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Competitiveness in Transition Economies: What Scope for Real Appreciation?

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We estimate equilibrium dollar wages for 15 transition economies of Central and Eastern Europe (CEE) and the former Soviet Union. Equilibrium dollar wages are interpreted as full employment wages consistent with a country's physical and human capital endowment, and estimated by regressing actual dollar wages on productivity and human capital proxies in a short (1990–95) panel of 85 countries. The main results are: (1) equilibrium dollar wages have appreciated steadily in the Baltic countries and fast-reforming CEE transition economies, but have been flat in most CIS countries; and (2) 1996 actual dollar wages remain below estimated equilibrium dollar wages for most but not all transition countries covered. [JEL F14, F21, F41, P20, P50]

In the last few years, the currencies of the transition countries of Central and Eastern Europe (CEE) and the former Soviet Union have undergone large real appreciations (Figures 1 and 2).[1] According to a widespread consensus, the process of appreciation set in following an initially strongly

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[1]The figures show consumption-based real exchange rate indices vis-á-vis the U.S. dollar since January 1990 for Hungary and Poland, January 1991 for the remaining Central and Eastern European countries, and January 1992 for the Baltic and Commonwealth of Independent States countries except for the Kyrgyz Republic and Kazakhstan, for which the date of introduction of the national currency was chosen.

Figure 1. *Real Exchange Rate Indices Versus the U.S. Dollar: Centeral and Eastern Europe*

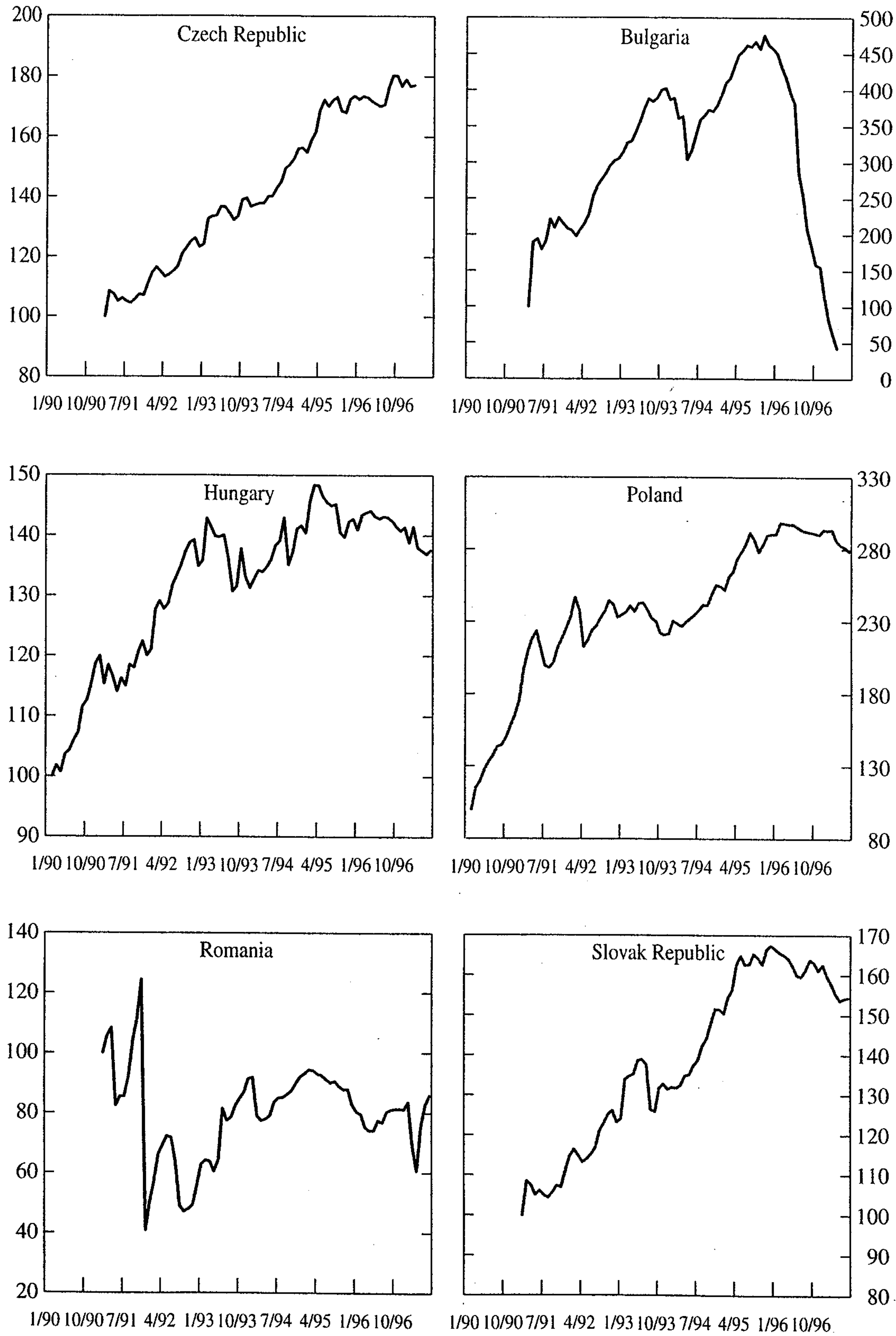
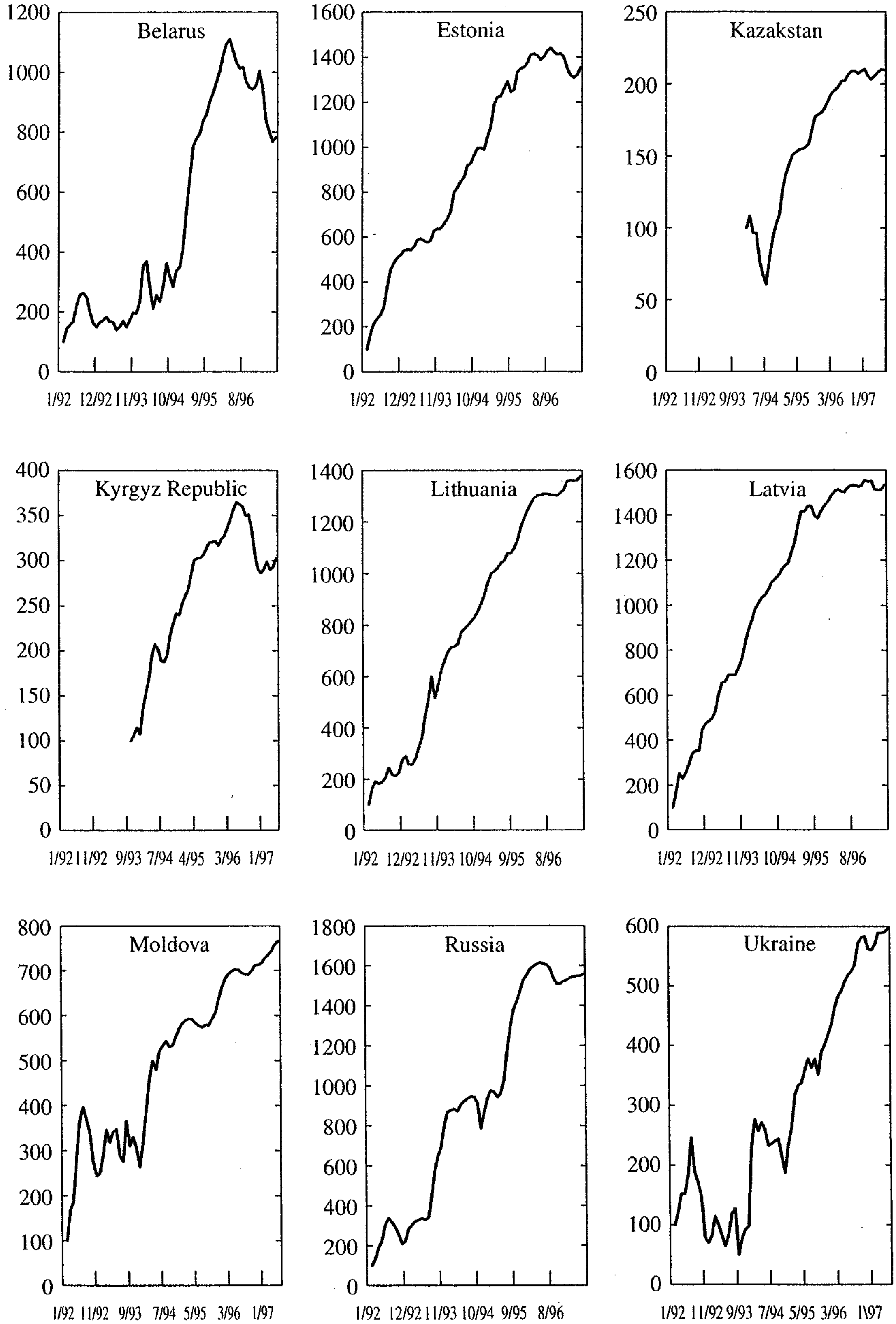


Figure 2. *Real Exchange Rate Indices Versus the U.S. Dollar: Baltic and CIS Countries*



undervalued position of these currencies.[2] Accordingly, it was not perceived to be threatening to the competitiveness positions of these countries. In view of the magnitude of real appreciations witnessed—up to fifteen-fold in some CIS economies, including Russia, since early 1992—this can no longer be taken for granted. The Czech currency crisis in May 1997, which led to a devaluation of the koruna by 8.5 percent following a large widening of the current account deficit in 1996 (Table 1), has focused new attention on the question of whether the scope for real appreciation has by now been exhausted in some or most transition economies. If the answer is yes, this would clearly have implications for the design of exchange rate policy, which in many countries so far has been guided primarily by the objective of reducing inflation, rather than maintaining competitiveness.

The objective of this paper is to obtain a sense of what scope, if any, remains for real appreciation in transition economies of Central and Eastern Europe and the former Soviet Union before competitiveness becomes an issue. At the outset, one might wonder whether the recent current account positions in these countries offer any information in this regard. While some countries have been in surplus—particularly energy exporters—most were in deficit (Table 1). But the latter is just what standard intertemporal equilibrium models would predict for countries that are rebuilding their capital stocks following a large structural shift. More generally, at a time when these economies are adapting to large relative price shocks, including terms of trade shocks, major trade disruptions, and other institutional changes affecting the trade regime, any inference about the appropriateness of the real exchange rate based on the current account position seems even more difficult than usual.[3]

What methodology could be usefully employed to decide whether the currencies of transition economies should by now be considered overvalued? The literature on competitiveness and real equilibrium exchange rates suggests a variety of approaches. Competitiveness in a given year is sometimes assessed by comparing the value of a real effective exchange rate (REER) index in that year with its value in some reference year in which

[2]See, for example, Richards and Tersman (1996) and Halpern and Wyplosz (1997), among others.

[3]This said, for countries that have completed major structural reforms in the external sector, it may be possible to gain information on current account sustainability by studying the composition of the balance of payments and the saving-investment balance on a case-by-case basis. This is the approach taken by Roubini and Wachtel (1998), who conclude that loss of competitiveness and current account sustainability problems may have arisen in a number of transition economies after 1995.

Table 1. *Current Account Balances*
(In percent of GDP)

Country	1992	1993	1994	1995	1996
Czech Republic	–1.5	2.0	–0.1	–2.7	–8.1
Bulgaria	–9.3	–12.8	–2.1	–0.5	0.5
Hungary	0.9	–9.0	–9.5	–5.6	–3.8
Poland	1.9	–0.1	2.3	3.3	–1.0
Romania	–7.8	–4.7	–1.7	–4.9	–6.6
Slovak Republic	–0.4	–5.4	4.8	2.3	–10.2
Belarus	5.5	–30.3	–13.1	–2.4	–6.6
Estonia	0.9	1.5	–7.1	–5.3	–8.2
Kazakhstan	–51.4	–9.4	–11.6	–4.2	–3.7
Kyrgyz Republic	–10.6	–16.1	–11.2	–19.3	–21.8
Latvia	1.8	7.0	–2.4	–3.5	–6.8
Lithuania	11.1	–4.6	–3.1	–4.4	–4.4
Moldova	–4.5	–13.4	–6.9	–8.6	–13.0
Russia	–1.4	1.4	3.7	1.1	1.7
Ukraine	–3.0	–5.9	–5.7	–4.2	–2.7

Source: International Monetary Fund.

the economy is regarded as being in external and internal equilibrium.[4] Alternatively, an equilibrium real exchange rate can be estimated, and compared to the actual real exchange rate. This is typically done either by first estimating or assuming an "equilibrium current account" and then estimating the real exchange rate that would generate it, or by estimating a reduced-form model in which the equilibrium real exchange rate is identified with the long-run real exchange rate that is associated with steady-state net foreign assets (NFA) and current account positions.[5]

The first of these approaches is difficult to apply to transition economies since it implicitly assumes a constant real equilibrium exchange rate over time and as such does not take account of changes in productivity, capital stocks, tastes, or commodity prices, which would, in general, imply a change in the equilibrium rate. Such changes presumably matter in the context of transition economies undergoing rapid structural transformation. Even more important, the REER index approach requires a reference year for which equilibrium of the real exchange rate can be assumed. No year before the beginning of transition can be taken as a reference, since trade and capital flows in that period were heavily restricted. On the other hand, the initial period of external opening and reform is usually associated with

[4]Feldman (1995) is an example for an application; Marsh and Tokarick (1994) and Lipschitz and McDonald (1992) discuss some theoretical underpinnings.
[5]See Clark and others (1994), Faruqee (1995), Debelle and Faruqee (1996), Kramer (1996), and Isard and Faruqee (1998).

very large real depreciations. In general, one would thus assume that transition economies have gone from artificially appreciated currencies—such as the official ruble exchange rate of 1.7 to the dollar in 1991—to undervalued currencies following currency convertibility and the initial floating or pegging of the exchange rate. There is no discernible state of equilibrium between these two states that could serve as a reference.

Unfortunately, the alternative and more sophisticated econometric approaches described above are not feasible either. The joint estimation of long-run equilibrium exchange rates and the current account or NFA positions is precluded by the absence of adequate time-series data for these countries, with only 2 to 5 years of data since the beginning of the transition process. The two-step approach, on the other hand, requires the estimation of real exchange rate elasticities of the current account. Even ignoring the fact that this estimation typically occurs in a time-series context and is thus subject to the same data limitations, it would seem difficult, if not impossible, to estimate the effects of real exchange rate movements on the current account at a time when fluctuations in exports are likely to be driven primarily by such changes as the removal of export quotas, the breakdown of traditional trading blocks, and changes in relative prices within the tradables sector.

The solution we propose to overcome these problems rests on two ideas. First, we use a real exchange rate measure that is both readily available for transition economies and—unlike REER indices—can be directly interpreted and compared in levels, namely, dollar wages in the manufacturing sector. Second, we estimate equilibrium dollar wages as a function of productivity measures using a short panel of countries, rather than a time series. Thus, the estimated equilibrium dollar wage represents an estimate of what the country could "afford" based on its stock of human and physical capital. We then go on to interpret competitiveness as the gap between actual dollar wages and estimated equilibrium dollar wages.

The approach pursued is inspired by the way in which macroeconomic practitioners often form a judgment of the international competitiveness status of a country—by comparing the country's average dollar wage with that of other countries that are considered "similar" in terms of the remaining determinants of profitability or unit cost, such as the quality and quantity of human and physical capital. Which countries are to be considered "similar" in this sense is usually decided ad hoc. We put this informal comparison on a more systematic footing by constructing a fictitious country with identical human and physical capital to the country we are assessing—as measured by crude proxies, which will be discussed in detail below—and estimating the dollar wage one would expect to prevail in this country.

Since one of the productivity measures on the right-hand side of our wage equation is (PPP-adjusted) GDP per employee (or per member of the labor force, or per capita), our approach encompasses competitiveness comparisons based on aggregate unit labor cost (see Havlik, 1996, who compares unit labor costs in Central and Eastern European transition economies with that of Austria). It is also related to a large literature comparing international price levels and relating deviations from purchasing power parities to differences in resource endowments across economies.[6] Unlike this literature, however, our left-hand side variable is dollar wages, not prices or price indices. Apart from the desire to stay close to the terms in which the discussion on competitiveness is led by practitioners, there are two reasons for this. First, the link from structural determinants to national price levels usually runs via factor prices, the argument typically being that factor prices are an important determinant of the price of nontradables relative to tradables, which in turn is an important determinant of the price level.[7] As a result, specifying a country-invariant structural equation relating resource endowments to price levels that can be estimated from observable variables seems even harder than specifying a corresponding regression model of dollar wages. Second, even if we wanted to estimate real equilibrium exchange rates for transition economies based on a comparison of national price levels, we would have been constrained by data availability, as the International Comparison Program, on which price level comparisons are typically based, was not extended to all of these countries at the time of this analysis.

Finally, we owe much to a recent paper by Halpern and Wyplosz (1997), who pursue broadly similar objectives for a somewhat different set of transition economies that does not include the Baltic and CIS economies studied here (except for Russia).[8] The basic similarity between Halpern and Wyplosz (1997) and this paper is the attempt to estimate equilibrium dollar wages in transition economies using a set of productivity proxies as right-hand side variables. However, the estimation approaches and data sets are different. Halpern and Wyplosz argue that to uncover the relationship between fundamentals and equilibrium dollar wages one needs to "observe each country for a long period of time" and consequently use a long panel (1970–90), which includes the planned economies (with a planned economy dummy) and a time trend. We argue that, since we are interested in equilibrium dollar wages as opposed to long-run steady-state wages, it is enough to use a cross section (or alternatively a short panel to test for coun-

[6]See Clague and Tanzi (1972), Isenman (1980), Clague (1986 and 1988), Kravis and Lipsey (1987 and 1988), Officer (1989), and Dollar (1992), among others.

[7]For example, Kravis and Lipsey (1988), and Dollar (1992).

[8]On the other hand, Halpern and Wyplosz include Slovenia, Croatia, and China, which are not in our country set.

try specific effects), provided the countries in our sample have been market economies for a sufficiently long time that we can assume that on average they are in equilibrium. On this basis, we use a short panel (1990–95) that includes 70 market economies and 15 transition economies with a transition dummy to capture out-of-equilibrium effects for this group (see Table A1). Apart from allowing us to estimate equilibrium wages for the Baltic and CIS countries, for which long panel data do not exist, this enables us to estimate equilibrium wages for transition economies within sample (1991 to 1995) rather than out of sample, as is the case for Halpern and Wyplosz. In spite of these differences, the main results of the two papers are fairly close for the countries they both study. Both papers suggest that while the gap between actual and equilibrium dollar wages declined in most countries since the beginning of transition, it remained substantial in 1995–96 in most (but not all) cases. For example, for Russia, Halpern and Wyplosz estimate equilibrium dollar wages of 400–500 dollars for 1995. Our estimate is lower (235–394 dollars, depending on the specification and data used), but this is still substantially higher than actual dollar wages in manufacturing during 1995 and 1996 (107 and 188 dollars, respectively).

I. Methodology

Basic Approach

In attempting to estimate equilibrium dollar wages for transition economies, we face two difficulties. First, as emphasized in the introduction, the virtual absence of time-series data for these economies; and second, the need to make inferences about equilibrium wages based on the observation of *actual* wages, which might be far off equilibrium. Our proposed solution to this problem is to estimate equilibrium dollar wages using a cross section (or alternatively—to deal with country-specific fixed effects—a short panel) of countries, including many nontransition economies. This is briefly justified as follows.

Suppose country i is a market economy. By this we mean that market forces have been determining prices and wages in country i for a long period of time (say, ten years). Suppose we knew the equilibrium dollar wage of country i. Then, assuming we have no further information about country i, our best guess of *actual* dollar wages in this country would be the equilibrium dollar wage:

$$E\left[w_{\$,i} \mid w^*_{\$,i}\right] = w^*_{\$,i} \,. \tag{1}$$

Now suppose we do not know $w^*_{\$,i}$ but we have a theory of how it is determined:

$$w^*_{\$,i} = G(Z_i), \tag{2}$$

where Z_i denotes a vector of observable "fundamentals" in country i. If we can also observe the actual dollar wages for each country, G could be *estimated*. Combining equations (1) and (2), we obtain

$$w^*_{\$,i} = G(Z_i) + \varepsilon_i, \tag{3}$$

with $E[\varepsilon_i \mid Z] = 0$. Thus, equilibrium dollar wages could be estimated by regressing actual dollar wages on Z_i in a cross section of countries for which equation (1) holds and computing the fitted wage for each country.

Three issues remain to be addressed. First, the above assumes that we know the model $G(Z_i)$. We thus need to propose such a model. Second, we need to find observable proxies for Z_i. Finally, we need to decide how to treat observations from transition economies, for which equation (1) cannot be assumed. The next section deals with the first of these issues, while the other two will be taken up in the sections that follow.

A Model of Equilibrium Dollar Wages

Short-Run Equilibrium

Consider a simple two-sector model of equilibrium dollar wages and the real exchange rate. We make two key assumptions. First, the tradables sector is assumed capital intensive relative to the nontradables sector. This is necessary and sufficient to generate a positive relationship between the relative price of nontradables and dollar wages and thus justify our use of dollar wages as a measure of the real exchange rate. For simplicity, take the special case in which we have a Ricardian technology in the nontradables sector:

$$Y_N = \delta L_N \tag{4}$$

and a Cobb-Douglas technology in the tradables sector:

$$Y_T = \gamma K^\alpha L_T^{1-\alpha}. \tag{5}$$

Assuming that the international (i.e., dollar) price of tradables is given, so that the domestic price of tradables equals the exchange rate ($P_T = E$, where E is defined as domestic currency per dollar), equilibrium in the nontradables sector then implies a linear relationship between the real exchange rate and dollar wages:

$$P_N = \frac{W}{\delta}; \text{ or } \frac{P_N}{P_T} \equiv P = \frac{1}{\delta}\frac{W}{E} \equiv \frac{1}{\delta} W_\$. \tag{6}$$

Our second key assumption is to assume a cost to capital adjustment, so that capital can be treated as given in the short run. This enables us to separate short-run equilibrium—that is, equilibrium dollar wages conditional on the level of the capital stock, which is all we need to estimate equation (3) in a cross section—from the dynamics of capital accumulation and long-run steady state. Tradables profit maximization leads to

$$\frac{W}{E} \equiv W_\$ = \gamma\,(1-\alpha)\left(\frac{K}{L_T}\right)^{\alpha} . \tag{7}$$

Assume now that labor is in fixed domestic supply $\overline{L}$ and mobile across the two sectors. Using the fact that all nontradables production is necessarily consumed at home, labor market clearing implies

$$W_\$ = \gamma\,(1-\alpha)\left(\frac{K}{\overline{L} - \frac{1}{\delta}C_N}\right)^{\alpha} \tag{8}$$

To close the model, we need to make some assumption about the consumption side. The simplest possibility is to assume that capital is owned by foreign investors and domestic workers-consumers are concerned with static utility maximization only, that is, they do not save and they only worry about how to allocate each period's wage bill across the two sectors. Assuming Cobb-Douglas preferences with expenditure shares β for nontradables and $1-\beta$ for tradables, this leads to

$$C_N = \frac{\beta W \overline{L}}{P_N} = \beta\delta\overline{L} \tag{9}$$

Substituting in equation (8) then gives us an expression for wage determination in general equilibrium:

$$W_\$ = \gamma\,(1-\alpha)\left(\frac{K}{(1-\beta)\overline{L}}\right)^{\alpha} . \tag{10}$$

In its log-linear version, this equation says that equilibrium dollar wages depend on a constant (which might be different across countries if technology is different) and the aggregate capital-labor ratio. The elasticity of dollar wages with respect to capital is the (constant) capital share α.

The price of simplicity in this way of closing the model is that—because we are not allowing consumers to accumulate debt or save—we are imposing current account balance. This is probably too strong: most economists would agree that the notion of "external equilibrium" defining the equilibrium dollar wage need not require current account balance at all times. Instead, it only requires the current account deficit or surplus to be consistent with an intertemporally optimizing, sustainable consumption path. In the simplest intertemporal extension of the case considered before, consumers solve

$$\max \sum_{t=0}^{\infty} \frac{1}{1+\rho} \ln \left(C_{N,t}^{\beta}\, C_{T,t}^{1-\beta} \right) s.t. \sum_{t=0}^{\infty} \frac{1}{1+r} \left(C_{N,t}\, P_t + C_{T,t} - W_{\$,t}\, \bar{L} \right) = 0. \quad (11)$$

The proportion of nontradables and tradables in intratemporal consumption is given (as before), as $C_{N,t}=\beta/(1-\beta)\ C_{T,t}/P_t$. In the simplest case, when $r=\rho$ and the intertemporal consumption profile is flat,[9] equation (8) becomes

$$W_{\$} = \gamma\,(1-\alpha) \left(\frac{K}{\bar{L} - \frac{\beta}{1-\beta} \bar{C}_T / W_{\$}} \right)^{\alpha}. \quad (12)$$

As in equation (8), the equilibrium dollar wage will depend on technology, capital and the labor force, but the simple linear dependence of $W_{\$}$ on the log of $K / \bar{L}$ disappears. This is because labor is no longer allocated across tradables and nontradables sectors in a fixed proportion. If the capital stock increases, dollar wages and nontradables prices rise, nontradables consumption declines and a larger proportion of the labor force moves to the tradables sector. This generates an offsetting effect that dampens the appreciation of $W_{\$}$.[10]

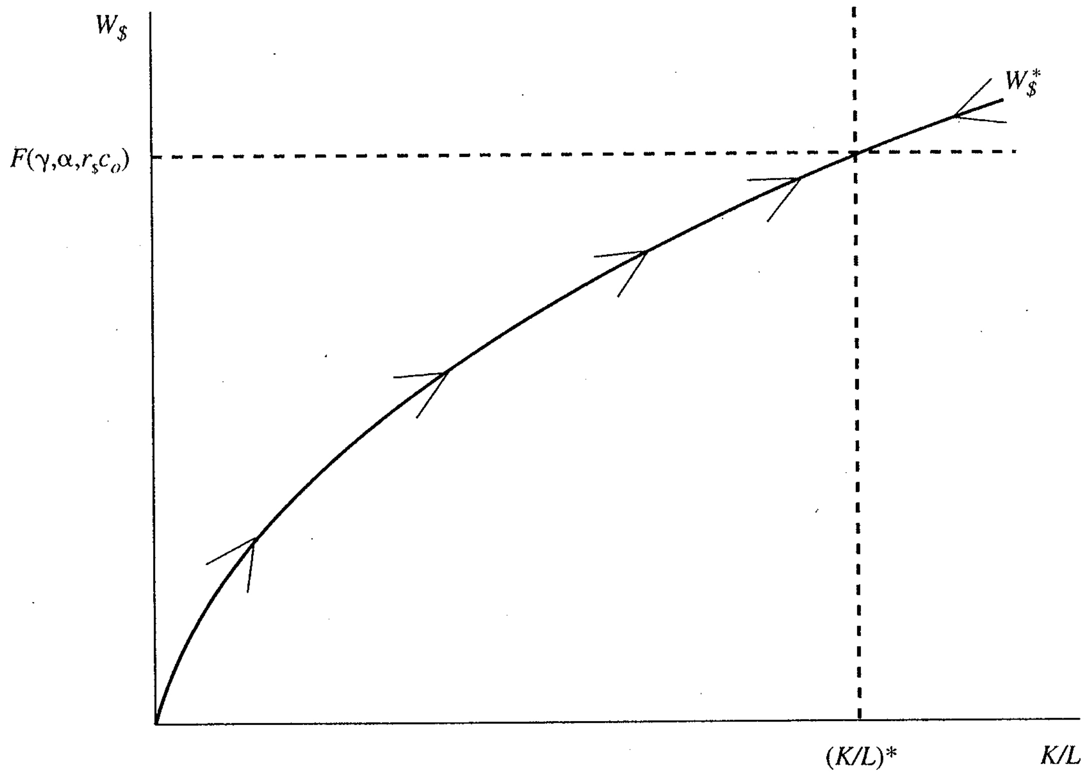
Dynamics and Long-Run Steady State

Equations (10) and (12) define relationships between equilibrium wages and fundamentals along the lines of equation (2), which can in principle be estimated using actual dollar wage data, as argued above. However, before turning to issues of empirical implementation it is instructive to take the model one step further to see what it implies about the dynamics of real exchange rates in transition economies. Note that nothing in our empirical approach, which we return to below, depends on the particular view taken in this subsection.

[9]See Dornbusch (1983).

[10]It is easy to show that the elasticity of $W_{\$}$ with respect to capital is always smaller than α. As $W_{\$}$ rises and labor moves into the tradables sector, the elasticity increases and reaches α in the limit.

Figure 3. *Example of Dollar Wage Dynamics with Inflation Inertia*



Equations (10) and (12) suggest that the dynamics of the equilibrium dollar wage will be determined by differential productivity improvements in the tradables and nontradables sector and capital accumulation in the tradables sector. If one abstracts from the former for the time being (i.e., treats the technology parameters as fixed), combining the previous section's assumptions about production and consumption with a standard neoclassical growth model will thus deliver real exchange rate dynamics.[11] Capital will be accumulated as long as the dollar profit from installing an extra unit of capital—which, inter alia, depends on the prevailing dollar wage—exceeds the unit installation cost times the international interest rate. As capital is installed, the marginal product of labor rises and equilibrium dollar wages increase. Thus, the adjustment of equilibrium dollar wages to their steady-state level can be depicted as a movement along the curve defined by equations (10) or (12) (see Figure 3). The steady-state level of equilibrium dollar wages itself will depend on the technology parameters of the model, the international

[11]One straightforward way of doing this is to combine equation (10) with a standard q-model of investment in the tradables sector, as was done in a previous version of this paper (available on request). More generally, it is possible to write down a two-sector open economy Ramsey model with costly capital adjustment that embodies the production structure presented above, along the lines of Obstfeld and Rogoff (1996), pp. 260–63.

interest rate, and installation costs. In particular, higher productivity in the tradables sector will imply higher steady state dollar wages.

From equation (6), Figure 3 equivalently traces out the dynamics of the real equilibrium exchange rate for given productivity parameters. In steady state, the "pure" Balassa-Samuelson mechanism reemerges: long-term trends in the real exchange rate will be driven by differential rates of (total factor) productivity growth in the tradables and nontradables sectors; faster productivity growth in the tradables sector generates a real appreciation. During the adjustment to steady state, however, there is an additional force behind real appreciation, namely capital accumulation in the tradables sector. If productivity of the service sectors rises faster than tradables productivity during the adjustment to steady state, we have two opposing forces acting on the real exchange rate. Real appreciation will still result if capital accumulation causes dollar wages to outpace relative productivity gains by the nontradables sector. Equilibrium dollar wages, on the other hand, will unambiguously rise during adjustment as long as productivity in the tradables sector does not decline.

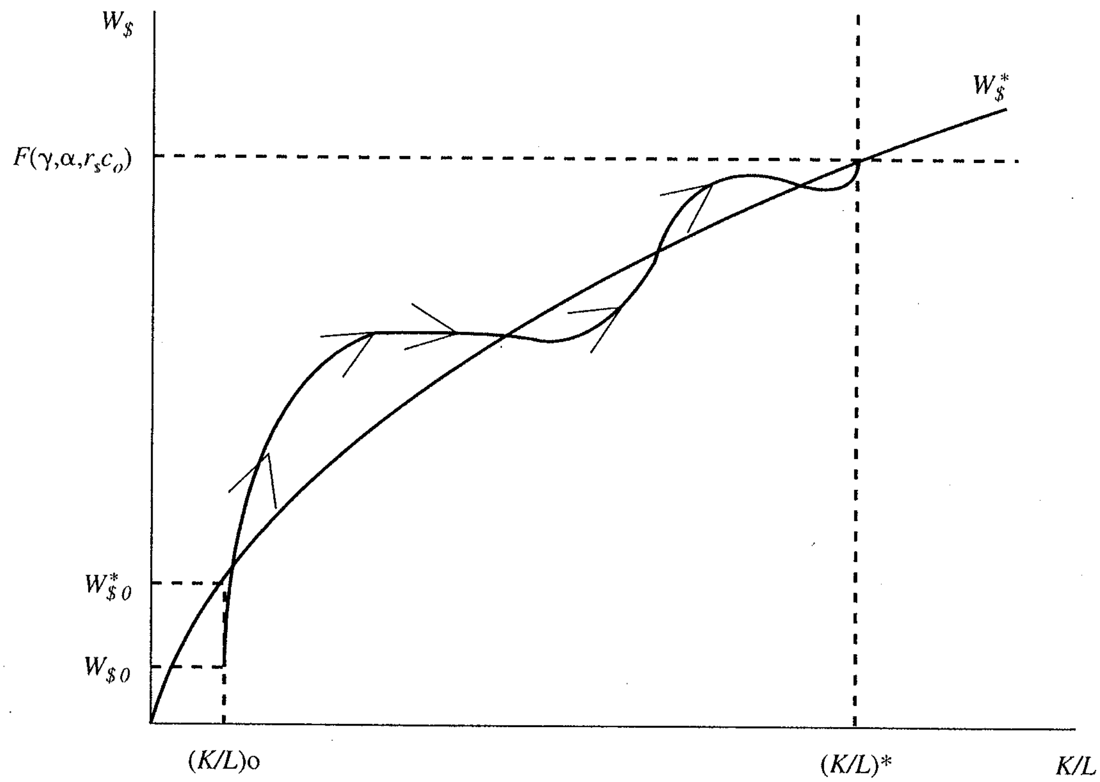
The discussion so far has focused on equilibrium real exchange rates and dollar wages because it was based on a market clearing, fully optimizing current account model. However, the equilibrium model may not provide a good description of *actual* short run dollar wage and real exchange rate movements for well-known reasons. For example, in a flexible exchange regime, exchange rates might be driven by external borrowing and portfolio investment in addition to capital investment; this could generate swings in the nominal exchange rate that, in the presence of short-run wage rigidities, will feed through to dollar wages. In a fixed exchange rate regime, on the other hand, the real exchange rate may become misaligned if there is inflation inertia, that is, if price or wage growth depends to some extent on past price or wage growth. In this case, wages could exhibit dampened oscillations around the equilibrium adjustment path, with unemployment arising whenever actual dollar wages are above their equilibrium levels for any given level of the capital-labor ratio (see Figure 4).[12]

The initial overshooting of the equilibrium dollar wage path will occur either if there is some inflation in the system to begin with, or—even in the absence of initial inflation—if the dollar wage was initially undervalued, as shown in Figure 4. At least one of these conditions is likely to apply in transition economies that fix exchange rates at the beginning of transition.

In summary, the basic stylized fact documented at the beginning of the paper—a sharp real appreciation since the inception of transition in practi-

[12] See Krajnyák and Zettelmeyer (1997), Appendix A, for a simple model that can generate this pattern.

Figure 4. *Example of Dollar Wage Dynamics with Inflation Inertia*



cally all transition economies—could be interpreted as follows. At the beginning of the transition, real exchange rates in transition economies are (1) below their steady-state levels and (2) undervalued, that is, below their equilibrium levels conditioning on existing levels of profitable capital (point $W_{\$0}$ in Figure 4). The former is a result of the capital obsolescence effect associated with external opening and price liberalization; the latter can be thought of either as consequence of an initial monetary overhang or initial capital flight/capital outflows that are not captured by the intertemporal current account model. As transition proceeds, new capital is accumulated in the tradables sectors, leading to an appreciation of equilibrium dollar wages and real exchange rates. Actual real exchange rates appreciate in the direction of this moving target, but in the presence of inflation inertia or capital-account driven appreciation there is no guarantee that this process will stop once equilibrium real exchange rate/equilibrium dollar wage levels have been reached.[13] Several years into transition, this raises the question whether dollar wages have by now overshot equilibrium or not.

[13]See Calvo, Sahay, and Végh (1996) for a documentation of the link between capital inflows and real appreciation in Central and Eastern European transition economies.

Empirical Specification

We now return to the problem of estimating equilibrium dollar wages on the basis of observable economic fundamentals at any given point in time. Equation (10) or (12) suggests that equilibrium dollar wages should depend on the capital share and equilibrium (i.e., full employment) productivity in the tradables sector, which in turn depends on total factor productivity in the tradables sector, the (equilibrium) capital labor ratio, consumption preferences, and possibly wealth (as a determinant of the consumption level). On this basis, it should be possible to run a regression along the lines of equation (3). Before we do so, we need to address some problems of empirical implementation.

Measuring Productivity

The first challenge is to find a set of observable right-hand side variables consistent with equations (10) or (12). Whether we pick measures that proxy individual variables or parameters on the right-hand side of equations (10) and (12) (such as total factor productivity and the capital labor ratio) or capture a combination of variables (such as tradables productivity) is a matter of empirical convenience, as we are not trying to isolate the effect of individual "fundamentals" on equilibrium wages. The approach we take follows Halpern and Wyplosz (1997) in using a wide set of relatively crude productivity measures or determinants, namely, (1) normalized PPP-adjusted GDP as a broad productivity measure, (2) a schooling or human capital variable, (3) the share of agriculture in GDP as a general proxy for economic development, and (4) a dummy for OECD membership, also as a proxy for economic development.[14] In addition, we try to include various indicators capturing institutional factors that potentially influence productivity (such as property rights). We justify this procedure as our best hope

[14] Unlike Halpern and Wyplosz, we do not consider inflation and the share of government consumption in GDP as determinants of total factor productivity or the marginal product of labor. As to inflation, the model suggests that equilibrium dollar wages should be determined by real factors only. Concerning the government share, the growth literature's familiar rationale for relating it to total factor productivity is unlikely to hold in our sample. Typically, it is assumed that higher government expenditure indicates better infrastructure and hence it points towards higher total factor productivity. Arguably, the inclusion of formerly centrally planned economies destroys the monotonicity of the relationship. First, a relatively large government share in these economies may merely indicate a difference in the ownership structure relative to market economies, and not necessarily a superior infrastructure. Second, countries more advanced in the transition process are likely to have downscaled the size of the government. In our view, these considerations warrant the exclusion of government share from the determinants of total factor productivity.

of proxying for productivity fundamentals in the tradables sector in view of (1) the unavailability of sector-specific productivity data for the tradables sector in our countries; and (2) the need to proxy *equilibrium* (i.e., full employment) productivity on the basis of observable measures, as follows.

Suppose we had data on equilibrium output in the tradables sectors across countries, valued at a set of average international prices. For simplicity, assume further that in equilibrium there is an unknown fixed ratio between the sizes of the tradables and nontradables sector. For example, take the case of equation (10). Denoting the average international price of tradables by I_T, we would have:

$$I_T Y_T^* = I_T \gamma K^{\alpha} L_T^{*(1-\alpha)} \tag{13}$$

$$L_T^* = (1-\beta)\bar{L} \tag{14}$$

Then, dividing equilibrium output at international prices by the labor force would result in a perfect measure of tradables productivity:

$$\frac{I_T Y_T^*}{\bar{L}} = I_T \gamma \left(\frac{K}{(1-\beta)\bar{L}} \right)^{\alpha} \equiv gdp_T^*. \tag{15}$$

From equation (10), $W_\$^* = (1-\alpha) / I_T \cdot gdp^*_T$. Thus, if equation (1) holds, running a regression of actual wages on gdp^*_T in a cross section or panel of countries will be sufficient to estimate equilibrium dollar wages. Note that this procedure would only assume that α is the same across countries; preferences and all remaining technology parameters could differ.

In fact, gdp^*_T is unavailable, but we do have PPP-adjusted GDP, that is, *aggregate* GDP at international prices.[15] Abstracting from the problem that this may not be measured at labor market equilibrium, dividing PPP-adjusted equilibrium GDP by the labor force yields

$$gdp^* \equiv \frac{I_T Y_T^* + I_N Y_N^*}{\bar{L}} = gdp_T^* + I_N \delta \beta, \tag{16}$$

[15]This is the Geary-Khamis (GK) definition of PPP-adjusted GDP; see Wagner (1995). In practice, this is not the only method used to estimate the PPP-adjusted GDP series we employ (see the data section below), but we assume that the estimates we use are close enough to what would have been obtained if application of the GK method had been feasible throughout.

where I_N denotes the international average price of nontradables. Using equation (10), we obtain:

$$W_{\$}^{*} = \frac{(\alpha-1)I_N\,\delta\beta}{I_T} + \frac{(1-\alpha)}{I_T}\,gdp^{*}. \tag{17}$$

It follows that a regression of actual dollar wages on *gdp** and a constant will only lead to consistent estimates of equilibrium wages under implausibly strong assumptions, such as cross-country equality not just of α but also of β and δ, that is, preferences and nontradables productivity. However, not only are β and δ likely to differ across our sample, but they are probably correlated with *gdp** (e.g., both tradables and nontradables productivity might depend on the stage of economic development). Moreover, PPP-GDP is not generally measured at full capacity (or equilibrium) but based on actual output and thus incorporates some degree of cyclical variation.

With these problems in mind, we adopt the Halpern-Wyplosz approach outlined at the beginning of this subsection. To the extent that cross-country differences in preferences and nontradables productivity are correlated with differences in economic development and economic and political institutions, additional regressors reflecting these differences should help control for differences in β and δ. This motivates the inclusion of variables such as human capital, agriculture share, OECD membership, and institutional indicators. On the one hand, these are likely to be directly related to productivity. On the other hand, they may also contain information about the relative size of nontradables consumption.

Finally, to minimize the possible distortions arising from cyclical variation in *gdp**, we explore the implications of different normalizations of PPP-adjusted GDP, using not just the labor force but also population and total employment in the denominator. The justification for this procedure will be briefly discussed in Section II below.

Country-Specific Effects

Based on the discussion so far, the error term in a regression of dollar wages on the variables suggested above may contain a country specific component. First, we might obviously still miss some country-varying determinants of equilibrium wages. In this case, running OLS on a cross section could generate a misspecification problem as these country varying determinants will probably be correlated with some of the right-hand side variables.[16] Second, even assuming that our right-hand side variables fully

[16] See, for example, Keane and Runkle (1992).

account for equilibrium wages, disequilibrium models along the lines of the example illustrated in Figure 4 suggest that the extent of disequilibrium at any given point in time could depend on the country's initial out-of-equilibrium position (or on large shocks or regime changes that have placed the country far out of equilibrium later in its history). In general, we have no information on this initial out-of-equilibrium position (the major exception being the transition economies, see below), which may or may not be correlated with the country's "fundamentals." If it is, we are back to the fixed-effects problem described previously. If it is not, $E[w_{\$,i}|w^*_{\$,i}] = w^*_{\$,i}$ will still hold—in other words, the country-specific effects will be random—and OLS will give us unbiased and consistent estimates. However, allowing for a non-time-varying country-specific component in the error term may still be a good idea from the point of view of efficiency.

To address the potential presence of country-specific effects, we run our regression on a short panel rather than a cross section. Apart from increasing our data set, this allows us to test for the presence of fixed effects and random effects and apply the appropriate estimator. The tests will be discussed below.

Transition Dummies

Recall the argument underlying assumption (1), namely, that countries have undergone a sufficiently long adjustment period in a regime governed by market forces to make it a priori impossible to guess whether their wages are undervalued or overvalued. This assumption is clearly violated for transition economies, which have only recently become subject to market forces, and whose dollar wages might have been far out of equilibrium at least in the early transition years (the conventional wisdom being that they were highly undervalued). Thus, if we include the transition economies in the regression sample without any further ado, our equilibrium wage estimates for given levels of the right-hand side measures are likely to be biased downward. To avoid this, one could either leave the transition economies out of the regression sample altogether, or include them in the sample but add a "transition dummy" to the regression. The latter procedure is more efficient, since it exploits information from the additional countries to estimate the slope parameters of the regression. Note that the transition dummy needs to be time varying, since one cannot assume that the extent to which the transition countries are out of equilibrium stays the same over the period covered by our panel. Similarly, it would seem advisable to include separate dummies for Central and Eastern Europe and countries of the former Soviet Union, since the former began their transition earlier than the latter.

In summary, our basic regression model is

$$w_{\$i,t} = a_0 + a_1 agr_{i,t} + a_2 gdp_{i,t} + a_3 school_{i,t} + a_4 oecd_i + \sum_{\tau=1}^{T} a_{\tau+4} cee_{i,t} + \sum_{\tau=1}^{T} a_{\tau+9} fsu_{i,t} + \mu_i + \varepsilon_{i,t}, \tag{18}$$

where $w_{\$i,t}$ stands for the log of dollar wages, $agr_{i,t}$ for the log of the percentage share of agriculture in GDP, $gdp_{i,t}$ for the log of normalized PPP-adjusted GDP, $school_{i,t}$ for a human capital proxy based on secondary school enrollment data, which are discussed below, $oecd_i$ for a dummy that equals 1 if i was an OECD member over the entire sample period and 0 otherwise, and $cee_{\tau i,t}$ for a dummy variable that takes on the value 1 in period t if $t = \tau$ and if country i is a transition economy of Central or Eastern Europe and 0 otherwise. $fsu_{i,t}$ is an analogous dummy for the CIS and Baltic countries and μ_i is a country-specific effect. Expressing $w_{\$i,t}$, $gdp_{i,t}$, and $agr_{i,t}$ in logs follows convention and contributes toward normality of the error in equation (18) in view of the skewness and nonnegativity of the wage and income distributions. The model was also extended to control for differences in taxation, the degree of government intervention, and property rights (see Section II).

We are left with a final issue, namely, what to do about the transition dummies when computing the fitted wages based on equation (18) for transition economies, which in accordance with equation (1) will be interpreted as equilibrium wage estimates. Clearly, it would be wrong to include the transition dummies when we compute fitted wages, since these dummies will pick up the potentially large initial exchange rate misalignment that might be common to transition economies—this is why they were added. Thus, comparing actual dollar wages in a transition economy with its fitted dollar wage including the transition dummy will tell us whether this economy's dollar wage is under- or overvalued relative to the average under- or overvaluation in the group of transition countries, not relative to the equilibrium dollar wage corresponding to its human and physical capital endowment. The relevant comparison is thus between the actual dollar wage in a transition economy and its fitted wage based on equation (18) after setting the transition dummies to zero. The implicit assumption is that in all aspects not captured by the three time-varying right-hand side variables (and later on, the additional variables introduced to account for institutional differences), the transition economy is structurally similar to an average non-OECD economy in our sample. Put differently, we need to assume that the transition dummies included in the regression reflect *only* the average extent of exchange rate misalignment in transition economies during the sample period, rather than structural differences between transition economies and nontransition developing countries. We return to this assumption below.

II. Estimation and Testing

Data

Our data set begins in 1990, the earliest starting date of economic transition in Eastern Europe, as defined by the first comprehensive attempt in a formerly planned economy to both liberalize prices and open the economy.[17] The endpoint was chosen to be 1995, the last year for which a large cross-sectional coverage could be achieved. Between these two dates, we attempted to include all market economies for which data were available for any number of consecutive years, and all transition economies for which a continuous sequence of annual data points was available from their first year of transition onward. This led to an unbalanced panel of 85 countries, including 15 transition economies: Bulgaria, the Czech Republic, Hungary, Poland, Romania, the Slovak Republic, the three Baltic states, Belarus, Kazakhstan, the Kyrgyz Republic, Moldova, Russia, and Ukraine. Our starting dates are 1990 for Hungary and Poland, 1991 for the remaining Eastern European countries, and 1992 for the Baltic and CIS countries. The market economies in our sample include all OECD countries, most Latin American countries, and some African and Asian economies.[18]

The economic variables used in our regressions were constructed as follows.

- *Dollar wages* are average monthly wages in manufacturing in U.S. dollars. Data on manufacturing wages in national currencies were obtained from ILO publications and from the OECD "Short Term Indicators Transition Economies" database. In addition, we used national statistical publications and, in some cases, the IMF's "Recent Economic Developments" country reports to fill in gaps and to broaden the coverage of the sample. In cases when information on hours was not available, the hourly wage data were converted into monthly wages by assuming an 8-hour working day and a 4.3-week month. Monthly wages were then expressed in U.S. dollars using annual average exchange rates from the IFS. In order to ensure cross-country comparability, we made every effort to obtain wage data for employees for each country. In some cases, however, only wage rates for workers were available, which tend to be substantially lower than employee wages. We included these countries in the sample but attempted to control for the difference

[17] See EBRD (1994), Appendix 2.1, for an overview of the major measures and their timing for each country.

[18] Appendix Table A1 reports the countries and years covered.

in the definition of the dependent variable by including an appropriately defined dummy variable on the right-hand side.[19]

- Data for *purchasing-power-parity-adjusted GDP* ("PPP GDP" for short) were obtained from the IMF's World Economic Outlook (WEO) database. The purchasing power parity estimates used by the WEO are based on the Penn World Table Mark 5 and (for transition countries) on a comparative study by the United Nations Economic Commission for Europe, and extended using "bridging equations"; see Wagner (1995). The problem with these data is that the estimates for transition countries are based on a pretransition production structure that is bound to become increasingly inaccurate as transition proceeds. The World Bank and the European Union (EU) provide estimates based on more recent price comparisons, but for the transition economies they are only available for one year, 1995, and in the case of the EU data, only for the Eastern European economies. Consequently, we generally relied on the WEO data for the purposes of estimation but used World Bank and EU data to check the robustness of our equilibrium wage estimates (see below).[20]

We experimented with normalizing PPP GDP by total population, labor force, and employment (obtained from the WEO database and from the World Bank's Social Indicators of Development database) to construct three alternative proxies for overall productivity. The argument presented in Section I was based on normalization by the labor force; however, this assumed that PPP GDP could be measured at its equilibrium (i.e., full capacity) level. In the presence of unemployment or overemployment, normalization by actual employment might be preferable, as this is likely to imply the least cyclical productivity measure. However, this measure has the problem that it will, all things being equal, overstate productivity in economies with a large subsistence sector. Normalizing by the labor force or, alternatively, population (as in Halpern and Wyplosz and most of the literature on explaining deviations from PPP) will avoid that particular problem. In the end, we decided to run our regressions using all three measures to check the robustness of the results to the choice of normalization. Normalized PPP GDP is denoted by "*gdp*".

[19]That is, a dummy that took the value of one whenever the average wage referred to workers and zero otherwise. In addition, we also reestimated all equations using only observations with employee wages on the left-hand side. The results, reported in Appendix Table A2, are very similar to the ones obtained using the larger sample.

[20]Sources: EU Agenda 2000 Comparison Table, available on-line at http://europa.eu.int/comm/dg1a/agenda2000/en/impact/annex.htm; and World Bank, 1997, Table 1 (pp. 214–15).

- The *share of agriculture in GDP* ("*agr*") was taken from the World Bank's 1996 "Social Indicators of Development" (SID) database. Since the SID database did not contain 1995 shares of agriculture, we decided to use the lagged, rather than contemporaneous, agriculture share in our regressions. This procedure can be justified on the grounds that the importance of agricultural production over the short time period considered is likely to capture cross-country differences in technological development rather than time trends; hence, the lagged share of agriculture is a satisfactory proxy to use.[21] A few missing values for 1994 and earlier years were filled in using national statistics or were interpolated.

- A *human capital variable* ("*school*") was constructed from the SID database as a measure of the average level of secondary school education of the labor force in each country and year. More precisely, $school_{i,t}$ is the average of secondary school enrollment ratios in country i between 1950 (the first year reported in the SID) and $t-1$, weighting the enrollments ratios in each year with the relative size of the cohort of fifteen-year-olds at the time. Since the first year reported in SID for school enrollment is 1965, we assumed 1965 schooling levels for the earliest cohorts. We also assumed lower participation rates for these cohorts in our sample period.[22] While this method is obviously crude, we consider the resulting indicator a better proxy for the human capital stock than either contemporaneous or lagged data on schooling for a particular cohort. In addition to the human capital variable thus constructed, we also experimented with alternative measures (average years of primary, secondary, and higher education in the population in the 15+ age group; share of people with different levels of education in the 15+ age group) from a database compiled by Barro and Lee (1996).

- To control for possible cross-country differences in the legal framework of economic activity, which are likely to influence productivity, we augmented the list of right-hand side variables by *indicators of government intervention, tax structure, and property rights* and an overall index of

[21]This assumption need of course not hold universally. For example, certain countries—such as Denmark—may have relatively high shares of agriculture in spite of an advanced level of technological development. If a negative coefficient on *agr* is estimated on the whole sample, this would lead to a fitted wage for Denmark that is probably an underestimation of the Danish equilibrium wage. There is nothing that can be done about this problem except to adjust the proxy ad hoc (at the stage of computing fitted wages) if for a specific country it seems clearly out of line with the underlying variable that it is supposed to proxy for.

[22] In particular, we assumed that the participation rate for the 1945–50 and 1950–55 cohorts were 60 and 80 percent, respectively.

"economic freedom" constructed by the Heritage Foundation. The indicators are based on 1994–95 information.

Econometrics

As discussed in Section I, the main rationale for estimating equation (18) in a panel regression is to address the presence of country-specific effects in the determination of equilibrium wages.[23] The simplest form of accounting for country-specific effects—that is, country-varying determinants of equilibrium wages that are not captured by the right-hand side variables in equation (18)—is to rewrite equation (18) as

$$w_{\$i,t} = a_0 + \sum_{j=1}^{3} a_j x_{ji,t} + a_4 oecd_i + \sum_{\tau=1}^{T} a_{\tau+4} cee_{i,t} + \sum_{\tau+1}^{T} a_{\tau+9} fsu_{i,t} + \mu_i + \varepsilon_{i,t}, \quad (19)$$

where x_j, $j = 1...3$ denotes the economic determinants of equilibrium wages in equation (18) and μ_i stands for unmeasured country-specific effects. If μ_i is uncorrelated with the remaining right-hand side variables, its presence will merely generate a serial correlation problem that can be corrected by estimating equation (19) using a random effects (RE) estimator. If, on the other hand, μ_i is correlated with any of the x_j, we face a much more serious problem. On the one hand, the endogeneity of the error term induced by μ_i will preclude consistent estimation of equation (19) using the pooled OLS or random effects estimators. On the other hand, first difference (FD) or fixed effects (FE) estimators, which are based on transforming equation (19) in a way that eliminates μ_i, do not allow the estimation of the *cee* and *fsu* dummies and of the constant a_0 and therefore preclude the computation of fitted wages for the transition economies (see Appendix for details).

Whether or not fixed effects are present, equation (19) may suffer from the standard endogeneity problem of a contemporaneous correlation between the residual $\varepsilon_{i,t}$ and one or more of the economic right-hand side variables. Since the agriculture share (*agr*) and the human capital indicator (*school*) are based on lagged GDP, agricultural production, schooling, and population growth, the most plausible candidate for this type of correlation is normalized PPP-GDP (*gdp*).[24]

[23]For an overview of the panel estimators discussed below, see Greene (1993), Chapter 16.4, and Keane and Runkle (1992)

[24]For example, consider an economy which receives a positive technology shock. PPP-GDP rises and the equilibrium exchange rate appreciates. If the market exchange rate follows suit and overshoots, this could result in an increase in dollar wages over and above the increase which corresponds to the increase in PPP-GDP, leading to a positive correlation between PPP-GDP and the error term..

To address these problems, we proceeded in two steps. First, we performed standard Hausman tests[25] based on estimating equation (19) in first differences, that is, comparing the plain OLS FD estimates of equation (19) with FD estimates using lagged right-hand side variables as instruments. Lagged endogenous variables are clearly less than ideal as instruments, but unfortunately we had no better alternative. The FD specification is appropriate because it is immune to any additional endogeneity problem through the presence of fixed effects. The null hypothesis of no endogeneity cannot be rejected at conventional significance levels for all three normalizations. p-values for testing H_0: "in the first-differenced version of equation (19), the residuals are orthogonal to the right-hand side variables" versus H_A: "residuals are correlated with the right-hand side variables" are reported in the last row of the FD columns in Table 2. We thus conclude that an endogeneity bias through the contemporaneous correlation of ε and any x_j is likely to be negligible.

Based on these findings, we assume that weak exogeneity is satisfied and as a result the FD estimator is consistent. The next step is to test for the presence of fixed effects. There are several ways of accomplishing this. Since the FD estimator is consistent, we could perform a Hausman-type test for the presence of fixed effects by comparing the FD estimates of the parameter vector **a** in equation (19) (consistent both under the null of no fixed effects and under the alternative) with the pooled OLS estimates of **a** (consistent only under the null of no fixed effects). Alternatively, we could perform the more conventional test for the presence of fixed effects by comparing the FE and the RE estimates of **a**, but in this case we need to test for strict exogeneity first ($x_{ji,t}$ uncorrelated with $\varepsilon_{i,t}$ at all leads and lags) since without this property both FE and RE will be inconsistent whether or not fixed effects are present. This can be achieved via a Hausman-type test of FD versus FE (see Appendix). Yet alternatively, we could test FD against RE, in which case we would be jointly testing strict exogeneity and the absence of fixed effects. As it turns out, all three procedures imply that in our case neither the absence of fixed effects nor strict exogeneity can be rejected, and as a result it is legitimate to use the RE estimator.

In the following, we present the results of the procedure of sequentially testing strict exogeneity and the absence of fixed effects. The columns headed by FD and FE in Table 2 report the point estimates and standard errors for $\hat{a}^{FE}$ and $\hat{a}^{FD}$, the parameter vector estimates obtained by the FE and FD estimators respectively. Table 2 also reports the rejection levels associated with the null hypothesis of strict exogeneity in the last row of columns FE; for details on how the test statistic underlying these p-values was computed, see the Appendix. As is apparent from the p-values reported,

[25] See Hausman (1978).

the null of strict exogeneity cannot be rejected at any conventional significance level.

We proceed to test H_0: "individual-specific effects are orthogonal to the residual" (i.e., absence of fixed effects but presence of random effects) versus H_A: "individual-specific effects are correlated with the residual" (presence of fixed effects). Assuming strict exogeneity, FE is consistent under both, while RE is only consistent under the null. In addition, under H_0 the RE estimator is (asymptotically) efficient—not just more efficient than FE—which implies that the hypothesis can be tested via a standard Hausman-test.[26] Point estimates and standard errors for the (untransformed) $\hat{a}^{RE}$ are presented in Table 2, in the columns headed by RE. The p-values for the test procedures (reported in the last row) indicate that for all three specifications, the null of random effects cannot be rejected at conventional significance levels.

Table 2 also shows the *estimated coefficients* for each model specification. The RE estimates of all coefficients have the expected sign. In particular, normalized PPP GDP (*gdp*) is positive and highly significant across specifications. This remains true for both the FE and FD estimates. Based on the RE estimates, a higher share of agriculture (*agr*), which in a cross-country setting is associated with a lower degree of development and thus lower productivity, results in lower dollar wages in all three specifications. The corresponding coefficient is significant at the 5 percent level in the specifications with PPP GDP normalized by employment and labor force, but significant only at the 15 percent level in the specification with PPP GDP per capita. The RE point estimate of the human capital variable (*school*) is positive in all specifications but it is significant only in the equation with PPP GDP normalized by employment. For both the agriculture share and human capital, the variation in the significance level of the estimated coefficient is due to differences in the magnitude of the coefficient. The OECD dummy (*oecd*) is positive and highly significant in all specifications, its magnitude suggesting a sizable "dollar wage bonus" for OECD countries. The FE and FD estimates for the coefficients on *agr* and *school* are typically insignificant and in some cases have the wrong sign. Since most of the variation in the underlying variables is due to cross-country differences and not to variation over time, this result is not surprising.

[26]That is, one in which the variance-covariance matrix of the differences between the two objects of comparison reduces to the difference between the variances of these objects. The only nonstandard element of the procedure is that, as in the previous step, the FE and RE estimators have to be transformed to obtain comparable parameter vectors, using the linear conditions governing the relationship between $\mathbf{a}^{FE}$ and $\mathbf{a}^{FD}$ (see Appendix). Thus, the test statistic is distributed $\chi^2(k)$, where k is the rank of the transformation matrices $\mathbf{B}$ and $\mathbf{C}$. In our case, $k = 13$.

Table 2. *Estimated Coefficients and Specification Test Results*
(Dependent variable: monthly average wages in manufacturing, in U.S. dollars; standard errors are in parentheses)

Variable	Specification (1)			Specification (2)			Specification (3)		
	FD	FE	RE	FD	FE	RE	FD	FE	RE
gdp	0.978	1.092	0.803	0.364	0.623	0.499	0.948	1.069	0.749
	(0.209)	(0.135)	(0.101)	(0.199)	(0.148)	(0.101)	(0.216)	(0.142)	(0.104)
agr	0.109	–0.047	–0.092	–0.009	–0.188	–0.238	0.102	–0.051	–0.136
	(0.114)	(0.086)	(0.063)	(0.116)	(0.093)	(0.061)	(0.115)	(0.088)	(0.063)
school	–0.304	–0.451	0.121	0.677	0.264	0.359	–0.110	–0.197	0.127
	(0.759)	(0.406)	(0.134)	(0.771)	(0.430)	(0.133)	(0.751)	(0.400)	(0.139)
oecd	—	—	0.641	—	—	0.905	—	—	0.761
	—	—	(0.158)	—	—	(0.157)	—	—	(0.158)
cee90	—	–0.408	–1.337	—	–0.267	–1.098	—	–0.410	–1.177
	—	(0.104)	(0.253)	—	(0.108)	(0.269)	—	(0.106)	(0.259)
cee91	0.308	–0.238	–1.146	0.223	–0.186	–1.003	0.306	–0.241	–1.000
	(0.131)	(0.057)	(0.230)	(0.133)	(0.060)	(0.245)	(0.132)	(0.057)	(0.237)
cee92	0.241	–0.033	–0.961	0.176	-0.058	–0.885	0.239	–0.037	–0.828
	(0.079)	(0.056)	(0.228)	(0.080)	(0.060)	(0.244)	(0.079)	(0.057)	(0.237)
cee93	0.149	0.109	–0.816	0.139	0.059	–0.762	0.151	0.106	–0.681
	(0.073)	(0.056)	(0.228)	(0.076)	(0.059)	(0.242)	(0.073)	(0.057)	(0.236)
cee94	–0.015	0.089	–0.820	0.017	0.060	–0.752	–0.010	0.091	–0.679
	(0.073)	(0.055)	(0.227)	(0.076)	(0.059)	(0.240)	(0.074)	(0.056)	(0.235)

cee95	0.129	0.209	–0.680	0.171	0.214	–0.590	0.134	0.217	–0.532
	(0.074)	(0.056)	(0.227)	(0.076)	(0.060)	(0.239)	(0.074)	(0.056)	(0.234)
fsu92	—	–0.755	–2.252	—	–0.630	–2.130	—	-0.753	–2.086
	—	(0.051)	(0.228)	—	(0.051)	(0.247)	—	(0.052)	(0.238)
fsu93	0.467	–0.279	–1.811	0.376	–0.241	–1.751	0.464	–0.280	–1.652
	(0.064)	(0.046)	(0.230)	(0.063)	(0.048)	(0.249)	(0.065)	(0.046)	(0.241)
fsu94	0.579	0.300	–1.272	0.475	0.239	–1.286	0.577	0.299	–1.122
	(0.067)	(0.047)	(0.232)	(0.065)	(0.049)	(0.251)	(0.068)	(0.048)	(0.245)
fsu95	0.440	0.734	–0.843	0.411	0.633	–0.890	0.441	0.735	–0.694
	(0.059)	(0.048)	(0.231)	(0.061)	(0.049)	(0.250)	(0.060)	(0.049)	(0.245)
workers_other	—	—	–0.360	—	—	–0.307	—	—	–0.368
	—	—	(0.165)	—	—	(0.170)	—	—	(0.168)
workers_oecd	—	—	–0.034	—	—	0.020	—	—	–0.013
	—	—	(0.213)	—	—	(0.219)	—	—	(0.217)
constant	—	—	–1.289	—	—	0.096	—	—	–1.421
	—	—	(0.858)	—	—	(0.989)	—	—	(0.934)
N	307	392	392	307	392	392	307	392	392
R^2	0.37	0.62	0.89	0.33	0.57	0.89	0.37	0.61	0.89
Test p-value	0.89	0.59	0.89	0.39	0.15	0.99	0.79	0.58	0.85

Notes: Specifications (1), (2), and (3) refer to normalization of GDP by population, employment, and the labor force, respectively. Coefficients on cee and fsu dummies are not directly comparable across FD, FE, and RE estimates, see Appendix II. For FE and RE models, "test p-value" refers to a specification test against the closest model to the left. For FD models, "test p-value" refers to a specification test against FD models using lagged variables as instruments.

The RE estimates show that the coefficients on the transition dummies (*cee* and *fsu*) are negative and significant throughout the sample period, suggesting lower than warranted dollar wages for these countries.[27] The absolute value of the dummies shows a steady decline.[28] This time path of the estimated dummies is consistent with the view that the transition economies were undervalued initially and that the gap has been closing over time.

Comparing the three specifications reveals differences in the estimated coefficients on the PPP GDP variables, agriculture share, human capital, and the OECD dummy. In particular, (i) the coefficients on *gdp* seem to be fairly close to one and to each other in the specifications which use labor force and population for normalization purposes, but only about half this size in specification (2), which uses employment; and (ii) in the presence of the employment-based productivity variable, the additional productivity or development proxies *agr* and *school* are significant, whereas in the other two specifications they tend to be insignificant. One possible interpretation for the latter is that, controlling for population and labor force, employment may show swings across countries depending on the size of the subsistence sector; these swings are unrelated to productivity in the tradables sector and thus to wages but they are correlated with *agr* and *school* as indicators of development.

Note, finally, that the three specifications produce virtually identical values for the regression R^2, that is, the right-hand side variables account for the same percentage of variation in dollar wages in the three equations. This, and the fact that there are noticeable differences in the coefficients across specifications, suggest that we examine the implications of all three specifications on estimated equilibrium dollar wages (See next section).

To check the *robustness* of the estimated coefficients we reestimated the equations with the following modifications. First, instead of using our human capital variable we used various combinations of the educational indicators by Barro and Lee (1996). Second, we restricted the sample to observations with average employee wages. Third, we experimented with using the share of mining in GDP as an additional regressor. Fourth, we attempted to account for cross-country differences in institutional factors influencing productivity by augmenting the list of variables by (1) an indicator of taxation; (2) an indicator of the extent of government intervention;

[27]Note that the FD, FE, and RE estimates of these dummies are not comparable. In particular, the coefficients on the transition dummies in the FD and FE specifications carry *no* information about the average effect of transition on the *level* of dollar wages, which is estimated by coefficients on the transition dummy in the RE specifications (see Appendix).

[28]The difference between the earliest (1990 for Central and Eastern Europe and 1992 for the Baltic and CIS countries) dummy and the latest one (1995) is statistically significant for both country groups in all three specifications.

(3) an indicator of the firmness of property rights; and (4) an overall index of economic freedom. Fifth, we experimented with alternative PPP GDP estimates for transition economies using EU and World Bank data. In this case, we used EU and World Bank data for the transition economies,[29] but kept IMF WEO data for the remaining countries, as neither of the other two data sources provide panel data for all countries and years covered in the sample.

As it turns out, our results are not sensitive to any of these modifications.

- The Barro-Lee schooling variables are strongly correlated with our human capital measure, and using them produces numerical results practically identical to our original equations.

- Restricting the sample to countries for which we were able to obtain average employee wages also produced results similar to the ones for the larger sample.

- The mining share had a positive but small (relative to the coefficient of the share of agriculture) and insignificant coefficient. Its inclusion had a negligible effect on the remaining coefficients.

- Of the economic freedom indicators, the overall index proved entirely insignificant, with a coefficient close to zero and a large standard error. Property rights and taxation were insignificant as well. Only the indicator of government intervention was statistically significant and *positive*. This indicates that stronger government intervention (as measured by an index based on the share of government consumption in GDP and on the significance of state ownership in the economy) is, all things being equal, associated with higher equilibrium wages. However, the inclusion of this variable did not influence the significance of our "core" variables and had a negligible impact on the magnitude of all except one coefficient, that on the OECD dummy, which declined substantially in all specifications. This suggests that the source of the positive impact of government intervention on warranted wages is the fact that the economic role of the government tends to be higher in OECD countries than in the rest of the sample. However, adding the government intervention variable to the list of right-hand side variables did not yield a considerable improvement in fit, nor did it lead to diminished significance and/or magnitude of the transition dummies. Moreover, the esti-

[29]For each transition country, 1991–95 real growth rates were applied to the available 1995 EU and World Bank estimates to generate a series covering the entire sample period.

mates of equilibrium wages were insignificantly different from those ignoring the government intervention index.

- Using different PPP-adjusted GDP measures for the transition economy group had negligible effects on the coefficient estimates.

On this basis, we decided to base our estimates of equilibrium dollar wages on the three parsimonious RE specifications reproduced in Table 2, and indicate the extent to which the estimated equilibrium wages are sensitive to the use of alternative PPP-adjusted GDP data (even if the regression coefficients are not, the fitted wages for individual transition economies may well be). For the purposes of deciding which currencies were undervalued in 1996 and which were not, we show estimates based on PPP-adjusted GDP data from all three sources.

Equilibrium Wage Estimates

Results

The RE estimates of the magnitude of the transition dummies suggest that dollar wages in countries of Central and Eastern Europe as a group were 75–65 percent under equilibrium (that is, their actual wages were at 25 to 35 percent of the warranted wage) in 1990 and 1991. This compares with 55 to 45 percent undervaluation in 1995. For the Baltics and the CIS countries, we see about 90 percent undervaluation in 1992 and about 60 percent in 1995. In conclusion, for our sample's transition economies *as a group*, though the gap between actual and equilibrium wages declined, it does not seem to have been eliminated by 1995.

We now examine whether this result holds for individual countries. Tables 3 and 4 report actual and estimated equilibrium wages for Central and Eastern European and Baltic and CIS countries in our sample. Our equilibrium wage estimates are fitted wages based on the RE model, setting both the estimated country-specific effect and the transition dummies to zero,[30] and using IMF WEO data on PPP-adjusted GDP per capita, for each of the three specifications. The sensitivity of these results to the use of PPP-adjusted GDP data from different sources is discussed at the end of the section. Note that

[30]For the argument on why transition dummies should be set to zero when computing equilibrium wages, see the end of Section I. The need to also set the country-specific effects to zero follows from the interpretation of random effects as capturing the initial out-of-equilibrium position of dollar wages (in this case, relative to other transition economies) rather than unobserved structural determinants of equilibrium wages; see Section I.

Table 3. *Actual and Estimated Dollar Wages in Manufacturing, US$/month: Selected Central and Eastern European Countries*
(Standard errors are in parentheses)

			Fitted					
		Actual	Specification (1)		Specification (2)		Specification (3)	
Country	Year	wage	Wage	Ratio	Wage	Ratio	Wage	Ratio
Bulgaria	1991	52	332	16	292	18	278	19
			(47)		(45)		(42)	
	1992	96	322	30	308	31	274	35
			(45)		(46)		(41)	
	1993	123	327	38	319	39	281	44
			(45)		(47)		(41)	
	1994	97	343	28	337	29	295	33
			(46)		(48)		(42)	
	1995	121	358	34	345	35	306	40
			(48)		(49)		(43)	
Czech Republic	1991	132	511	26	417	32	420	31
			(67)		(62)		(59)	
	1992	162	507	32	442	37	422	38
			(66)		(66)		(59)	
	1993	196	513	38	450	44	424	46
			(66)		(66)		(60)	
	1994	231	529	44	447	52	433	53
			(68)		(65)		(60)	
	1995	303	566	54	480	63	462	66
			(72)		(68)		(63)	
Hungary	1990	177	395	45	284	62	356	50
			(48)		(39)		(44)	
	1991	187	372	50	282	66	338	55
			(44)		(39)		(41)	
	1992	226	388	58	331	68	359	63
			(43)		(44)		(42)	
	1993	287	403	71	369	78	375	77
			(44)		(47)		(44)	
	1994	304	425	72	392	78	394	77
			(46)		(49)		(45)	
	1995	309	438	71	396	78	403	77
			(48)		(48)		(46)	
Poland	1990	105	344	30	298	35	298	35
			(45)		(44)		(41)	
	1991	153	349	44	332	46	308	50
			(44)		(49)		(42)	
	1992	200	371	54	362	55	329	61
			(46)		(52)		(45)	
	1993	207	389	53	389	53	344	60
			(48)		(53)		(46)	
	1994	229	415	55	404	57	364	63
			(50)		(54)		(47)	
	1995	288	449	64	428	67	391	74
			(53)		(56)		(50)	

Table 3. *(concluded)*

Country	Year	Actual wage	Fitted: Specification (1) Wage	Ratio	Specification (2) Wage	Ratio	Specification (3) Wage	Ratio
Romania	1991	90	223	40	187	48	200	45
			(30)		(31)		(29)	
	1992	61	215	28	190	32	195	31
			(29)		(32)		(28)	
	1993	74	222	33	200	37	200	37
			(30)		(32)		(29)	
	1994	80	232	34	206	39	207	39
			(31)		(33)		(30)	
	1995	102	252	41	218	47	223	46
			(34)		(34)		(31)	
Slovak Republic	1991	129	441	29	431	30	373	34
			(60)		(64)		(55)	
	1992	161	435	37	450	36	372	43
			(59)		(67)		(56)	
	1993	175	427	41	439	40	362	48
			(59)		(66)		(55)	
	1994	196	448	44	451	44	376	52
			(61)		(66)		(56)	
	1995	242	477	51	456	53	395	61
			(64)		(66)		(57)	

Note: "ratio" denotes actual wage as a percentage of fitted wage.

the "ratio" columns of the tables give actual wages in percent of estimated equilibrium wages.

The main points emerging from Tables 3 and 4 are as follows.

First, with minor exceptions, *actual* dollar wages exhibit a similar *profile* across countries, namely monotonically rising throughout the transition process, with particularly sharp increases in the first two years. However, dollar wage *levels* vary greatly across countries, and these differences are quite persistent through time. While it is possible to identify subgroups for which there is considerable convergence in actual dollar wages over the period (e.g., the Czech Republic, Estonia, Hungary, Poland, the Slovak Republic, and possibly Latvia), the dispersion within the transition country group as a whole does not seem to decline substantially during the period examined.

Second, there are interesting differences in the profiles of estimated *equilibrium* dollar wages across countries. With the caveat that the standard errors around these estimates are fairly high,[31] two broad patterns can be

[31]In particular, estimates of the equilibrium wages at time t are typically within the one-standard-error bands around the equilibrium wage estimates at time $t-1$ and $t+1$.

Table 4. *Actual and Estimated Dollar Wages in Manufacturing, US$/month: Baltic and Selected CIS Countries*
(Standard errors are in parentheses)

			Fitted					
		Actual	Specification (1)		Specification (2)		Specification (3)	
Country	Year	wage	Wage	Ratio	Wage	Ratio	Wage	Ratio
Belarus	1992	26	406	6	303	9	328	8
			(67)		(53)		(55)	
	1993	26	374	7	285	9	303	8
			(63)		(51)		(52)	
	1994	32	355	9	298	11	294	11
			(59)		(52)		(48)	
	1995	75	333	22	292	26	276	27
			(62)		(51)		(46)	
Estonia	1992	45	387	12	317	14	309	14
			(63)		(55)		(51)	
	1993	78	374	21	332	24	303	26
			(62)		(57)		(50)	
	1994	137	414	33	378	36	338	41
			(65)		(61)		(53)	
	1995	211	444	48	399	53	360	59
			(68)		(63)		(56)	
Latvia	1992	60	308	19	246	24	246	24
			(54)		(46)		(42)	
	1993	70	282	25	249	28	230	30
			(51)		(46)		(40)	
	1994	143	299	48	270	53	244	59
			(53)		(48)		(41)	
	1995	195	308	63	279	70	251	78
			(58)		(49)		(42)	
Kazakhstan	1992	26	232	11	233	11	203	13
			(52)		(48)		(41)	
	1993	70	213	33	221	32	184	38
			(49)		(46)		(38)	
	1994	69	170	41	196	35	148	47
			(44)		(43)		(33)	
	1995	112	161	69	203	55	140	80
			(42)		(44)		(32)	
Kyrgyz Republic	1992	14	172	8	169	8	167	9
			(63)		(37)		(33)	
	1993	21	153	14	167	13	149	14
			(62)		(37)		(31)	
	1994	35	129	28	152	23	126	28
			(65)		(35)		(27)	
	1995	53	132	40	168	31	129	41
			(68)		(37)		(28)	

Table 4. *(concluded)*

			Fitted					
		Actual	Specification (1)		Specification (2)		Specification (3)	
Country	Year	wage	Wage	Ratio	Wage	Ratio	Wage	Ratio
Moldova	1992	20	164	12	167	12	141	14
			(54)		(37)		(30)	
	1993	23	165	14	166	14	141	17
			(51)		(36)		(30)	
	1994	35	123	28	146	24	107	33
			(53)		(34)		(25)	
	1995	46	123	38	147	31	106	43
			(58)		(34)		(25)	
Russia	1992	32	307	10	273	12	257	12
			(42)		(50)		(44)	
	1993	63	289	22	276	23	245	26
			(39)		(52)		(43)	
	1994	96	265	36	278	34	230	42
			(40)		(55)		(43)	
	1995	107	269	40	300	36	235	45
			(44)		(59)		(44)	
Ukraine	1992	28	257	11	221	13	218	13
			(52)		(44)		(40)	
	1993	14	228	6	213	7	197	7
			(49)		(43)		(37)	
	1994	26	181	15	173	15	155	17
			(44)		(39)		(32)	
	1995	47	166	28	164	29	143	33
			(42)		(38)		(31)	
Lithuania	1992	20	218	9	196	10	185	11
			(42)		(39)		(33)	
	1993	48	174	27	169	28	149	32
			(39)		(36)		(29)	
	1994	84	183	46	186	45	158	53
			(40)		(38)		(30)	
	1995	124	194	64	196	63	167	74
			(44)		(39)		(31)	

Note: "ratio" denotes actual wage as a percentage of fitted wage.

distinguished. In one group of countries—the Czech Republic, Hungary, Poland, the Baltics, and to a lesser extent, Bulgaria and Romania—estimated equilibrium wages increase more or less monotonously throughout the transition, typically after falling once at the beginning of our sample period. Except for the initial decline, which could be due to an overestimation of productivity in the early stages of transition, when the obsolescence of part of the capital stock has not yet been reflected in a decline in output, these countries broadly follow the pattern of continuous real equilibrium appreciation predicted by the benchmark neoclassical model. In the CIS countries, on the other hand, we observe declining or flat wages over the entire observation period. This is particularly true for Belarus, Kazakhstan, Moldova, and Ukraine, where equilibrium wages first decline and then stay flat, while in Russia and the Kyrgyz Republic equilibrium wages seem to turn the corner only in 1995. One interpretation might be that in these countries the process of new capital accumulation that drives recovery and real appreciation in transition countries according to the simple neoclassical model had not set in by 1995, for reasons that are not captured in that model, but that could relate to differences in the speed and consistency of reforms, the legal and political environment, and possibly location.[32]

Third, looking at actual and equilibrium dollar wages on a country-by-country basis confirms our earlier conclusion that despite their rapid increase, actual dollar wages did not catch up with equilibrium wages by 1995. The results in Tables 3 and 4 suggest that at the outset of the transition process, dollar wages in manufacturing in the Czech Republic, Poland, Romania, and the Slovak Republic were about 30–40 percent of their equilibrium level, while dollar wages in Bulgaria, the Baltics, and the CIS countries seem to have started out with an even larger discrepancy from their equilibrium level (10–20 percent). At the other extreme, our estimates put dollar wages in Hungary in 1990 at 50–60 percent of the warranted wage. This clearly indicates that—based upon our numerical estimates—all transition countries in our sample entered the transition with "too low" dollar wages (or, in our interpretation, with an undervalued exchange rate). By 1995, actual dollar wages in Central and Eastern European countries stood between 35 and 75 percent of equilibrium wages, with Bulgaria at the low and Hungary and Poland at the high end of the range. The dispersion in the extent of undervaluation in 1995 is at least as pronounced for the Baltics and the CIS countries, where actual dollar wages range from 25–30 percent of estimated equilibrium wages (Belarus, Ukraine) to around 70 percent (Latvia, Lithuania, Kazakhstan). These results imply that for specifications (1) and (2), the actual wage is outside the one-standard-error band around the equilibrium wage estimate in all cases and the hypothesis that the equilibrium wage is smaller than the actual wage is formally rejected throughout at either the 1 or 5 percent level.[33] According to specification (3), finally, the null is rejected at the 1 or 5 percent levels for

[32]Note, however, that these differences were not picked up when we extended the basic model by including political and institutional indicators for 1994–95; see subsection on robustness.

[33]The latter refers to Hungary (1993–95) and Latvia (1995); the p-values are 0.023, 0.019, 0.20, and 0.018, respectively.

all countries except Kazakhstan, Latvia, and Lithuania in 1995 (p-values 0.153, 0.059, and 0.051, respectively).

Using EU or World Bank data PPP-adjusted GDP data for transition economies does not affect any of these qualitative conclusions. In particular, the same observations apply regarding cross-country differences in the profiles of fitted wages, and the overall conclusion that actual dollar wages remained below estimated equilibrium wages in 1995 and before. However, there are noteworthy differences in the equilibrium wage estimates for individual countries. In particular, using EU or World Bank data leads to equilibrium wage estimates for Estonia, Latvia, and Lithuania that are much closer together, implying that the hypothesis that the equilibrium wage is smaller than the actual wage could now be formally rejected in the case of Lithuania but perhaps not Estonia (depending on the specification). In addition, using World Bank data makes Kazakhstan look better, increasing its equilibrium wage estimates by about 60 dollars.

In summary, even in countries where dollar wages rose sharply in the course of transition, they seem to remain below equilibrium, in most cases at less than 70 percent of the equilibrium estimate. Thus, we are confident that most transition countries did not have overvalued currencies by 1995. Unfortunately, owing to the unreliability of PPP-adjusted GDP data, there is considerable uncertainty in determining the set of countries that might constitute exceptions. We return to this thorny issue in the subsection below on competitiveness in 1996.

Discussion

Before proceeding, we now ask whether and to what extent the findings of the previous subsection could exaggerate the actual state of competitiveness in transition countries. In principle, we could have erred in two ways: equilibrium wages for transition economies could be overestimated, and reported wages in these countries could understate actual wages. We discuss each in turn.

Equilibrium wages for transition economies could be overestimated as a result of suppressing the transition dummies when computing fitted wages. As explained in Section I, transition dummies were included to capture the possibility that dollar wages in transition economies might be far off their equilibrium path, particularly at the early stages of transition. The implicit assumption was that in any aspects relevant to equilibrium dollar wages not captured by the other right-hand side variables, transition economies were similar to a representative developing country during the period in question. If this assumption is violated, then any unrepresented structural idiosyncracies would be absorbed into the transition dummies. If the net effect of these factors is to depress equilibrium dollar wages in transition economies relative to nontransition developing countries, suppressing the transition dummy in the calculation of equilibrium dollar wages for transition economies would lead to overestimation.

In our view, this problem could arise for the early years of transition, in particular because at this time PPP GDP per capita or worker could be a very poor proxy of productivity in the tradables sector. At the beginning of transition, measured real GDP has typically not yet declined to reflect the obsolescence of large sectors of the economy, thus, PPP-adjusted GDP will overstate profitable productive capacity. However, once output begins bottoming out after several years of transition and real output declines on the order of 30–50 percent, this effect is unlikely to be present; if anything, one would suspect that PPP-GDP per capita now understates actual productivity since much emerging private sector activity is unrecorded. Thus, from this angle, one would not expect 1995 equilibrium dollar wages to be overestimated. This still leaves the possibility that important determinants of equilibrium dollar wages—such as property rights, the legal framework, or political uncertainty—might be poorly captured by the productivity measures used. However, including direct measures of these variables in the regression did not seem to affect substantially the estimates of equilibrium wages, as discussed previously. Moreover, to introduce an upward bias in the equilibrium dollar wage estimates for transition economies, transition economies would need to do significantly worse in terms of unmeasured institutional conditions than the average developing country in our sample. This may be plausible for the beginning of transition, but not—in most cases—for the later years. On this basis, the equilibrium wage estimates for 1995/96 in countries such as Hungary, the Czech Republic, the Slovak Republic, Poland, Estonia, Latvia, Lithuania, and the Kyrgyz Republic is more likely to be biased downward than upward. In other cases—such as Belarus, which, in a sense, is still at the beginning of transition—the possibility that equilibrium wages could indeed be overestimated seems more plausible.

We are left with the question whether the undervaluation of currencies in transition economies might be exaggerated due to the systematic underreporting of actual wages in transition economies, rather than the overestimation of equilibrium wages. There is good reason to suspect that wage underreporting could in fact be present. First, reported wages for transition economies could understate actual remuneration relative to nontransition economies because in-kind payments and other nonwage benefits may be a large share of earnings in transition economies, particularly at the beginning of the transition process.[34] This may contribute toward explaining

[34]For Russia, Commander and Jackman (1993) estimate that noncash benefits comprised about 35 percent of total labor income in 1992.

some of the unbelievably low actual dollar wages contained in Table 2 for the early transition years. Second, while this bias is likely to decline as transition progresses, it could be replaced by a second bias, namely, inadequate reporting of private sector wages. To the extent that newly emerging private firms are not fully captured, the official wage data reported for manufacturing or "industry" will mostly reflect public and formerly public (privatized) enterprises, whose behavior may be more similar to state-owned enterprises than to new private enterprises.[35] If these enterprises pay lower wages than new private sector enterprises, this would imply lower reported wage levels not only relative to the true sector average but, more important, relative to the market wage level relevant for new entrants. Moreover, even if total remuneration in the public or formerly public sector and the new private sector is similar, public or formerly public firms might still pay a larger share in the form of noncash benefits, and thus reported wages would be higher if based on the new private sector.[36]

The evidence on public-private sector wage differentials, typically based on survey data, is mixed. For Poland (1993), Rutkowski (1994) finds a substantial earning differential for employees with post-secondary and university educations, who earn 22–26 percent more in the private sector, but not for lower education levels, where private sector jobs tend to pay somewhat less than public sector jobs. However, he also notes that "statisticians at CSO believe that earnings in the private sector are underreported to a larger degree than in the public sector" (p. 35). In a more recent paper, (Rutkowski, 1996), he finds 27–32 percent higher earnings at the post-secondary level, and 3–11 percent higher private earnings at the pre-secondary level of educational attainment. For the Czech Republic, Flanagan (1995) finds that after controlling for schooling and experience, "workers in new private firms earn about 18 percent more than workers in current or former state enterprises" in 1993. Using similar data but without controlling for schooling or experience, Vecerník (1995) reports wage differentials of 23 percent vis-à-vis the state sector and about 9 percent vis-à-vis the privatized sector for 1993 and about the same (24 and 10 percent, respectively) for 1994. This is roughly consistent with data reported in Ham, Svejnar, and Terrell (1995), who find that in the Czech Republic "private firms generally tend to pay slightly higher (0–10 percent) average wages than state enterprises" without distinguishing between privatized and new emerging private firms. On the other hand, for 1993, Blanchard, Commander, and Coricelli (1995) find that private sector wages in Poland and Hungary are

[35]See Aghion and Carlin (1996).

[36]Estrin, Schaffer, and Singh (1995) compare state-owned, privatized and new private firms in Poland and find that the two former categories offer substantially higher levels of social benefits.

somewhat lower than public sector wages (by 7 and 14 percent, respectively), while they are somewhat higher in Bulgaria (16–50 percent) and much higher in Russia (82 percent). On Bulgaria, Beleva, Jackman, and Nenova-Amar (1995) report that "private firms tend to pay higher wages, but this is to some extent offset by lower bonuses and more limited provision of nonwage benefits" (p. 219).

From Tables 3 and 4, it is clear that correcting for wage underreporting on the order of magnitude of the private-public wage differentials reported for each country would somewhat weaken, but not reverse our basic result. In other words, even if one takes the view that wages are underreported by the full extent of the private-public sector wage differential—which amounts to assuming that private sector wages do not enter the aggregate statistics at all—actual wages would still appear undervalued. This is even true for the outlier among the cases reported above, namely, Russia: applying the factor of 1.82 to the ratio of actual and equilibrium dollar wages in Table 3 still gives a ratio of only 0.73. Thus, in spite of the caveats discussed, it is hard to escape the overall conclusion that the currencies of the transition economies studied in this paper remain undervalued at least up to 1995, substantially so in many cases.

A final issue is that actual dollar wages in transition economies may underrepresent wage *costs* owing to differences in payroll taxes. From the firm perspective, payroll taxes are typically extremely high in economies at the early stages of transition, implying a downward bias in measured wage costs. Over time, this downward bias tends to become smaller as the payroll tax structure becomes more similar to that prevailing in market economies and nonwage benefits decline. As a result, our data may substantially overstate the gap between actual and equilibrium dollar wages in the initial years of transition, and create the illusion that this gap is closed fast.[37] However, most of this effect will be concentrated in the first few years of reforms, and as such it should, in most cases, not constitute a problem in assessing whether or not transition economies were wage competitive in the recent past. To the extent that specific countries are known to still have extraordinary high payroll taxes as of 1995 or 1996, a corresponding adjustment can be made individually.

Relative Competitiveness

The equilibrium wage argument pursued so far corresponds to a competitiveness concept that focuses on the attractiveness of a country to international capital flows. The implied perspective is that of an international

[37]We thank Chris Lane for this observation.

investor, who compares all countries and invests in those where dollar wages are lowest relative to productivity, taking into account a convex adjustment cost. However, this may not be the right perspective if we take the view that comparisons *within* the transition group are, for any reason, more relevant than comparisons between transition economies and developing market economies, or if we believe that it is more relevant to compare dollar wages and productivity within groups of actual trading partners rather than globally across all potential competitors. In this case, the right question is not whether dollar wages in a given transition economy are below equilibrium wages for that transition economy but rather whether dollar wages are *relatively more* below equilibrium in that economy than the dollar wages of its trading partners.

To answer this question, we constructed the following "index of relative competitiveness" for country i:

$$u_i + \frac{\frac{W_i^*}{W_i}}{\Pi_j\left(\frac{W_j^*}{W_j}\right)^{\theta_{ij}}}, \tag{20}$$

where w^*_I / w_i denotes the ratio of equilibrium and actual dollar wages for country i. θ_{ij} denotes country i's trade share with country j, so the denominator of equation (20) is a trade-weighted average of the equilibrium to actual dollar wage ratios of country i's trading partners. In practice, we picked the six most important trading partners for each transition economy in our sample based on 1994 export and import data from the IMF's Direction of Trade Statistics, and used 1994 trade weights throughout the sample period.[38] The index u_i is increasing in relative competitiveness; it takes a value smaller than one if country i does not have a cost-competitive edge over its trading partners, and greater than one if country i's cost-competitive position is favorable.

For transition economies with a large share of trade with other transition economies (as is the case for most CIS countries), the index u_i has the added advantage that it is less sensitive to omissions of transition-specific structural determinants of equilibrium wages, as such omissions would affect all

[38] Arguably, using the shares of partner countries in *total* trade can be misleading in the case of countries with very dissimilar import and export compositions, as these countries may not "compete" with their import partners in any meaningful sense. Specifically, in the case of countries that tend to import raw materials, but export mostly manufactures—a criterion met by most transition economies in our sample—manufacturing trade shares might constitute more meaningful weights than total trade weights. Unfortunately, data unavailability prevents us from calculating relative competitiveness indices based on manufacturing trade shares for most transition economies in the sample.

transition economies in the same direction. In other words, we may believe that the simple comparison between actual and estimated equilibrium wages exaggerates competitiveness for some or all CIS economies because of unmeasured idiosyncracies of those economies that tend to depress the equilibrium wage, but this argument would not apply if a country is judged competitive on the grounds of the index u_i. On the other hand, a disadvantage of using u_i relative to the earlier approach is that u_i is sensitive to mismeasurement of the actual wages among the country's trading partners, rather than just mismeasurement of its own wage. For example, gross underrecording of the average wage of a major trading partner can render a country "uncompetitive" based on u_i, even if its own wage is far from its correctly estimated equilibrium wage. In addition, u_i will typically be less precise than point estimates of equilibrium dollar wages, as its construction involves the use of several (typically seven) such estimates.

Tables 5 and 6 show the result of our calculations for Central and Eastern European countries, the Baltics, and CIS countries in our sample. Our results show all Central European transition economies as competitive during the 1991–95 period, although Hungary's competitiveness edge was never very large (and basically zero during 1993–94) and the Czech Republic's edge has narrowed continuously since 1991 (Table 5). For the Baltics and the CIS countries, a more mixed impression emerges. While the competitiveness margins are generally larger than those of the CEE economies, three countries—Kazakhstan, Lithuania, and Latvia—appear less competitive than their trading partners by the end of the period (u_i falls below the threshold value of one). For Latvia and Lithuania, this result is driven by fast dollar wage growth during 1993–95 relative to their main transition economy trading partners (especially Russia) and by the relatively low estimated equilibrium wage for Lithuania implied by IMF WEO estimates of PPP-adjusted GDP.[39] For Kazakhstan, the driving force is that actual dollar wages were of roughly the same magnitude as Russia's from about 1993 onward, while estimated equilibrium wages were substantially lower. Using PPP-adjusted GDP data from the EU and the World Bank changes these results in the case of Lithuania, whose relative competitiveness index rises above one (see below) but not for Kazakhstan and Latvia. Moreover, while the alternative GDP data imply a smaller estimated competitiveness margin for Estonia, this margin remains positive. Thus, the results of this subsection appear somewhat more robust to the use of PPP-adjusted GDP data from alternative sources than those of the previous subsection.

[39]Note, however, that the Lithuanian relative competitiveness index rises to about unity if it is based on manufacturing shares rather than total trade shares (using WEO PPP-adjusted GDP estimates). This is driven by a drop in the trade share with Russia from about 50 percent to about 25 percent if only manufacturing trade is considered, and a corresponding rise in the trade shares with Western Europe.

Table 5. *Competitive Position Relative to Trading Partners: Selected Central and Eastern European Countries*

		Index of relative competitiveness		
Country	Year	Specification (1)	Specification (2)	Specification (3)
Bulgaria	1991	3.74	3.29	3.20
	1992	2.02	1.93	1.76
	1993	1.78	1.69	1.56
	1994	2.67	2.49	2.32
	1995	2.44	2.22	2.12
Czech Republic	1991	2.27	1.81	1.97
	1992	2.07	1.73	1.81
	1993	1.83	1.51	1.58
	1994	1.72	1.35	1.47
	1995	1.58	1.27	1.36
Hungary	1990	1.46	1.02	1.35
	1991	1.29	0.97	1.20
	1992	1.20	1.00	1.13
	1993	1.06	0.92	1.00
	1994	1.16	1.00	1.09
	1995	1.31	1.10	1.22
Poland	1990	2.37	1.97	2.10
	1991	1.63	1.51	1.47
	1992	1.44	1.35	1.30
	1993	1.50	1.41	1.35
	1994	1.55	1.39	1.37
	1995	1.48	1.31	1.30
Romania	1991	1.48	1.24	1.34
	1992	2.20	1.94	2.02
	1993	2.01	1.75	1.83
	1994	2.12	1.78	1.91
	1995	1.92	1.57	1.71
Slovak Republic	1991	1.38	1.50	1.31
	1992	1.26	1.39	1.19
	1993	1.31	1.41	1.23
	1994	1.37	1.45	1.28
	1995	1.38	1.38	1.27

Note: Indices prior to 1992 use 1992 information on Russia's competitiveness.

Competitiveness in 1996

We now assemble the main estimation results in this paper to evaluate the competitive position of transition economies using the most recent available annual wage and exchange rate data. In doing so, the main limitation is that we are forced to use equilibrium wage estimates out of sample, that is, we need to compare 1995 equilibrium wage estimates with 1996

Table 6. *Competitive Position Relative to Trading Partners: Baltic and Selected CIS Countries*

		Index of relative competitiveness		
Country	Year	Specification (1)	Specification (2)	Specification (3)
Belarus	1992	2.14	1.76	2.02
	1993	3.37	2.65	3.15
	1994	4.16	3.35	3.92
	1995	1.94	1.55	1.82
Estonia	1992	4.19	3.50	3.68
	1993	2.64	2.26	2.33
	1994	2.10	1.81	1.86
	1995	1.70	1.44	1.49
Kazakhstan	1992	1.30	1.43	1.32
	1993	0.74	0.79	0.73
	1994	0.96	1.05	0.94
	1995	0.65	0.74	0.63
Kyrgyz Republic	1992	1.77	1.85	1.97
	1993	1.92	2.10	2.12
	1994	1.39	1.54	1.54
	1995	1.26	1.43	1.39
Latvia	1992	1.73	1.41	1.52
	1993	1.89	1.63	1.68
	1994	1.28	1.08	1.13
	1995	1.09	0.91	0.96
Lithuania	1992	2.90	2.74	2.76
	1993	1.30	1.26	1.24
	1994	1.08	1.05	1.03
	1995	0.91	0.85	0.85
Moldova	1992	1.47	1.65	1.44
	1993	1.77	1.87	1.71
	1994	1.25	1.48	1.22
	1995	1.14	1.32	1.11
Russia	1992	5.20	4.70	4.57
	1993	2.03	1.92	1.80
	1994	1.56	1.60	1.41
	1995	1.78	1.94	1.63
Ukraine	1992	1.66	1.45	1.59
	1993	3.22	3.04	3.14
	1994	2.25	1.94	2.15
	1995	1.57	1.38	1.51

dollar wage data. The justification is that, from Tables 3 and 4, by 1995 equilibrium dollar wages are either flat or rising in almost all countries. Thus, if actual dollar wages in 1996 continue to be significantly below 1995 estimated equilibrium wages (or the "relative competitiveness index" using 1995 equilibrium and 1996 actual wage data continues to be significantly larger than one), we can generally infer that the economy under discussion was indeed competitive in 1996.

Table 7 shows actual 1996 dollar wage data for our group of countries and compares it with 1995 estimated equilibrium wages using PPP-adjusted GDP data from the WEO, and a relative competitiveness index based on these estimates and 1996 actual wages, for each of the three RE specifications discussed in the previous subsections. Table 8 shows how these estimates would be affected if World Bank or EU GDP data were used instead of WEO data.[40] Thus, the tables capture the sensitivity of our results along two dimensions: (1) the choice of normalization of PPP-adjusted GDP; and (2) the source of PPP-adjusted GDP data. Information on competitiveness is provided from two perspectives. First, one can compare actual wages in transition economies with estimated equilibrium wages, which are the wages one would expect in an average developing market economy with identical right-hand side "fundamentals" as the transition economy we are studying. Since this average is taken over all 85 economies in our sample, the implicit assumption is that the transition economy of interest is in a potential competitive situation with all other economies that we use in our estimation. The second perspective is one where the gap between actual and estimated equilibrium dollar wages matters only relative to the corresponding trade-weighted gap between actual and equilibrium wages among the economy's trade partners. This is expressed in the competitiveness index, which takes a value greater than one if the economy's dollar wages are more undervalued (or less overvalued) than that of its trading partners. The implicit assumption is that the economy is in a competitive situation only with its trading partners; however, rather than directly comparing dollar wages within this group we compare "undervaluation gaps" to control for differences in fundamentals across countries.

The main conclusions from Tables 7 and 8 are as follows:

- Based on the direct comparison between actual and 1995 equilibrium dollar wages, none of the countries surveyed appears overvalued in the

[40]This involves reestimating the equations and calculating fitted wages and relative competitiveness indices using the alternative PPP-adjusted GDP numbers for transition countries, and WEO data for nontransition countries. To address the problem that the three data sources are based on somewhat different sets of international prices and goods baskets, we scaled the EU and World Bank data such that their 1995 PPP-adjusted GDP for Poland is identical to the 1995 WEO estimate. Without this adjustment, the EU-based estimates in Table 8 would be approximately 8 percent higher and the World Bank based estimates approximately 15 percent lower than shown in Table 8.

Table 7. *Estimated Competitive Position in 1996 Using Alternative Specifications: Selected Transition Economies*

Country	Actual wage	1995 estimated equilibrium wage			Competitiveness index [a]		
		Spec. 1	Spec. 2	Spec. 3	Spec. 1	Spec. 2	Spec. 3
Bulgaria	101	358	345	306	3.44	3.12	2.98
Czech Republic	349	566	480	462	1.54	1.23	1.32
Hungary	305	438	396	403	1.52	1.28	1.42
Poland	319	449	428	391	1.51	1.34	1.33
Romania	104	252	218	223	2.18	1.78	1.95
Slovakia	265	477	456	395	1.45	1.45	1.33
Estonia	248	444	399	360	1.74	1.47	1.53
Latvia	236	308	279	251	1.22	1.01	1.07
Lithuania	177	194	196	167	0.90	0.84	0.85
Belarus	101	333	292	276	2.33	1.86	2.18
Kazakhstan	156	161	203	140	0.74	0.84	0.72
Kyrgyz Republic	58	132	168	129	1.74	1.97	1.92
Moldova	66	123	147	106	1.12	1.31	1.09
Russia	188	269	300	235	1.19	1.31	1.09
Ukraine	75	166	164	143	1.55	1.42	1.48

[a] Based on 1996 actual wages and 1995 estimated equilibrium wages.

sense that actual wages *exceed* the estimated equilibrium wage range. On the other hand, some countries—specifically, Latvia, Lithuania and Kazakhstan—fall *within* this estimated range for *some* specifications and some PPP-adjusted GDP per capita estimates. Generally, the tables thus suggest that some scope for real appreciation remains, although its extent varies greatly across countries.

- Based on the relative competitiveness index, the emerging picture is more complex. All Central and Eastern European transition economies appear to have maintained their competitiveness edge in 1996. This includes the Czech Republic; thus, the Czech currency crisis would not have been predicted on the basis of our estimates. The Baltic and CIS economies, on the other hand, show much greater heterogeneity. Based on WEO data, only Lithuania and Kazakhstan would seem to have lost their edge in 1996. Based on World Bank and EU data, Lithuania appears competitive (in particular, if manufacturing trade shares are used, see footnote 39) but Latvia does not, and Estonia's position looks less comfortable. For Kazakhstan, the picture is unchanged.

Table 8. *Estimated Competitive Position in 1996 Using Alternative GDP Estimates: Selected Transition Economies*

Country	Actual wage	1995 estimated equilibrium wage			Competitiveness index [a]		
		WEO	EU	WB	WEO	EC	WB
Bulgaria	101	306–358	293–334	310–360	2.98–3.44	2.81–3.22	2.75–3.15
Czech Republic	349	462–566	560–724	566–734	1.23–1.54	1.40–1.88	1.52–2.12
Hungary	305	396–438	434–506	434–504	1.28–1.52	1.39–1.74	1.34–1.63
Poland	319	391–449	391–447	394–449	1.33–1.51	1.31–1.48	1.27–1.43
Romania	104	218–252	278–319	258–329	1.78–218	2.05–2.73	2.02–2.64
Slovakia	265	395–477	460–559	282–361	1.33–145	1.37–1.49	0.81–1.03
Estonia	248	360–444	287–346	302–363	1.47–1.74	1.19–1.33	1.14–1.26
Latvia	236	251–308	225–273	239–290	1.01–1.22	0.91–1.02	0.84–0.93
Lithuania	177	167–194	272–328	270–323	0.85–0.90	1.16–1.51	1.01–1.21
Belarus	101	276–333	...	284–340	1.86–2.33	...	1.55–1.73
Kazakhstan	156	140–203	...	205–260	0.72–0.84	...	0.78–0.89
Kyrgyz Republic	58	129–168	...	157–190	1.74–1.97	...	1.56–1.84
Moldova	66	106–147	...	...	1.09–1.31	...	...
Russia	188	235–300	...	339–394	1.09–1.31	...	1.49–1.65
Ukraine	75	143–166	...	174–203	1.42–1.55	...	1.37–1.46

[a] Based on 1996 actual wages and 1995 estimated equilibrium wages. Ranges are based on results from different specifications.

III. Conclusion

This paper finds that even though dollar wages in transition economies generally continue to remain below their "equilibrium" level, the scope for continuing real appreciation varies greatly across the transition countries examined and to some extent depends on the competitiveness concept adopted and the productivity data used. In addition, the case of the Czech Republic serves as a reminder that the cost-competitiveness indicators studied above should not be relied on for the purposes of predicting currency crises, which depend on a much broader range of variables driving short term capital movements.[41]

With these limitations in mind, we conclude that (1) most countries in our sample continue to have scope for real appreciation; and (2) for a minority of countries, this scope may be small or nil and enhanced vigilance is appropriate. Subject to the caveats regarding productivity and wage data which were discussed at length in the course of the paper, this set would appear to include Kazakhstan, and perhaps the Baltic economies. One would thus

[41]For a recent survey in this area, see Kaminsky, Lizondo, and Reinhardt (1998)

expect competitiveness considerations to gain greater prominence in the discussion of exchange rate policy in these countries.

APPENDIX

Testing for Strict Exogeneity

To test for strict exogeneity, we used a "Hausman-like" test on the FE and FD estimates. Under the null hypothesis of strict exogeneity, both FE and FD estimates are consistent, but FE is more efficient. Under the alternative, only FD is consistent. The test statistic in our particular case is

$$T_{FE,FD} = \left(A\hat{a}^{FE} - \hat{a}^{FD}\right)' Var\left(A\hat{a}^{FE} - \hat{a}^{FD}\right)^{-1}\left(A\hat{a}^{FE} - \hat{a}^{FD}\right) \tag{A1}$$

where $\mathbf{\hat{a}}^{FE}$ and $\mathbf{\hat{a}}^{FD}$ are vectors of unequal length[42] containing the FE and FD parameter estimates based on equation (19), respectively, **A** is the matrix that transforms the true parameter vector $\mathbf{\hat{a}}^{FE}$ into $\mathbf{\hat{a}}^{FD}$, and $Var(\mathbf{\hat{a}}^{FE} - \mathbf{\hat{a}}^{FD})$ is a consistent estimate of the variance-covariance matrix of the difference between the transformed FE and the FD coefficient estimators. Under the null, the test statistics' distribution is $\chi^2(k)$, where k is the rank of A (or, equivalently, the dimension of $\mathbf{\hat{a}}^{FD}$, that is, 11 in our case).

The variance-covariance matrix was estimated along the lines of a method suggested by Keane and Runkle (1992), assuming that the residuals are uncorrelated across individuals (countries) but allowing a general covariance structure for each country, which is assumed to be the same for all countries. In particular, we first used a Newey-West (1987) procedure to estimate a within-country variance-covariance matrix of the residuals, denoted FE in the FE case (of dimension 6×6) and FD in the FD case (dimension 5×5). Since our panel was unbalanced, we then chose the appropriate submatrix of FE and FD for each country to construct the estimated variance-covariance matrix of the residuals for FE and FD, respectively. These matrices take the form of "quasi-diagonal" matrices of dimension 392×392 in the FE and 307×307 in the FD case, where the quasi-diagonal contains a sequence of submatrices of FE and FD, respectively, and all other elements are zero by assumption. Similar assumptions and procedures were applied to construct an estimate of the 392×307 cross-covariance matrix of the residuals in the two specifications.

Treatment of Transition Dummies in the FE and FD Specifications

Consider a version of equation (19) where, for simplicity, we have ignored the *oecd* dummy and the distinction between *cee* and *fsu* dummies and have assumed a three-period panel:

[42] See below.

$$w_{\$i,t} = b_0 + b_1 x_{i,t} + b_2 d_{1i,t} + b_3 d_{2i,t} + b_4 d_{3i,t} + \mu_i + \varepsilon_{i,t} \tag{A2}$$

where $x_{i,t}$ represents the economic variables in equation (19) and the dummies $d_{\tau i,t}$ take on the value 1 if $t = \tau$ and i is a transition economy and 0 otherwise.

Suppose now that we try to estimate equation (A2) using the FD or FE estimators. Trivially, if we mechanically first-difference or groupwise demean the data, this will transform the full set of linearly independent dummies into a set of linearly dependent vectors, creating perfect multicollinearity. However, suppose we first-difference or demean both the left-hand side and right-hand side economic data and then go on to running OLS after adding a *full* set of time-varying dummies, as before:

$$\Delta w_{\$i,t} = b_1^{FD} \Delta x_{i,t} + b_3^{FD} d_{2i,t} + b_4^{FD} d_{3i,t} + \Delta \varepsilon_{i,t}, \tag{A3}$$

where Δ denotes first differences; for example, $\Delta w_{i,t} = w_{i,t-l}$, with t running from 2 to 3, and

$$\tilde{w}_{\$i,t} = b_1^{FE} \tilde{x}_{i,t} + b_2^{FE} d_{1i,t} + b_3^{FE} d_{2i,t} + b_4^{FE} d_{3i,t} + \tilde{\varepsilon}_{i,t}, \tag{A4}$$

where the ~ denotes demeaned data and $t = 1 \ldots 3$. It is easy to show that $b_l = b_l^{FD} = b_l^{FE}$ and that the coefficients on the dummy variables are related by the condition $b_j^{FD} = b_j - b_{j-1} = b_j^{FE} - b_{j-1}^{FE}$ for $j = 3,4$. Thus, while the dummy coefficients of equations (A2) and (A4) can easily be transformed into the dummy coefficients of equation (A3), knowing b_j^{FD} ($j= 1,3,4$) is insufficient to recover the corresponding coefficients b_j ($j = 0 \ldots 4$) or b_j^{FE} ($j =1 \ldots 4$). Similarly, knowledge of b_j^{FE} if ($j =1 \ldots 4$) is insufficient to determine b_j ($j = 0 \ldots 4$) and vice versa. Thus, neither the first difference nor the fixed effects parameter estimates enable us to compute equilibrium dollar wage levels for transition economies, because neither allows us to recover the parameter b_0 of equation (A2). Put differently, even though we set the transition dummies to zero, when computing estimated equilibrium wages for transition economies we need to have coefficient estimates for these dummies, because without this information we cannot determine the intercept b_0, which we require for the purposes of estimating equilibrium wages.[43] The problem is that these dummies are impossible to estimate through either the FD or FE techniques.

Note, however, that by making use of the conditions relating the parameters of equations (A2), (A3), and (A4), we can perform specification tests based on the comparison of these parameters in a largely standard fashion. What we need to do is to perform Hausman-type tests on the transformations of the parameter estimates which we know would generate identical vectors if the estimates were equal to the true parameters, and thus should be close if both of the alternative estimators are consistent. These tests are described in the first section of this Appendix.

[43] Note that in the more general equation we actually estimate, which includes a time-invariant OECD dummy, the coefficient on that dummy could not be recovered either

Table A1. *Data Coverage*

	Country	Dates		Country	Dates
1	Argentina	1990–94	44	Kyrgyz Republic	1992–95
2	Australia	1990–95	45	Latvia	1992–95
3	Austria	1990–95	46	Lithuania	1992–95
4	Bahrain	1990–94	47	Luxembourg	1990–93
5	Barbados	1990–91	48	Malawi	1990–91
6	Belarus	1992–95	49	Malaysia	1990–92
7	Belgium	1990–95	50	Malta	1990–93
8	Bolivia	1990–94	51	Mauritius	1990–94
9	Botswana	1990–94	52	Mexico	1990–95
10	Brazil	1990–94	53	Moldova	1992–95
11	Bulgaria	1991–95	54	Namibia	1992
12	Cameroon	1990–92	55	Netherlands	1990–95
13	Canada	1990–95	56	New Zealand	1990–95
14	Chile	1990–95	57	Nicaragua	1991–95
15	Colombia	1990–95	58	Norway	1990–95
16	Costa Rica	1990–94	59	Pakistan	1991
17	Cyprus	1990–93	60	Panama	1990–92
18	Czech Republic	1991–95	61	Paraguay	1990–94
19	Denmark	1990–95	62	Peru	1991–94
20	Ecuador	1990–93	63	Philippines	1990–91
21	Egypt	1990–92	64	Poland	1990–95
22	El Salvador	1990–94	65	Portugal	1990–95
23	Estonia	1992–95	66	Romania	1991–95
24	Fiji	1990–94	67	Russia	1992–95
25	Finland	1990–95	68	Seychelles	1990–94
26	France	1990–95	69	Singapore	1990–94
27	Germany	1990–95	70	Slovak Republic	1991–95
28	Ghana	1990–91	71	South Africa	1990–94
29	Greece	1990–95	72	Spain	1990–95
30	Guatemala	1990–95	73	Swaziland	1990–92
31	Hong Kong	1992–95	74	Sweden	1990–95
32	Hungary	1990–95	75	Switzerland	1990–95
33	India	1990	76	Taiwan Prov. of China	1990–95
34	Indonesia	1990–91	77	Thailand	1990–94
35	Ireland	1990–95	78	Trinidad and Tobago	1990–93
36	Israel	1990–95	79	Turkey	1990–93
37	Italy	1990–95	80	Ukraine	1992–95
38	Jamaica	1990–93	81	United Kingdom	1990–95
39	Japan	1990–95	82	United States	1990–95
40	Jordan	1990–93	83	Venezuela	1990–93
41	Kazakhstan	1992–95	84	Zambia	1991–92
42	Kenya	1990–91	85	Zimbabwe	1990–94
43	Korea	1990–95			

Table A2. *Estimated Coefficients and Specification Test Results, Restricted Sample*
(Dependent variable: monthly average wages in manufacturing, in U.S. dollars; standard errors are in parentheses)

Variable	Specification (1)			Specification (2)			Specification (3)		
	FD	FE	RE	FD	FE	RE	FD	FE	RE
gdp	0.828	0.954	0.727	0.702	0.883	0.572	0.820	0.962	0.700
	(0.197)	(0.124)	(0.099)	(0.185)	(0.132)	(0.104)	(0.203)	(0.131)	(0.102)
agr	0.139	–0.087	–0.101	0.104	–0.119	–0.207	0.135	–0.083	–0.137
	(0.109)	(0.082)	(0.062)	(0.109)	(0.083)	(0.060)	(0.110)	(0.083)	(0.061)
school	0.092	–0.218	0.184	0.294	–0.098	0.308	0.207	–0.059	0.153
	(0.700)	(0.365)	(0.133)	(0.696)	(0.373)	(0.139)	(0.695)	(0.359)	(0.139)
oecd			0.676			0.911			0.792
			(0.154)			(0.159)			(0.154)
cee90		–0.388	–1.370		–0.273	–1.053		–0.394	–1.199
		(0.092)	(0.242)		(0.091)	(0.264)		(0.093)	(0.248)
cee91	0.305	–0.233	–1.182	0.264	–0.177	–0.948	0.305	–0.237	–1.022
	(0.117)	(0.049)	(0.222)	(0.116)	(0.050)	(0.246)	(0.118)	(0.049)	(0.230)
cee92	0.243	–0.043	–1.002	0.201	–0.035	–0.823	0.242	–0.045	–0.851
	(0.071)	(0.049)	(0.220)	(0.069)	(0.050)	(0.244)	(0.071)	(0.049)	(0.229)

cee93	0.152	0.100	−0.856	0.120	0.067	−0.704	0.153	0.099	−0.704
	(0.064)	(0.049)	(0.220)	(0.066)	(0.049)	(0.242)	(0.064)	(0.049)	(0.229)
cee94	−0.007	0.088	−0.855	−0.006	0.051	−0.699	−0.003	0.090	−0.699
	(0.065)	(0.048)	(0.219)	(0.066)	(0.049)	(0.241)	(0.065)	(0.048)	(0.228)
cee95	0.139	0.218	−0.710	0.148	0.186	−0.542	0.142	0.223	−0.548
	(0.065)	(0.048)	(0.219)	(0.066)	(0.051)	(0.239)	(0.066)	(0.049)	(0.226)
fsu92		−0.730	−2.315		−0.638	−2.067		−0.734	−2.127
		(0.045)	(0.222)		(0.043)	(0.250)		(0.046)	(0.232)
fsu93	0.451	−0.272	−1.883	0.403	−0.249	−1.682	0.449	-0.274	−1.699
	(0.058)	(0.039)	(0.224)	(0.055)	(0.040)	(0.253)	(0.058)	(0.040)	(0.236)
fsu94	0.563	0.286	−1.355	0.513	−0.259	−1.208	0.564	0.288	−1.175
	(0.061)	(0.041)	(0.225)	(0.056)	(0.041)	(0.256)	(0.061)	(0.042)	(0.240)
fsu95	0.439	0.716	−0.927	0.404	0.647	−0.813	0.441	0.721	−0.748
	(0.053)	(0.042)	(0.225)	(0.053)	(0.041)	(0.254)	(0.053)	(0.043)	(0.240)
constant			−0.851			−0.493			−1.055
			(0.848)			(1.003)			(0.918)
N	254	323	323	254	323	323	254	323	323
R^2	0.89	0.71	0.72	0.46	0.70	0.88	0.46	0.71	0.89

Notes: Restricted sample comprises only countries for which employee wage data was available. Specifications (1), (2), and (3) refer to normalization of GDP by population, employment, and the labor force, respectively. Coefficients on cee and fsu dummies are not directly comparable across FD, FE, and RE estimates, see Appendix II.

REFERENCES

Aghion, Philippe, and Wendy Carlin, 1996, "Restructuring Outcomes and the Evolution of Ownership Patterns in Central and Eastern Europe," *Economics of Transition*, Vol. 4 (October), pp. 371–88.

Barro, Robert J., and Jong Wha Lee, 1996, "International Measures of Schooling Years and Schooling Quality," *American Economic Review, Papers and Proceedings*, Vol. 86 (May), pp. 218–23.

Beleva, Iskra, Richard Jackman, and Mariela Nenova-Amar, 1995, "Bulgaria," in *Unemployment, Restructuring and the Labor Market in Eastern Europe and Russia*, ed. by Simon Commander and Fabrizio Coricelli (Washington: World Bank), pp. 193–232.

Blanchard, Olivier, Simon Commander, and Fabrizio Coricelli, 1995, "Unemployment and Restructuring in Eastern Europe and Russia," in *Unemployment, Restructuring and the Labor Market in Eastern Europe and Russia*, ed. by Simon Commander and Fabrizio Coricelli (Washington: World Bank), pp. 289–330.

Calvo, Guillermo A., Ratna Sahay, and Carlos A. Végh, 1996, "Capital Flows in Central and Eastern Europe: Evidence and Policy Options," in *Private Capital Flows to Emerging Markets After the Mexican Crisis*, ed. by Guillermo A. Calvo, Morris Goldstein, and Eduard Hochreiter (Washington: Institute for International Economics).

Clague, Christopher K., 1986, "Determinants of the National Price Level: Some Empirical Results," *Review of Economics and Statistics*, Vol. 68 (May), pp. 320–23.

———, 1988, "Purchasing-Parities and Exchange Rates in Latin America," *Economic Development and Cultural Change*, Vol. 36, April 1988, 529–41.

———, and Vito Tanzi, 1972, "Human Capital, Natural Resources and the Purchasing-Power Parity Doctrine: Some Empirical Results," *Economia Internazionale*, Vol. 25, (February), pp. 3–18.

Clark, Peter, Leonardo Bartolini, Tamim Bayoumi, and Steven Symansky, 1994, *Exchange Rates and Economic Fundamentals: A Framework for Analysis*, IMF Occasional Paper No. 115 (Washington: International Monetary Fund).

Commander, Simon, and Richard Jackman, 1993, "Providing Social Benefits in Russia. Redefining the Roles of Firms and Government," World Bank Policy Research Working Paper No. 1184 (Washington: World Bank, September).

Debelle, Guy, and Hamid Faruqee, 1996, "What Determines the Current Account? A Cross-Sectional and Panel Approach," IMF Working Paper 96/58 (Washington: International Monetary Fund).

Dollar, David, 1992, "Outward-Oriented Developing Economies Really Do Grow More Rapidly: Evidence from 95 LDCs, 1976–1985," *Economic Development and Cultural Change*, Vol. 40 (April), pp. 523–44.

Dornbusch, Rudiger, 1983, Real Interest Rates, Home Goods and Optimal External Borrowing, *Journal of Political Economy*, Vol. 91 (February), pp. 141–53.

EBRD, 1994, *Transition Report* (London: European Bank for Reconstruction and Development).

_______, 1995, *Transition Report* (London: European Bank for Reconstruction and Development).

_______, 1996, *Transition Report* (London: European Bank for Reconstruction and Development).

Estrin, Saul, M.E. Schaffer and I.J. Singh, 1995, "The Provision of Social Benefits in State-Owned, Privatized and Private Firms in Poland," Centre for Economic Performance Discussion Paper No. 223 (London: London School of Economics and Political Science).

Faruqee, Hamid, 1995, "Long-Run Determinants of the Real Exchange Rate: A Stock-Flow Perspective," *Staff Papers*, International Monetary Fund, Vol. 42 (March), pp. 80–107.

Feldman, Robert A., 1995, "External Competitiveness," in *United Germany: The First Five Years*, IMF Occasional Paper No. 125 (Washington: International Monetary Fund), pp. 9–20.

Flanagan, Robert J., 1995, "Wage Structures in the Transition of the Czech Economy," IMF Working Paper 95/36 (Washington: International Monetary Fund).

Green, William H., 1993, *Econometric Analysis* (New York: Macmillan, 2nd. ed.).

Halpern, László, and Charles Wyplosz, 1997, "Equilibrium Exchange Rates in Transition Economies," IMF *Staff Papers*, Vol 44 (December 1977), pp. 430–61.

Ham, John, Jan Svejnar and Katherine Terrell, 1995, "Czech Republic and Slovakia," in *Unemployment, Restructuring and the Labor Market in Eastern Europe and Russia*, ed. by Simon Commander and Fabrizio Coricelli (Washington: World Bank), pp. 91–146.

Hausman, Jerry, 1978, "Specification Tests in Econometrics," *Econometrica*, Vol. 46, pp. 1251–71.

Havlik, Peter, 1996, "Exchange Rates, Competitiveness and Labour Costs in Central and Eastern Europe," Vienna Institute for Comparative Economic Studies (WIIW), October 1996.

Isard, Peter, and Hamid Faruqee, eds., 1998, "Exchange Rate Assessment: Some Recent Extensions and Applications of the Macroeconomic Balance Approach," forthcoming as IMF Occasional Paper (Washington: International Monetary Fund).

Isenman, Paul, 1980, "Inter-Country Comparison of 'Real' (PPP) Incomes: Revised Estimates and Unresolved Questions," *World Development*, Vol. 8, pp. 61–72.

Kaminsky, Graciela, Saul Lizondo and Carmen M. Reinhardt, 1998, "Leading Indicators of Currency Crises," *Staff Papers*, International Monetary Fund, Vol. 45 (March), pp. 1–48.

Keane, Michael P., and David E. Runkle, 1992, "On the Estimation of Panel Data Models with Serial Correlation When Instruments Are Not Strictly Exogenous," *Journal of Business and Economic Statistics*, Vol. 10 (January), pp. 433–41.

Krajnyák, Kornélia and Jeromin Zettelmeyer, 1997, "Competitiveness in Transition Economies: What Scope for Real Appreciation?" IMF Working Paper 97/149 (Washington: International Monetary Fund).

Kramer, Charles, 1996, "FEERs and Uncertainty: Confidence Intervals for the Fundamental Equilibrium Exchange Rate of the Canadian Dollar," IMF Working Paper 96/68 (Washington: International Monetary Fund).

Kravis, Irving B., and Robert E. Lipsey, 1987, "The Assessment of National Price Levels," in *Real-Financial Linkages among Open Economies*, ed. by Sven W. Arndt and J. David Richardson (Cambridge, Massachusetts: MIT Press), pp. 97–134.

———, 1988, "National Price Levels and the Prices of Tradables and Nontradables," *American Economic Review, Papers and Proceedings*, Vol. 78 (May), pp. 474–78.

Lipschitz, Leslie, and Donogh McDonald, 1992, "Real Exchange Rates and Competitiveness: A Clarification of Concepts, and Some Measurements for Europe," *Empirica—Austrian Economic Papers*, Vol. 19 (No.1), pp. 37–69.

Marsh, Ian W., and Stephen P. Tokarick, 1994, "Competitiveness Indicators: A Theoretical and Empirical Assessment," IMF Working Paper 94/29 (Washington: International Monetary Fund).

Newey, Whitney, and Kenneth West, 1987, "A Simple, Positive Semi-Definite, Heteroscedasticity and Autocorrelation Consistent Covariance Matrix," *Econometrica*, Vol. 55, pp. 703–708.

Obstfeld, Maurice, and Kenneth Rogoff, 1996, *Foundations of International Macroeconomics* (Cambridge, Massachusetts: MIT Press).

Officer, Lawrence H., 1989, "The National Price Level: Theory and Estimation," *Journal of Macroeconomics*, Vol. 11 (Summer), pp. 351–73.

Richards, Anthony J., and Gunnar H.R. Tersman, 1996, "Growth, Nontradables, and Price Convergence in the Baltics," *Journal of Comparative Economics*, Vol. 23 (October), pp. 121–45.

Roubini, Nouriel, and Paul Wachtel, 1998, "Current Account Sustainability in Transition Economies," NBER Working Paper 6468 (Cambridge: Massachusetts: National Bureau of Economic Research, May).

Rutkowski, Jan, 1994, "Labor Market Transition and Changes in the Wage Structure: The Case of Poland," PPRG Discussion Paper No.32 (Warsaw: Polish Policy Research Group, Warsaw University, September).

———, 1996, "High Skills Pay Off: The Changing Wage Structure During Economic Transition in Poland," *Economics of Transition* (EBRD),Vol. 4 (May), pp. 89–112.

Vecerník, Jiri, 1995, "Changing Earnings Distribution in the Czech Republic: Survey Evidence from 1988-1994," *Economics of Transition* (EBRD), Vol. 3 (September), pp. 355–71.

Wagner, Nancy, 1995, "A Review of PPP-Adjusted GDP Estimation and Its Potential Use for the Fund's Operational Purposes," IMF Working Paper 95/18 (Washington: International Monetary Fund).

World Bank, 1997, *World Development Report 1997* (Washington: World Bank).