Productivity Ratio (*T/N*) and Relative Prices (*N/T*) – Two Sectors (*M+C*)



Note and Source: See Figure 1.

3. Perhaps surprisingly, the comparison of productivity levels in the two sectors reveals that the T/N line, although upward-sloping, remains in some cases below 1. This indicates the presence of higher productivity in the nontraded sector. For example, in the UK, the ratio of sectoral productivity levels, T/N, is persistently below 1.[15]

To analyze the phenomena of dual inflation more accurately, we use the unbalanced panel data from Tables 1–3 to estimate equations (1) and (2). The results of the estimations are summarized in *Table 4*.

In the cases of unbalanced panels I and II, the F test justifies the use of equation (1), because the test reveals equality of βs.[16]

As far as unbalanced panel III is concerned (where data for the Czech Republic are also included), the situation is different. The result of the F test necessitates the adoption of two βs (i.e. using equation (2)) because of two different marginal effects of sectoral productivity on relative prices.[17]

Nevertheless, in the case of unbalanced panel III, we further compute both models in parallel for the sake of comparison with the existing literature on this topic.

All the estimations include testing of fixed effects across countries (F test), cross-sectional heteroskedasticity (Lagrange Multiplier [LM] test), cross-sectional correlation (LM test), serial correlation (DW statistics), and the existence of a common slope across countries (F test). Table 4 presents the results.

The cross-sectional correlation is insignificant, as is the cross-sectional heteroskedasticity. The serial correlation proves to be a serious problem and has been accounted for by using the DW iterative procedure, which leads to the most efficient removal of serial correlation.[18]

We find significant fixed effects which means that each country has its own idiosyncratic constant price ratio between the nontraded and traded sectors. The slopes across countries, however, do not vary statistically sig-

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[15] These results are most frequently present for the 1970s and 1980s. Because of this, they cannot be the product of merging the OECD and Eurostat data bases and we have to take them for granted. Moreover, the core of our analysis is unbalanced panel III, where only recent developments are included.

[16] The F test enables us to compare the restricted model with a single β with the unrestricted model with two different βs and use the residuals from both models to test against F statistics critical tables. This shows whether or not we can restrict the model and use only a single β. The probability of having a model with one β proves to be satisfactorily high in the case of unbalanced panels I and II (the corresponding p-values being 0.24 and 0.21 respectively, see Table 4).

[17] The probability of rejecting the null hypothesis (i.e., that there is a single β coefficient for both marginal contributions) is 0.999, as presented in Table 4. Hence, we have proven the existence of two βs at the 1% significance level.

[18] This technique consists of a repeated Cochrane-Orcutt transformation of the data, as long as the DW statistic falls into an inconclusive region. This allows us to keep the structure of the model in the given form, i.e. without lagged terms in the regression. The repeated data transformation is as follows: $y_t - \rho y_{t-1} = x_t - \rho x_{t-1} + \varepsilon_t$, where y_t is the dependent variable, x_t is the independent variable, ρ is the coefficient of first-order autocorrelation and ε_t denotes the error term, which is presumed to be serially uncorrelated. This is, however, tested again in the next stage. If the DW statistic shows again that the residuals exhibit serial correlation, the procedure is repeated. See (Green, 2000) for details.

TABLE 4 Estimation Results

Model Specification (1): $\ln P^{rel}{}_{it} = \alpha_i + \beta \ln (LP^{tr}{}_{it} / LP^{ntr}{}_{it}) + \varepsilon_t$

Model Specification (2): $\ln P^{rel}{}_{it} = \alpha_i + \beta_1 \ln (LP^{tr}{}_{it}) - \beta_2 \ln (LP^{ntr}{}_{it}) + \varepsilon_t$

Model Specification: Panel Time span No. of observations	(1) UB I 4S[a] 1970–1997 208	(1) UB II 2S[a] 1970–1999 226	(1) UB III 2S[b] 1986–2001 72	(2) UB III 2S[b] 1986–2001 72
Common slope (standard error)	0.45*** (0.053)	0.36*** (0.058)	0.59*** (0.082)	β_1 β_2 0.65*** 0.53*** (0.073) (0.111)
F test $\beta = \beta_1 = \beta_2$	$1.35 < F_{1,192}$ (0.2467)	$1.57 < F_{1,216}$ (0.2116)	$16.43 > F_{1,61}$ (0.00015)	$16.43 < F_{1,61}$ (0.99985)
Cross-sect. correl. LM test[c]	$4.2 < \chi_{28}^2$ (1.0)	$5.2 < \chi_{28}^2$ (1.0)	$4.1 < \chi_{36}^2$ (1.0)	$4.12 < \chi_{36}^2$ (1.0)
Cross-sectional heteroskedasticity LM test[d]	$1.4 < \chi_8^2$ (0.99)	$2.4 < \chi_8^2$ (0.97)	$0.03 < \chi_9^2$ (1.0)	$0.02 < \chi_9^2$ (1.0)
Serial correlation DW test[e]	1.8 (14)	1.8 (14)	1.63 (16)	1.63 (16)
R^2	0.76	0.76	0.85	0.89
Fixed country effect[f] test	$8.72 > F_{7,199}$ (0.000)	$16.05 > F_{7,217}$ (0.000)	$4.17 > F_{8,62}$ (0.000)	$4.16 > F_{8,61}$ (0.000)
Common slope test[g]	$0.68 > F_{7,192}$ (0.6824)	$0.24 > F_{7,210}$ (0.9757)	$1.27 > F_{8,54}$ (0.2793)	$1.06 > F_{16,46}$ (0.4172)

[a] Belgium, Denmark, Finland, the Netherlands, France, Italy, the U.K., and W. Germany, four sectors

[b] Includes the Czech Republic, two sectors. Without the Czech Republic data, β = 0.61*** (0.081) in model specification (1) and β_1 = 0.55*** (0.1) β_2 = 0.27*** (0.03) in model specification (2).

[c] Lagrange multiplier test $\lambda_{LM} = T \sum_{i=2}^{N} \sum_{j=1}^{i-1} r_{ij}^2$ where r_{ij}^2 is the ij^{th} residual correlation coefficient, which was calculated using OLS residuals. Presented are the probabilities of not rejecting H_o: No cross-sectional correlation.

[d] White test for heteroskedasticity $LM = \frac{T}{2} \sum_i [\frac{s_i^2}{s^2} - 1]^2$, where s^2 is the variance of the OLS residuals. Presented are the probabilities of not rejecting H_o: No cross-sectional heteroskedasticity.

[e] DW statistics after the iterative method leads to the state of no first-order autocorrelation. The numbers of iterations needed are shown in parentheses.

[f] Using F test: $FT = \frac{(SSRR - SSRU)/R}{SSRU/(n-k)} \sim F(R, n-k)$, where R is the number of restrictions and k is the number of regressors in the unrestricted model. SSR stands for the sum of the squared residuals. Presented are the probabilities of not rejecting H_o: No fixed effect.

[g] Presented are the probabilities of not rejecting H_o: No country specific slope.

nificantly and the tests of the restricted model prove the existence of a common slope coefficient(s) β for all countries.[19]

In all four regressions (unbalanced panels I–III, UB III in two specifications, as explained above), the β coefficients are significant at the 1% significance level (in Table 4, this is denoted by ***). The fit of all four regressions turns out to be satisfactorily high (see the R^2 values in Table 4).

First, we deal in detail with unbalanced panel I (which includes the EU countries only, and four sectors). The single common slope coefficient β is significant and positive. These findings are in accordance with the BSEF theory. At the same time, however, the value of the coefficient is 0.45, thus indicating a lower long-term impact of relative productivity developments on relative prices than the BSEF theory would predict (according to the BSEF theory, $\beta = 1$).

When only two sectors are included (unbalanced panel II), the results are similar, though the value of the common slope coefficient β is slightly lower (0.36) than in the previous estimate (0.45). Nonetheless, the use of a two-sector-model is justified, since it generates statistically significant results of approximately the same range as the four-sector model. This finding is important for the further analysis, where the Czech Republic data is included.

Finally, we use the data in a different structure (unbalanced panel III, model specifications (1) and (2)), in order to include the Czech Republic and also to reflect predominantly the recent developments abroad. Eight annual observations (1994–2001) are available for the Czech Republic, and the eight most recently available observations are included for the remaining countries as well (ranging from 1986–1993 for the UK to 1992–1999 for Belgium, Italy, Finland and France, see Table 3 for details).

When adopting model (1) for the data in unbalanced panel III, we see for recent developments in the EU countries a higher degree of dual inflation ($\beta = 0.61$)[20] than that prevailing over the whole period 1970–1997 ($\beta = 0.36$).[21] Most importantly, however, the inclusion of the Czech Republic data does not alter the estimates significantly ($\beta = 0.59$).[22]

For model specification (2), which also explores the data of unbalanced panel III (including the Czech Republic), we find coefficients for the traded sector $\beta_1 = 0.65$ and for the nontraded sector $\beta_2 = 0.53$. Without the Czech Republic, the coefficients are as follows: $\beta_1 = 0.55$ and $\beta_2 = 0.27$.

While the inclusion of the Czech data does not alter the estimates of β(s) significantly in the case of model specification (1), in the case of specification (2) it does. At this stage of our research, we have yet to find the factors behind such a result.

Egert (2002) uses the Johansen cointegration test to analyze the link between relative productivity and relative prices in the Czech Republic. He estimates coefficient β in the 95% confidence interval ranging between 0.56 and 0.72 (compared to our $\beta = 0.59$). The same author also explores panel cointegration analysis, with a similar result, i.e. with $\beta < 1$.

[19] The F test examines whether the βs differ statistically significantly across countries, or they can be considered equal. The corresponding p-values of not rejecting the null hypothesis of having one β(s) for unbalanced panels I–III (0.68; 0.98; 0.28; 0.41) prove the existence of a common slope coefficient(s) β. See Table 4.

[20] See note b below Table 4.

[21] See the estimation results for UB II 2S in Table 4.

[22] See the estimation results for UB III 2S; model specification (1).

Mihaljek (2002) estimates a country regression model for the Czech Republic and arrives at an even lower β compared to our results. One must stress, however, that both Egert (2002) and Mihaljek (2002) use model specification (1) in the above-quoted cases.

In contrast, we test the impact of relative productivity on relative prices also in model specification (2) and conclude that the impact of the non-traded sector's productivity on relative prices is lower than model specification (1) assumes and the estimated difference between β_1 and β_2 is statistically significant at the 1% significance level.[23] It follows that conclusions based solely on model specification (1) could be biased.

As noted above, we obtain $\beta_1 > \beta_2$, with a ratio of about 1.23. Čihák and Holub (2001) use cross-country regressions for about 30 commodity groups (instead of sectors) and their ratio of the estimated coefficients is approximately 1.3, thus resembling our results once again.

As a result, we find that sectoral productivity developments have a statistically significant impact on relative prices in the EU countries and also in the Czech Republic, but the magnitude of the impact is not as strong as the BSEF would predict (in both model specifications).

4. Real Exchange Rate Evolution in the Light of Cross-Country Inflation Differentials

The tests performed in Section 3 suggest the presence of dual inflation for both the Czech Republic and the EU countries. In addition to the value of the β coefficient(s), the final "implied" impact on domestic CPI inflation depends, under stable nominal exchange rates, on the country-specific share of nontradables in consumption and also on the quantity of sectoral productivity divergence (see model specifications (3.1) and (3.2) in Section 2).

Comparing the magnitude of "domestic" or "implied" CPI inflation internationally, we obtain cross-country inflation differentials whose interpretation is, under stable nominal exchange rates, analogous to real exchange rate changes due to a dual productivity differential (see model specifications (4.1) and (4.2) in Section 2).

By leaving nominal exchange rates constant, we do not have to deal with the problem of which currency to use as a benchmark for comparison. Yet the final impact on the real exchange rate depends on the difference between "implied" inflation at home and in the reference country, which in turn depends on the productivity differential at home against that in the reference country.

Tables 5 and 6 summarize, in accordance with our equations (3.1) and (3.2), the results of simulations of the "domestic" or "implied" CPI inflation for nine countries.

In *Table 5*, the second column includes country-specific productivity growth differentials between the traded and nontraded sectors

[23] The results of the F test are presented in Table 4. The probability of rejecting the hypothesis that the two β coefficients are equal is 0.99.

TABLE 5 Simulated Annual BSEF-implied Impact on Domestic CPI (in percentage points) (model specification (3.1), for $\beta = 0.59$; unbalanced panel III)

Country	Relative Productivity growth[a] $\Delta\ln(LP_{tr}/LP_{ntr})$	Country share[b] $Q_n/(Q_n+Q_t)$		"Implied" inflation $\Delta\ln CPI$	
		δ_1	δ_2	δ_1	δ_2
Belgium	0.0228	0.21	0.72	0.28	0.97
Denmark	0.0132	0.22	0.61	0.17	0.47
Finland	0.0545	0.23	0.51	0.62	1.37
France	0.0379	0.19	0.59	0.42	1.32
Germany	0.0194	0.15	0.56	0.17	0.64
Italy	0.0134	0.20	0.68	0.16	0.54
Netherlands	0.0404	0.21	0.63	0.50	1.50
United Kingdom	–0.0016	0.23	0.60	–0.02	–0.06
Czech Republic	–0.0059	0.10	0.65	–0.04	–0.22

Notes: [a] LP_{tr} is represented here by manufacturing and LP_{ntr} by construction.
[b] δ_1: real value added in manufacturing represents the traded sector Q_t, and real value added in construction the nontraded sector Q_n. δ_2: Q_t = agriculture and manufacturing, Q_n = transport, construction, retailing and financial, social and personal services. Average shares δ are used. In the Czech case, δ_1 is calculated from nominal value added in construction and manufacturing, as the average for the investigated period; δ_2 is the estimated share of nontradables in GDP, as the average for the investigated period.
Sources: See Table 3.

$\ln(LP^i_{tr}/LP^i_{ntr})$, while the third column contains two alternative approximations of the country-specific shares of nontradables in consumption (δ). Using the estimated value $\beta = 0.59$ for these two specifications of δ, we obtain for each country two alternative results for "domestic" or "implied" CPI inflation ($\Delta\ln CPI$); see the fourth column.

As can be seen from Table 5, the countries with the highest productivity differential, such as Finland, France, the Netherlands and Belgium, record the highest "domestic" or "implied" inflation.

Table 5 also shows that, on average, there has not been faster productivity growth in the Czech traded sector. This manifests itself in the negative sign of $\Delta\ln(LP_{tr}/LP_{ntr})$, and, subsequently, determines the sign of $\Delta\ln CPI$. It is, however, more appropriate to say that there is close-to-zero "implied" annual inflation in the Czech Republic, ranging between –0.04 and –0.22 percentage points.

Table 5 presents the results of the standard assumed (long-run) specification of the BSEF, as specified in model (1). The data of unbalanced panel III are used. Owing to the rejection of the commonly assumed restriction (i.e. of model specification (1)), we do not comment on these results in detail and instead deal now with the evidence on the BSEF as presented in *Table 6*.

In contrast with Table 5, in Table 6 the second column includes productivity growth separated by sectors, i.e., $\Delta\ln(LP_{tr})$ and $\Delta\ln(LP_{ntr})$. Two βs are used, in line with model specification (3.2).

When using the above model specification, the country-specific, average contributions to annual domestic CPI inflation vary between approxima-

TABLE 6 Simulated Annual BSEF-implied Impact on Domestic CPI (in percentage points) (model specification (3.2), for β_1 = 0.65 and β_2 = 0.53; unbalanced panel III)

Country	Relative productivity growth		Country share $Q_n/(Q_n+Q_t)$		"Implied" inflation $\Delta \ln CPI$	
	$\Delta \ln LP_{tr}$	$\Delta \ln LP_{ntr}$	δ_1	δ_2	δ_1	δ_2
Belgium	0.0220	–0.0005	0.21	0.72	0.31	1.07[a]
Denmark	0.0377	0.0245	0.22	0.61	0.25	0.70
Finland	0.0611	0.0157	0.23	0.51	0.72[a]	1.60[a]
France	0.0532	0.0152	0.19	0.59	0.50[a]	1.56[a]
Germany	0.0221	0.0026	0.15	0.56	0.19	0.73
Italy	0.0246	0.0112	0.20	0.68	0.20	0.68
Netherlands	0.0347	–0.0060	0.21	0.63	0.53[a]	1.61[a]
United Kingdom	0.0424	0.0439	0.23	0.60	0.09	0.25
Czech Republic	0.0639	0.0697	0.10	0.65	0.05	0.29

Note: [a] Country-specific, BSEF-implied inflation above the sample average.
Source: See Table 3.

tely 1.6 p. p. for the Netherlands (Finland) and 0.05 p. p. for the Czech Republic.

For δ_1, the average annual "implied" or "domestic" inflation for nine countries is 0.31 p. p. and for δ_2 it is 0.94 p. p. Depending on the coefficient δ used, the deviation from the mean value of "implied" inflation (measured by the coefficient of variation) is within the interval (0.14–0.28).

When one considers the relatively low inflation rates in the EU during the 1990s, the "implied" annual inflation rates for particular countries exceeding 0.5 % and even 1 % cannot be overlooked.[24]

This is obviously not the case for the Czech Republic, where the "implied" annual inflation rate, ranging between 0.05 % and 0.3 %, is negligible (actually the lowest within the sample of nine countries), both in absolute terms and with respect to total inflation.[25]

In general, we interpret the results in Table 6 as meaning that during the investigated period the Czech economy did not experience any inflation (real appreciation) pressure due to the existence of the BSEF. In contrast, Belgium, Finland, France and the Netherlands appear to be most affected by the influence of sectorally unbalanced productivity growth on inflation (real exchange rate appreciation).[26]

Table 7 summarizes in the second column the results of calculating the BSEF-implied cross-country inflation differentials in line with model specification (4.2) and also signals the absence of any real exchange rate appreciation pressure for the Czech koruna (with Germany as the reference country).

[24] Kohler (1999) and Weidmann (2002) find for selected EU countries almost identical (or slightly higher) "implied" annual inflation rates.

[25] Using model specification (3.1), with β obtained from country regression and with different data, sector composition and probably also with different (unreported) δ, Mihaljek (2002) obtains for the Czech Republic a magnitude of "implied" inflation of 0.32, which is comfortably close to our result for δ_2 (0.29) in Table 6. One must admit, however, that the above-mentioned differences in econometric approaches complicate the comparability of these results.

TABLE 7 Cross-country BSEF-implied Inflation Differentials (in percentage points) (Germany as the reference country)

Country	"BSEF-implied" cross-country inflation differentials		"BSEF-neutral" labour productivity acceleration		"BSEF-significant" labour productivity acceleration	
	δ_1	δ_2	δ_1	δ_2	δ_1	δ_2
Belgium	0.12	0.34	0.61	0.16	3.94	1.13
Denmark	0.06	–0.03	0.88	0.66	2.74	1.32
Finland	0.53	0.87	0.42	0.30	1.51	0.79
France	0.31	0.83	0.52	0.32	2.04	0.82
Italy	0.01	–0.05	0.96	0.55	4.09	1.46
Netherlands	0.34	0.88	0.26	0.00	2.37	0.96
United Kingdom	–0.10	–0.48	1.14	0.96	2.72	1.56
Czech Republic	–0.14	–0.44	1.35	0.96	3.75	1.33

Now, still using model specification (4.2), we adopt a condition that the annual "domestic" or "implied" inflation is the same in the Czech Republic and in Germany and calculate in the third column the increase in productivity in the Czech traded sector necessary to reach such a value of annual inflation. This approximates the maximum productivity growth in the Czech traded sector which will not bring about a positive cross-country inflation differential (real exchange rate appreciation pressure).

Second, we perform the same exercise assuming that annual "domestic" or "implied" inflation in the Czech Republic is 1 p. p. higher than in Germany.

One has to admit that even in the case of relatively rapid labor productivity growth acceleration in the Czech traded sector (i.e., by 35 %), there would be no BSEF-based impact on the inflation differential (real exchange rate appreciation) against Germany.

The results also show that labor productivity growth in the Czech traded sector would have to be 1.33–3.75 times greater than it actually is to contribute 1 p. p. to the BSEF-implied inflation differential against Germany (i.e. instead of the current annual average rate of labor productivity growth of 6.4 %, it would have to reach a minimum of 8.5 %).

This suggests that the BSEF-based impact on inflation (real exchange rate appreciation) will also remain rather insignificant in the future,[27] should productivity growth in the Czech traded sector not accelerate quite dramatically.

[26] The question obviously remains how would our results be influenced by relaxing the condition of constant nominal exchange rates. According to Mihaljek (2002) a one percentage point increase in the dual productivity differential leads to a 0.15 p. p. increase in the respective CPI inflation differential and to a 0.1% nominal exchange rate appreciation (the respective differentials are measured between the Czech Republic and the Euro area).

[27] The results obtained for the Czech Republic are to a great extent determined by relatively fast average annual productivity growth in construction of around 7 %. One could obviously question whether this is an appropriate assumption for the nontraded sector as a whole. If it is not, our simulation results may be biased even if we have good estimates of β(s). This potential problem is an additional reason for viewing the results cautiously.

5. Conclusion

As we performed our experiments, we always kept in mind the statistical imperfections and model simplifications. Nevertheless, we find that the impact so far of the BSEF on inflation (the real exchange rate) is likely to be very low, if not negligible, in the Czech Republic. We verify such a statement by using two basic model specifications of the problem.

First, we deal with model specification (1), which is shared by the bulk of the literature on this topic. We conclude that a statistically significant impact of relative productivity developments on relative prices does exist in all the investigated countries, even though it is much lower than the BSEF would predict.

In the case of the Czech Republic, one has to note that the difference between the sectoral productivity growth rates is actually very low. This, coupled with a relatively low value of the coefficient β, makes it more appropriate to say that there is a close-to-zero impact of the BSEF on the CPI inflation (real exchange rate appreciation). This is documented by calculations (simulations) in line with model specification (3.1)

Our extension to the standard approach is embodied in model specifications (2), (3.2) and (4.2), where we allow for a more general statement of the problem, which proves to be superior to the standard approach. The resulting impact of relative productivity on inflation (real exchange rate) is close to zero once again.

Our results are generally supported by two still unpublished papers: (Egert, 2002) and (Mihaljek, 2002). Thus, they probably cannot be solely attributed to the simplifications or omissions that we made when testing the presented model. Even when the traded and nontraded sectors are separated in a different manner and different econometric frameworks are used, as in (Egert, 2002) or (Mihaljek, 2002), the estimates of the BSEF for the Czech Republic remain very close to zero.[28]

These recently collected findings differ from those established in the existing literature. For example, according to (Golinelli – Orsi, 2001), the annual contribution of the BSEF to inflation in the Czech Republic is 4.3 % and according to (Sinn – Reutter, 2001) it is 2.88 %.

Halpern and Wyplosz (2001) explore unbalanced panel data for EU accession countries and Russia and find, on average, a 3% "equilibrium" real exchange rate appreciation which can be attributed to the BSEF. Co-

[28] Egert (2002, p. 33) argues that "the equilibrium real exchange rate appreciation [...] may actually have been close to zero in the cases of the Czech Republic, Slovakia and Slovenia, and around 1 % and 3 % for Hungary and Poland, respectively", when the BSEF is adopted as a model for equilibrium real exchange rate appreciation". Mihaljek (2002, p. 18) concludes "[...] productivity differentials vis-à-vis the euro area explain only a small proportion of inflation differentials. Moreover, productivity differentials between tradable and nontradable industries in general seem to explain only a small portion of the domestic inflation in Central European countries." Also Beneš and Klíma (2002), who employed rather provisional simulations, as well as a simple accounting framework for the real exchange rate, find (p. 11) that "decomposing the internal price movements into two subperiods reveals that most recent trends defy the importance of the BS effect altogether".

ricelli and Jazbec (2001) use an even broader unbalanced panel with 19 transition countries and find a “sustainable” real appreciation of about 1 %.[29]

Yet the United Nations (2002, p. 183) claim the presence of “the Balassa-Samuelson effect, which is an equilibrium phenomena and is a fundamental feature of a fast-growing, catching-up economy”.

The above results have been extensively discussed among professional economists and have made the BSEF a fashionable subject. Moreover, they have influenced the framework within which policy makers perceive the magnitude of the “equilibrium” real exchange rate appreciation, as well as the macroeconomic “sustainability” of the real catching-up process. This is not surprising, because these “pro-BSEF” results are clearly of use in two “policy-relevant” directions:

1. In countries with relatively high inflation, they appear to “justify” suggestions to modify the Maastricht inflation criterion because fulfillment of the price stability criterion is allegedly at odds with “real” convergence.
2. Even in low-inflation candidate countries with rapid (real and nominal) exchange rate appreciation, they aspire to explain how much of this process is “sustainable” from the viewpoint of macroeconomic stability and in such a way provide policy makers with important guidance.

In contrast, the most recent findings suggest that the BSEF is a rather poor explanatory variable and that other, as yet less highlighted factors should be tested as determinants of the evolution of the equilibrium real exchange rate, or that the notion of equilibrium itself needs to be redefined.

A partial analogy with our results can also be found in (Kohler, 1999), who uses model specification (1) and reports that the value of “implied inflation” is close to zero for Asian and African developing countries.[30]

Thus, summarising our results and also making reference to (Kohler, 1999), (Egert, 2002) and (Mihaljek, 2002), it seems that the BSEF mechanism works predominantly, if at all, in highly developed countries such as the EU member states or the USA, and perhaps also in fast-growing, catching-up economies such as Hungary or Poland. For the Czech economy, however, the impact of the effect is rather negligible, just as it is in Slovakia, Slovenia and some non-European developing countries.

It follows that when one tries to define and subsequently quantify the factors of “equilibrium” real exchange appreciation in the Czech Republic, one still can doubt that the BSEF really belongs on the list of plausible explanatory variables.

But does it really mean, as, for example, Kovács (2002, pp. 3–4) argues, that “[...] real convergence should not necessarily endanger the fulfillment

[29] Quoted according to Egert (2002)

[30] Making reference to (Kohler, 1999), we can mention, as examples, countries such as China or Zimbabwe, which have recorded very low values of annual implied inflation.

of the Maastricht Treaty Criteria" and the BSEF "might easily become [even] smaller for the future as the catch-up process is more complete"?

The point is that, with regard to real convergence, the Czech Republic did not record any remarkable progress in this direction throughout the 1990s, at least in terms of economic level or total factor productivity indicators.[31]

Therefore, contrary to (Kovács, 2002), the impact of the BSEF should become, in fact, stronger in the future as the catching-up process gathers pace. This should manifest itself, among other effects, in an acceleration of productivity growth in the traded sector.

As our simulations demonstrate, however, even in the case of relatively rapid future productivity growth in the traded sector, the magnitude of the BSEF-based impact on the real exchange rate (or on the CPI-inflation differential against Germany) would hardly exceed 1 p. p., as compared with the current close-to-zero impact. Therefore, the BSEF will probably not be a major explanatory factor for future real exchange rate developments either.

APPENDIX

We assume two sectoral Cobb-Douglas production functions:

$$Y_{tr} = A_{tr}\, L^{\sigma} K^{1-\sigma}$$
$$Y_{ntr} = A_{ntr}\, L^{\gamma} K^{1-\gamma}$$

where Y denotes the output of the traded (tr) and nontraded (ntr) sectors. A is total factor productivity, while K and L are capital and labor inputs. Finally, σ and γ denote the labor shares in the traded and nontraded sectors and $(1-\sigma)$; $(1-\gamma)$ represent the respective capital shares. Assuming perfect factor mobility between the two sectors (in the case of capital also internationally), the profit maximization conditions imply:

$$R_{tr} = (1-\sigma)\, P_{tr} A_{tr}\, L^{\sigma} K^{-\sigma} \qquad \text{(i)}$$
$$R_{ntr} = (1-\gamma)\, P_{ntr} A_{ntr}\, L^{\gamma} K^{-\gamma} \qquad \text{(ii)}$$
$$W_{tr} = \sigma P_{tr} A_{tr}\, L^{\sigma-1} K^{1-\sigma} \qquad \text{(iii)}$$
$$W_{ntr} = \gamma P_{ntr} A_{ntr}\, L^{\gamma-1} K^{1-\gamma} \qquad \text{(iv)}$$

where R is the interest rate, W represents the wage rate, and P_{tr} and P_{ntr} stand for prices in the traded and nontraded sectors. Log-differentiating (i)–(iv) yields the following relation:

$$\ln P_{tr} - \ln P_{ntr} = c + \ln(\gamma/\sigma) \ln A_{ntr} - \ln A_{tr} \quad \text{(v)}$$

This standard theoretical framework cannot be easily tested. Since all the studies at our disposal have used average labor productivity (Y/L) instead of total fac-

[31] See (Nachtigal – Tomšík – Votavová, 2001), (United Nations, 2002), (Flek – Hájek – Hurník – Prokop – Racková, 2001) and (Hájek – Hurník – Hrnčíř, 2002) for detailed empirical analysis.

tor productivity (A), we derive our "testable" hypothesis for average labor productivity as well. Expressing in terms of average labor productivity, we obtain, similarly to Kohler (1999), the following relation:

$$\ln P_{tr} - \ln P_{ntr} = \ln(\gamma/\sigma) + \beta_1 \ln LP_{ntr} - \beta_2 \ln LP_{tr} \quad \text{(vi)}$$

where $\ln(\gamma/\sigma)$ = const. and $\beta_1 = \beta_2 = 1$. See equation (1) in the text.

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Sectoral Productivity and Real Exchange Rate Appreciation: Much Ado about Nothing?

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Using panel data for selected national economies, we estimate relative price changes stemming from fluctuations in sectoral productivity. Subsequently, we calculate the cross-country CPI-inflation differentials implied by sectorally unbalanced productivity growth, taking into account country-specific weights of nontradables in consumption (value added) and assuming there are no adjustments in nominal exchange rates. We find that sectoral productivity developments have a statistically significant impact on relative prices in the EU countries and also in the Czech Republic, but the magnitude of the impact is not as strong as the Balassa-Samuelson Effect (BSEF) would predict. The final impact of relative productivity on inflation (on the real exchange rate) is even weaker and, moreover, in the case of the Czech Republic the impact is negligible. Thus, contrary to the prevailing view, we question the meaning of the BSEF as a plausible explanatory variable of (equilibrium) real exchange rate determination in the Czech Republic. The same situation we simulate for the future, provided productivity growth in the traded sector does not accelerate dramatically.