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Do Long-Run Productivity Differentials Explain Long-Run Real Exchange Rates?

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Abstract

We develop a two-country, balanced-growth intertemporal general equilibrium model to examine two predictions of the Balassa-Samuelson model, namely that (i) productivity differentials determine the domestic relative price of nontradables and (ii) deviations from purchasing power parity reflect differences in the relative price of nontradables. In our model, the equilibrium relative price of nontradables along the long-run balanced-growth path is determined by the ratio of the marginal products of labor in the tradable and nontradable sectors. The empirical relevance of the Balassa-Samuelson predictions is examined using the Hodrick-Prescott filter to extract long-run components from a panel database for fourteen OECD countries. The evidence indicates that labor productivity differentials do explain long-run, cross-country differences in relative prices. The predicted relative prices, however, are of little help in explaining long-run deviations from purchasing power parity.

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Summary

In celebration of thirty years of the Balassa-Samuelson model, this paper attempts to provide an appraisal of the static theory of Balassa (1964) and Samuelson (1964) by embedding it in an explicitly dynamic general equilibrium setting. The paper's appraisal of this model focuses on two of its key implications; namely that, (i) cross-country differences in the relative price of nontradables reflect differences in the relative marginal productivity of labor of tradable and nontradable sectors, and (ii) cross-country differences in the level of real exchange rates are explained by differences in the relative price of nontradables. These two propositions are developed as long-run, balanced-growth, implications of a two-country intertemporal equilibrium model and several tests are conducted to examine their empirical relevance. For the empirical analysis the authors identify restrictions imposed on the cross-sectional, low-frequency behavior of the data implied by the model, and construct a cross-country sectoral database from existing OECD data to conduct econometric tests based on panel data methods.

The empirical analysis suggests that the Balassa-Samuelson proposition that cross-country differences in long-run domestic relative prices of nontradables are determined by differences in the ratio of long-run sectoral marginal products of labor cannot be rejected by the data. However, the analysis also indicates that long-run relative prices (as measured in the data or as predicted by regressions) are of little help in explaining long-run, cross-country differences in the level of real exchange rates based on CPIs or GDP deflators. Thus, while the paper finds that the Balassa-Samuelson general equilibrium model performs well as a theory of relative prices, it indicates that the model seems unable to account for long-run deviations from PPP. The authors state that this finding echoes a quotation by Paul Samuelson that prefaces the paper: "Unless very sophisticated indeed, PPP is a misleading pretentious doctrine, promising us what is rare in economics, detailed numerical predictions."

"Unless very sophisticated indeed, PPP is a misleadingly pretentious doctrine, promising us what is rare in economics, detailed numerical predictions". [Paul Samuelson (1964)]

I. Introduction

In two seminal papers, Balassa (1964) and Samuelson (1964), independently argued that labor productivity differentials between tradable and nontradable sectors will lead to changes in real costs and relative prices, bringing about divergences in exchange rate adjusted national price levels. 1/ In the last thirty years this insight has been the guiding principle for most theoretical and empirical research on real exchange rates.

Several different predictions of the Balassa-Samuelson model have been explored in the literature. 2/ Some empirical studies have focused on Balassa's finding that real exchange rates bear a strong positive relationship to the level of output per-capita across countries. Others examine the relevance of sectoral inflation differentials in explaining differences in real exchange rates. 3/ Furthermore, several theoretical papers have focused on the determinants of the equilibrium relative price of nontradables in intertemporal models (Dornbusch, 1983; Greenwood 1984).

However, surprisingly, little empirical work has been carried out on developing intertemporal equilibrium models to investigate the predictions of the Balassa-Samuelson model. Exceptions are Rogoff (1991) and Obstfeld (1993). Obstfeld provides evidence of deterministic trends in real exchange rates for Japan and the United States. He develops a small open economy model with unbalanced growth to capture this important stylized fact. Our analysis differs from his in that we model a two-country world with balanced-growth in which long-run relative price differentials reflect differentials in factor productivity growth. 4/ For the empirical analysis we focus on differences *across* countries in long-run levels of real exchange rates and domestic relative prices of nontradable goods.

1/ Hereafter, by "relative price" we mean the price of nontradables relative to tradables with tradables acting as the numeraire.

2/ For want of a unified name in the literature we have chosen to refer to the arguments supporting the empirical regularities observed by Balassa (1964) and Samuelson (1964) as the Balassa-Samuelson model. Elsewhere in the literature it has been called either the Balassa effect, the Balassa-Ricardo effect or the productivity bias hypothesis.

3/ For recent empirical studies along these lines see De Gregorio, Giovannini and Wolf (1994) and Micosi and Milesi (1993). See also Hsieh (1982), Kravis, Heston and Summers (1983), Kravis and Lipsey (1987), Marston (1987), Yoshikawa (1990) and Bergstrand (1991) for other empirical tests of the predictions of the Balassa-Samuelson model.

4/ In our model sectoral output, consumption and investment grow at the same rate. There is still a differential in total factor productivity growth, however, to the extent that labor shares in the tradable and nontradable sectors differ.

Thus, unlike Obstfeld (1993), we are concerned with the cross-sectional implications of the Balassa-Samuelson model rather than its time series implications.

In a closely related strand of the intertemporal equilibrium literature, Stockman and Tesar (1990) and Mendoza (1992) have studied the quantitative implications of multisector equilibrium models of the business cycle. The authors use numerical methods popularized in the real business cycle literature to evaluate the role of productivity shocks and terms-of-trade disturbances in determining the cyclical properties of the relative price of nontradables and the real exchange rate. In a recent contribution to this literature, Backus and Smith (1993) derive closed-form solutions linking deviations from purchasing power parity (PPP) and real interest parity to international consumption patterns. They use a two-country general equilibrium exchange economy to examine the possibility that nontraded goods may explain the persistent deviations from PPP observed in the data.

This paper aims to contribute to the empirical literature analyzing real exchange rates from a dynamic general equilibrium perspective. Our objective is to examine two basic propositions of the Balassa-Samuelson model, namely that: (i) productivity differentials determine the domestic relative price of nontradables and, (ii) productivity differentials explain deviations from PPP. We carry out the analysis in the context of a two-country, balanced-growth model driven by labor-augmenting (Harrod-neutral) technological progress. The Balassa-Samuelson propositions are derived as *long-run* implications of the model and closed-form solutions are obtained for the long-run relative price of nontradables and the real exchange rate.

We show that along the long-run balanced-growth path, the relative price of nontradables is determined by the ratio of the marginal products of labor in the tradable and nontradable sectors. Assuming Cobb-Douglas technologies, this ratio can be expressed as a log-linear function of sectoral capital-output ratios or of the investment-output ratio in the tradable sector. The investment-output ratio is shown to be a function of exogenous parameters describing preferences and technology. We then derive three empirically implementable equations from this version of the Balassa-Samuelson model. The empirical tests take into account the long-run nature of the Balassa-Samuelson model by extracting low frequency components from time series for 14 OECD countries with the Hodrick-Prescott (1980) filter. The empirical tests also exploit the panel structure of the data.

The empirical evidence we provide suggests that low frequency differences in relative labor productivities do explain differences in long-run relative prices in our sample of OECD countries. We conclude that the first proposition of the Balassa-Samuelson model is consistent with the long-run implications of the balanced-growth general equilibrium model developed in this paper. We then follow Balassa (1964) and examine the extent to which the theory can explain low frequency deviations from PPP observed in the data. The results suggest that while relative labor

productivity differentials do explain the long-run behavior of the domestic relative price of nontradables, the relative price of nontradables is far less successful in explaining observed cross-country differences in long-run CPI-based and GDP deflator-based real exchange rates. In our equilibrium model this negative result can be attributed to the failure of PPP in tradable goods; or to a rejection of either the constant-elasticity forms of the production and utility functions or the balanced-growth constraints.

As a by-product of our analysis we are able to clarify two theoretical results that are important in assessing the findings of some empirical studies of the Balassa-Samuelson model. First, the proposition that sectoral labor productivity differentials are the only determinants of equilibrium domestic relative prices is, in general, only a *long-run* implication of neoclassical models. We show that in the *short-run*, the ratio of marginal products of labor determines only the supply of nontradable goods relative to tradable goods. Demand is determined by the households' marginal rate of substitution between the two goods. Thus, the short-run determination of the equilibrium relative price of nontradables cannot be studied without modeling the households' optimization problem. This result casts doubt on empirical studies of the Balassa-Samuelson model that consider only the supply-side and *time series* properties of the relative price of nontradables, without distinguishing between the long- and short-run components of the data.

Second, a key finding of the original Balassa paper is that there is a positive relationship between *aggregate* output per head and the real exchange rate (or the relative price of nontradables). However, the theoretical analysis shows that in the long-run, it is the ratio of *sectoral* marginal products of labor that determines the relative price of nontradables. Therefore, the original Balassa-Samuelson model cannot predict how *aggregate* output per-capita should relate to domestic relative prices. This holds even if it is assumed that sectoral technologies are such that average and marginal products are proportional to each other and that population is a good proxy for labor services or hours worked. We conclude that, although the observed positive relationship between aggregate output per head and the real exchange rate (or the relative price of nontradables) remains an important stylized fact, it cannot be easily derived from the theoretical principles underlying Balassa and Samuelson's original formulation.

The paper is organized as follows. In Section II we outline the theoretical framework and establish the Balassa-Samuelson propositions as steady-state implications of a standard dynamic neoclassical model. In Section III we discuss data analysis and filtering issues. In Section IV we provide the empirical results. Section V presents some concluding remarks.

II. The Theoretical Framework

In this section we describe the structure of our two-country, two-sector, intertemporal general equilibrium model. The model we examine is similar to that developed by Stockman and Tesar (1990), but differs in that our analysis focuses on the long-run rather than on business cycle frequencies. The conditions we derive for the long-run behavior of the relative price of nontradables are robust to alternative specifications within the class of multisector intertemporal equilibrium models of the open economy. In particular, our results hold for models with or without complete contingent claims markets and with or without distortionary taxes (see Mendoza and Tesar (1993)).

Consider a two-country world economy where households in each country consume tradable and nontradable goods and supply labor services to firms producing those goods. Households formulate optimal intertemporal plans to maximize expected lifetime utility. Firms produce tradable and nontradable goods by hiring the services of labor and capital and by combining them according to Cobb-Douglas technologies subject to stationary productivity disturbances. Households and firms are free to trade goods, equity, and financial assets internationally. For notational clarity we only describe the characteristics of preferences and production in the home country. Foreign country characteristics are symmetric and, where necessary, identified by an asterisk.

1. Firms

Firms in the home country produce two types of goods tradables (T) and nontradables (NT) according to the following constant returns to scale Cobb-Douglas technologies:

$$Y_t^T = F(K_t^T, N_t^T) = A_t^T (X_t N_t^T)^{\alpha T} (K_t^T)^{1-\alpha T} \qquad 0 \leq \alpha^T \leq 1 \tag{1}$$

$$Y_t^{NT} = F(K_t^{NT}, N_t^{NT}) = A_t^{NT} (X_t N_t^{NT})^{\alpha NT} (K_t^{NT})^{1-\alpha NT} \qquad 0 \leq \alpha^{NT} \leq 1 \tag{2}$$

where the production function, F(.), in each sector is assumed to be concave, increasing and twice continuously differentiable. Y^i_t, i = T, NT is the output of tradable and nontradable goods at time t respectively; K^i_t, i = T, NT are the stocks of physical capital allocated to the production of tradable and nontradable goods at time t. Factors of production are assumed to be perfectly mobile across tradable and nontradable sectors and capital may be owned by households in either country. N^i_t, i = T, NT represents labor inputs required for the production of each good at time t, X_t is an index of Harrod-neutral labor-augmenting technological progress at time t and A^i_t, i = T, NT, are stochastic

productivity disturbances. 1/ Total factor productivity in each sector is given by:

$$\theta_t^T = A_t^T (X_t)^{\alpha T} \tag{3}$$

$$\theta_t^{NT} = A_t^{NT} (X_t)^{\alpha NT} \tag{4}$$

The stationary productivity shocks induce fluctuations of macroeconomic variables around long-run deterministic trends. 2/ These long-run trends are identified by imposing the balanced-growth conditions discussed in King, Plosser, and Rebelo (1988), where growth is driven by exogenous, labor-augmenting technological progress as in (1) and (2). Technological change evolves over time at the rate γ (where γ is the rate of growth of labor-augmenting technological change, i.e., the aggregate growth rate). For conventional preferences and technology this results in balanced-growth for all components of aggregate demand. Moreover, from (3) and (4) it follows that the differential in total factor productivity growth, which has played a key role in previous studies of the Balassa-Samuelson model, is:

$$\ln\left(\frac{\theta_{t+1}^T}{\theta_t^T}\right) - \ln\left(\frac{\theta_{t+1}^{NT}}{\theta_t^{NT}}\right) = (\alpha T - \alpha NT)\ln\gamma + \epsilon_{t+1} \tag{5}$$

where: $\gamma = \dfrac{X_{t+1}^T}{X_t^T} = \dfrac{X_{t+1}^{NT}}{X_t^T}$ and $\epsilon_{t+1} = \ln\left(\dfrac{A_{t+1}^T}{A_t^T}\right) - \ln\left(\dfrac{A_{t+1}^{NT}}{A_t^{NT}}\right)$

and ϵ is a stationary random process. Thus, for a given rate of balanced growth (γ), the differential in total factor productivity growth is determined by the difference in labor income shares.

It is well known that with labor-augmenting technological progress the model exhibits steady-state growth. Therefore, a transformation is required to render the representative household's optimization problem stationary.

1/ See Swan (1963) and Phelps (1966) who show that the assumption of labor-augmenting technological progress is a necessary condition for steady-state growth in neoclassical growth models.

2/ Obstfeld (1993) notes that this is a reasonable approximation for industrial country multilateral real exchange rates.

This transformation is achieved by deflating all variables (except labor and leisure) by the index of technological progress X_t. 1/

The first order conditions for the firm's optimization problem, given the rental rate for capital r_t and the wage rate for labor w_t in each sector, yield the following zero-profit conditions:

$$f(k_t^T,N_t^T)=r_t^T k_t^t+w_t^T N_t^t \tag{6}$$

$$f(k_t^{NT},N_t^{NT})=r_t^{NT} k_t^{NT}+w_t^{NT} N_t^{NT} \tag{7}$$

where f(.) and k^i_t, i = T, NT represent the transformed (detrended) production functions and capital stock, respectively. r^i_t, i = T, NT are the rental rates for capital in the tradable and nontradable sectors at time t and w^i_t, i = T, NT are real wages in each sector at time t.

2. Households

The economy is inhabited by an infinitely lived representative household with a time separable utility function defined over the consumption of tradables, nontradables and leisure. The household maximizes the discounted sum of expected lifetime utility.

$$E\left[\sum_{t=0}^{\infty} \beta^t U(c_t^T,c_t^{NT},L_t)\right] \quad 0<\beta<1 \tag{8}$$

where E is the expectations operator conditioned on the time t information set; β is the subjective discount factor; c^T_t and c^{NT}_t are the consumption of tradables and nontradables at time t respectively and L_t is the time devoted to leisure. The instantaneous utility function is twice continuously differentiable in each of its arguments.

We assume a constant elasticity of substitution (CES) instantaneous utility function:

$$U(.)=\frac{\left[\left[\Omega(c_t^T)^{-\mu}+(1-\Omega)(c_t^{NT})^{-\mu}\right]^{\frac{-1}{\mu}} L_t^w\right]^{1-\sigma}}{1-\sigma} \tag{9}$$

1/ The discount factor and law of motion for capital are also properly adjusted. All deflated variables are written in lower case caps.

where Ω is the share of tradables in consumption, $1/1+\mu$ is the elasticity of substitution between tradable and nontradables, ω is the elasticity of leisure, and σ is the coefficient of relative risk aversion.

Households maximize utility subject to the budget constraint:

$$p^{NT}_t c^{NT}_t + c^T_t = \left[r^T_t k^H_t + r^{T*}_t k^F_t + p^{NT}_t r^{NT}_t k^{NT}_t\right] + \left[w^T_t N^T_t + p^{NT}_t w^{NT}_t N^{NT}_t\right]$$
$$\gamma\left[k^H_{t+1} + k^F_{t+1} + p^{NT}_t k^{NT}\right] + (1-\delta)\left[k^H_t + k^F_t + p^{NT}_t k^{NT}_t\right] \qquad (10)$$
$$-\gamma R_t b_{t+1} + b$$

and the normalized time constraint:

$$L_t + N^{NT}_t + N^T_t = 1 \qquad (11)$$

where p^{NT}_t is the relative price of nontradables, k^H_t, k^F_t and k^{NT}_t are the stocks of physical capital owned by households in the home country in the domestic tradables sector, the foreign tradables sector and the domestic nontradables sector respectively. Capital in both sectors is assumed to depreciate at the same rate δ.

Households accumulate net foreign assets, b, that yield the world interest rate i_t. R is the inverse of the real gross rate of return paid on international bonds. Thus, we assume a financial market structure in which countries trade only equity and noncontingent bonds and therefore, insurance markets are incomplete. The household's problem, therefore incorporates the period-by-period constraint (10) instead of the present value of wealth typical of complete market models. 1/

For the transformation procedure discussed earlier to produce stationary equilibrium allocations that correspond to nonstationary, balanced-growth equilibrium allocations, two additional adjustments are required. First, the discount factor must be transformed so that $\tilde{\beta} = \beta \, .\gamma^{1-\sigma}$ where $\beta = 1/1 + \rho$ is the rate of time preference. Second, it is required that γ be introduced as a multiplicative factor in the accumulation of capital and bonds in the budget constraint.

1/ See Cole (1988) for a discussion of this issue. Our results still hold in a model like that of Stockman and Tesar (1990) where markets are complete.

3. Competitive equilibrium

In a competitive equilibrium for this world economy, home and foreign households maximize utility, home and foreign firms maximize profits and the goods, services, and financial asset markets clear. The equilibrium is characterized by allocations of consumption, labor supply, capital and international bonds that satisfy the following optimality conditions in the home country:

$$\frac{U_1(t)}{U_2(t)} = p_t^{NT} \tag{12}$$

$$\frac{U_3(t)}{U_1(t)} = w_t^{T} \tag{13}$$

$$\frac{U_3(t)}{U_2(t)} = w_t^{NT} \tag{14}$$

$$\gamma R_t U_1(t) = \beta E[U_1(t+1)] \tag{15}$$

$$\gamma U_1(t) = \beta E\left[U_1(t+1)\left[r_{t+1}^{T}+1-\delta\right]\right] \tag{16}$$

$$\gamma U_1(t) = \beta E\left[U_1(t+1)\left[r_{t+1}^{T*}+1-\delta\right]\right] \tag{17}$$

$$\gamma p_t^{NT} U_1(t) = \beta E\left[p_{t+1}^{NT} U_1(t+1)\left[r_{t+1}^{NT}+1-\delta\right]\right] \tag{18}$$

$$w_t^{T} = f_2(k_t^{T}, N_t^{T}) \tag{20}$$

$$r_t^{NT} = f_1(k_t^{NT}, N_t^{NT}) \tag{21}$$

$$w_t^{NT} = f_2(k_t^{T}, N_t^{T}) \tag{22}$$

The market-clearing conditions are:

$$f(k_t^{NT},N_t^{NT}) = c_t^{NT}+\gamma k_{t+1}^{NT}-(1-\delta)k_T^{NT} \tag{23}$$

$$f(k_t^{NT*},N_t^{NT*}) = c_t^{NT*}+\gamma^* k_{t+1}^{NT}-(1-\delta)k_t^{NT*} \tag{24}$$

$$f(k_t^T,N_t^T)+f(k_t^{T*},N_t^{T*}) = c_t^T+c_t^{T*}+\gamma k_{t+1}^T-(1-\delta)k_t^T+\gamma^* k_{t+1}^T-(1-\delta)k_t^{T*} \tag{25}$$

$$b_t + b_t^* = 0 \tag{26}$$

U_i, i=1,2,3 are the partial derivatives with respect to the first (c^T), second, (c^{NT}) or third (L)arguments of the utility function. The corresponding conditions in the foreign country and the budget constraints are also part of the set of optimality conditions describing world equilibrium. Conditions (12)-(22) have the usual interpretation in terms of marginal productivities and rental prices of inputs.

Of considerable importance in our analysis of the Balassa-Samuelson model are equations (12)-(14) and (18)-(22), that determine the equilibrium relative price of nontradables. Equation (12) states that from the demand-side, the equilibrium relative price of nontradables at time t is equal to the marginal rate of substitution between tradable and nontradable goods. By dividing (14) by (13), substituting the result in (12), and displacing the rental prices of labor with the marginal products as stated in (20) and (22), one can also show that from the supply-side the equilibrium relative price of nontradables at time t is the ratio of the marginal products of labor in the tradable and nontradable sectors.

This static characterization of the relative price of nontradables in terms of the ratio of the marginal products of labor is the principle emphasized by Balassa and Samuelson. However, in world general equilibrium both demand- and supply-side conditions must be satisfied by the market-clearing relative price of nontradables. Moreover, these two conditions are not independent of the rest of the equilibrium system. In deterministic form, (18) is an Euler condition linking the intertemporal marginal rate of substitution to the change in the relative price of nontradables over time. This Euler condition introduces intertemporal income and substitution effects in the determination of the relative price of nontradables at date t. This means that optimal intertemporal plans concerning consumption and investment affect atemporal decisions regarding allocations of consumption across tradables and nontradables and of capital and labor across sectors; hence affecting the relative price of nontradables.

4. The long-run price of nontradables

In general, as the above discussion showed, the original Balassa-Samuelson principle is only a characterization of supply-side determinants of the relative price of nontradables. In this section we show that the Balassa-Samuelson principle can be interpreted as an *equilibrium* outcome along the long-run balanced-growth path.

To establish the Balassa-Samuelson principle as a long-run equilibrium outcome we proceed by assuming the random shocks to the production technologies are stationary and that certainty equivalence holds. This, enables us to examine the long-run balanced growth world equilibrium by focusing on the model's deterministic stationary state. In this steady-state, the equilibrium relative price of nontradables reduces to expressions closely related to the Balassa-Samuelson framework.

Consider the supply-side equilibrium condition that equates the relative price of nontradables to the ratio of the marginal products of labor in the tradable and nontradable sectors, within a country:

$$p^{NT} = \frac{f_2(k_t^T, N_t^T)}{f_2(k_t^{NT}, N_t^{NT})}$$

Exploiting the fact that Cobb-Douglas production functions have the property that output per man-hour is a monotonic transformation of the capital-output ratio, $(y/N)=(k/y)^{(1-\alpha)/\alpha}$, enables us to write the relative price of nontradables as:

$$p^{NT} = \left(\frac{\alpha T}{\alpha NT}\right)\left[\frac{\left(\frac{k^T}{y^T}\right)^{\frac{1-\alpha T}{\alpha T}}}{\left(\frac{k^{NT}}{y^{NT}}\right)^{\frac{1-\alpha NT}{\alpha NT}}}\right] \tag{27}$$

Thus, (27) is a supply-side condition that states that the relative price of nontradables is a function of sectoral labor shares and sectoral capital-output ratios. Note that the relative price of nontradables is higher the higher is output per man-hour in the tradable goods sector *relative* to the nontradable goods sector. Therefore the theory, as developed here, cannot predict how *aggregate output per capita* relates to domestic relative

prices. 1/ Even if it is assumed that technology is such that average and marginal products are proportional to each other, as in the Cobb-Douglas case, and that population is a good proxy for labor services or hours worked, it is the ratio of sectoral output-per capita levels that determines the relative price of nontradables and not the aggregate level of output.

From (16) and (18) it follows that in a deterministic stationary equilibrium with perfect sectoral capital mobility, the marginal products of capital in the tradable and nontradable sectors are equalized:

$$f_1\left(k_t^T, N_t^T\right) = f_1\left(k_t^{NT}, N_t^{NT}\right)$$

With Cobb-Douglas production functions this relationship reduces to:

$$\frac{k_t^{NT}}{y_t^{NT}} = \left(\frac{1-\alpha NT}{1-\alpha T}\right) \frac{k_t^T}{y_t^T}$$

Equation (27) can therefore be rewritten to express the relative price of nontradables as a function of the labor shares in both sectors and the capital-output ratio in the tradables sector:

$$p^{NT} = \left(\frac{\alpha T}{\alpha NT}\right) \left(\frac{1-\alpha NT}{1-\alpha T}\right)^{\frac{\alpha NT-1}{\alpha NT}} \left(\frac{k_t^T}{y_t^T}\right)^{\frac{1-\alpha T}{\alpha T} - \frac{1-\alpha NT}{\alpha NT}} \tag{28}$$

Up to this point, we have derived expressions for the relative price of nontradables that depend on capital-output ratios and represent either the supply-side condition (27) or that condition jointly with the steady-state equality of sectoral marginal products of capital (28). To argue that these conditions explain equilibrium allocations along the balanced-growth path, we need to establish that capital-output ratios are exogenously determined by structural parameters. We do this by imposing steady-state conditions on all of the equations (12)-(22). After manipulation of (16), in long-run balanced-growth equilibrium the capital-output ratio in the tradables sector is:

1/ One reason for this is that the theory precludes by assumption the potential supply-side relationship between aggregate output per capita and relative price of nontradables due to nonhomothetic tastes, see Bergstrand (1991) and De Gregorio, Giovannini and Wolf (1994).

$$\frac{k_t^T}{y_t^T} = \frac{\tilde{\beta}(1-\alpha T)}{\gamma-\tilde{\beta}(1-\delta)} \tag{29}$$

This equation incorporates the steady-state equality of the intertemporal marginal rate of substitution in consumption and the real rate of return on capital (net of depreciation) required to produce balanced-growth at the rate γ in the components of aggregate demand.

What emerges from the analysis, at this point, is that in long-run growth equilibrium the capital-output ratio in the tradables sector is determined by exogenous structural parameters, β, γ, σ, αT, δ. Therefore, at low frequencies (27) and (28) can be interpreted as expressions that determine the *equilibrium* relative price of nontradables and not simply the supply-side of the economy.

Working with (29) and the steady state definition of the investment rate, $\frac{i^T}{y^T} = [\gamma-(1-\delta)]\frac{k^T}{y^T}$, we find an alternative representation of the equilibrium relative price of nontradables as a function of the investment output ratio:

$$p^{NT} = \left(\frac{\alpha T}{\alpha NT}\right)\left(\frac{1-\alpha NT}{1-\alpha T}\right)^{\frac{\alpha NT-1}{\alpha NT}}\left[\frac{i^T}{y^T}[\gamma-(1-\delta)]^{-1}\right]^{\frac{1-\alpha T}{\alpha T}-\frac{1-\alpha NT}{\alpha NT}} \tag{30}$$

or as a function of deep structural parameters:

$$p^{NT} = \left(\frac{\alpha T}{\alpha NT}\right)\left(\frac{1-\alpha NT}{1-\alpha T}\right)^{\frac{\alpha NT-1}{\alpha NT}}\left[\frac{\gamma^{-\sigma}\beta(1-\alpha T)}{1-\gamma^{-\sigma}\beta(1-\delta)}\right]^{\frac{1-\alpha T}{\alpha T}-\frac{1-\alpha NT}{\alpha NT}} \tag{31}$$

Finally, note that the expressions we have derived for the equilibrium relative price of nontradables in (27), (28) and (30) are consistent with those from earlier studies of the Balassa-Samuelson model that emphasize sectoral differentials in factor productivity growth. 1/ This is evident from the fact that in our model, given capital-output or investment-output ratios, the relative price of nontradables is determined by the relative size of α^{NT} and α^{T}. These two parameters in turn determine the differential in sectoral total factor productivity growth given in (5).

1/ See De Gregorio, Giovannini and Wolf (1994), and Kravis Heston and Summers (1983).

5. The long-run real exchange rate

In this subsection we link real exchange rates to the equilibrium relative price of nontradables. We establish the connection between the model's equilibrium relative price of nontradables and the real exchange rate by following the convention of the intertemporal equilibrium literature. 1/ The convention is to proceed by noting that the households problem has a dual representation with an expenditure function P_tC_t where C_t is a composite consumption good represented by $C_t = [\Omega(c^T{}_t)^{-\mu}+(1-\Omega)(c^{NT}{}_t)^{-\mu}]^{-1/\mu}$, and P_t is the price index of the composite consumption good represented as:

$$P_t(P^T_t, P^{NT}_t) = \left[\Omega^{\frac{1}{1+\mu}} P^{T\,\frac{\mu}{1+\mu}}_t + (1-\Omega)^{\frac{1}{1+\mu}} P^{NT\,\frac{\mu}{1+\mu}}_t\right]^{\frac{1+\mu}{\mu}}$$

Define the real exchange rate as $s_t = P^*{}_t/P_t$. 2/ Then, if the law of one price holds for tradable goods, the real exchange rate is:

$$s_t = \frac{\left[\Omega^{*\frac{1}{1+\mu*}} + (1-\Omega^*)^{\frac{1}{1+\mu*}} P^{NT*\,\frac{\mu*}{1+\mu}}_t\right]^{\frac{1+\mu*}{\mu*}}}{\left[\Omega^{\frac{1}{1+\mu}} + (1-\Omega)^{\frac{1}{1+\mu}} P^{NT\,\frac{\mu}{1+\mu}}_t\right]^{\frac{1+\mu}{\mu}}}$$

From this expression it is evident that the real exchange rate is a function of the relative price of nontradables in the two countries. In long-run, balanced-growth equilibrium the real exchange rate is therefore a function of the same structural parameters that determine the ratio of sectoral marginal products of labor, which as we showed earlier determine the relative price of nontradables.

Assuming Cobb-Douglas preferences, i.e., $(1/1+\mu=1)$, enables us to conveniently express the real exchange rate for empirical implementation as:

$$s_t = \left[\frac{(\Omega^*)^{\Omega^*}(1-\Omega^*)^{1-\Omega^*}}{\Omega^{\Omega}(1-\Omega)^{1-\Omega}}\right]\left[\frac{P^{NT*(1-\Omega^*)}_t}{P^{NT(1-\Omega)}_t}\right] \qquad (32)$$

1/ See Frenkel and Razin (1987), Backus and Smith (1993), Greenwood (1984) and Mendoza (1992).

2/ The convention at the International Monetary Fund is to define the real exchange rate as $P_t/P^*{}_t$. This should be kept in mind for the empirical analysis.

III. Data Analysis and Filtering

Estimating (27), (28), (30) and (32) requires data on the relative price of nontradables, the investment-output ratio in the tradable sector and the capital-output ratios in the tradable and nontradable sectors. These variables do not exist in ready form, so the first task was to construct these variables from existing sources.

As our focus is on the cross-country properties of the data, we constructed a panel dataset. The dataset provides a rich source of cross-country information and consists of annual data for 14 countries, 20 sectors spanning 1970-85 and was obtained from the OECD intersectoral database. 1/2/ The database includes information on sectoral real and nominal valued added, capital stocks, investment, employment and factor returns for each of the 20 sectors. From this database we constructed series for the relative price of nontradables, the investment-output ratio in the tradables sector and the capital-output ratio in tradables and nontradables sectors for each country in our sample.

In order to construct the required data, the first issue is to decide which sectors are to be considered tradable and nontradable. We choose De Gregorio, Giovannini and Wolf's (1994) classification scheme. This scheme is based on the ratio of the actual shares of total exports to total production across all 14 countries for each sector. This results in a sector being classified as tradable if more than 10 percent of total production is exported. 3/ The 10 percent threshold classifies agriculture, mining, all of manufacturing and transportation as tradables with the remaining sectors classified as nontradables. Annual data on real exchange rates based on trade weighted consumer price indices (CPI) were obtained from the IMF International Financial Statistics while GDP deflator-based real exchange rates were taken from Micosi and Milesi (1993).

We decided to extract the long-run growth component of the data before estimation for the following two reasons. First, we have shown that the Balassa-Samuelson predictions are long-run equilibrium implications. Thus,

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1/ The countries studied are: Australia, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, the Netherlands, Norway, Sweden, the United Kingdom and the United States.

2/ The 20 sectors are: (1) Agriculture (2) Mining (3) food, beverages and tobacco, (4) textiles (5) wood and wood products (6) paper, printing and publishing (7) chemical (8) nonmetallic mineral products (9) basic metal products (10) machinery equipment (11) other manufactured products (12) electricity, gas and water (13) construction (14) wholesale and retail trade (15) restaurants, hotels (16) transport, storage and communications (17) finance, insurance (18) real estate (19) community, social and personal services (20) government services.

3/ For details see De Gregorio, Giovannini and Wolf (1994). Their classification is similar to that of Stockman and Tesar (1990).

to be consistent with the theory, any tests of the predictions of our model must be based on the long-run components of the data. Second, it is well known that employment adjusts gradually to changes in output and as a result labor productivity rises in an economic upturn and declines in a downturn. By extracting the growth component from the data, we isolate the factors that are more closely related to long-run labor productivity and abstract from short-run cyclical changes that may bias the results.

Several statistical procedures have been used to filter data in macroeconomic analysis. The most common ones are the linear-trend filter, the Hodrick-Prescott (HP) filter, the Beveridge-Nelson filter and random-walk detrending (Canova and Dellas 1993). Unfortunately, a consensus on the appropriate use of filters in macroeconomic analysis does not exist. However, Baxter (1991) and Singleton (1988) have argued that the choice of filtering procedure should be governed by the theoretical model at hand. We find their arguments compelling and choose two filters: the linear-trend and HP filters because they are consistent with our version of the Balassa-Samuelson model (i.e, deterministically trending variables uncorrelated with the cyclical components of the data).

The linear-trend filter removes a deterministic linear trend from the data and is attractive for its simplicity. However, the simplicity of the linear-trend filter presents a drawback when applied to highly nonstationary processes such as exchange rates and relative prices. To confirm that the data does exhibit nonstationarity, we carried out Dickey-Fuller and Augmented Dickey-Fuller stationarity tests. As expected, the tests fail to reject the presence of unit roots in all three of the data series.

The HP-filter has certain attractions relative to the linear-trend filter. Like the linear-trend filter, the HP-filter assumes that the cyclical and growth components of the data are uncorrelated. However, unlike the linear-trend filter, the HP-filter will render stationary any integrated process up to fourth order (King and Rebelo 1993). Furthermore, the HP-filter permits the data generating process to have a deterministic as well as a stochastic growth component.

Figure 1 plots the actual observations and the HP-filtered trends of the relative price of nontradables, the investment-output and capital-output ratios in tradables and the capital-output ratios in nontradables for Germany. Visual examination of Figure 1 suggests that linear trends are not likely to differ significantly from the HP-filtered trends. We confirmed this by plotting both filters. While the two filtering procedures are remarkably similar for some variables like the investment-output ratio the HP-filter captures a slow moving trend that the linear trend filter misses. Given these results, we decided to use the HP-filter in the empirical analysis reported in the remainder of the paper. 1/

1/ Plots of the HP-filter and actual data for the other countries are similar and not reported here to conserve space.

A striking feature that is evident from Figure 1 is the smoothness of the trend component that emerges from the HP-filtering procedure. Harvey and Jaeger (1993) argue that to avoid blind application of the HP-filter, the assumption of a smooth deterministic trend should be empirically verified by estimating a structural time series model: 1/

$$y_t = \mu_t + \Gamma_t + \epsilon_t \qquad t=1\ldots T$$

where y_t is the series; μ_t is the trend; Γ_t is the cycle and ϵ_t is a random error term. The trend is:

$$u_t = \mu_{t-1} + \beta_{t-1} + \eta \qquad \eta_t \sim N(0, \sigma^2_\eta)$$

$$\beta_t = \beta_{t-1} + \xi_t \qquad \xi_t \sim N(0, \sigma^2_\xi)$$

where β_t is the slope parameter and ξ_t and η_t are independent and normally distributed white noise.

The cyclical term is stochastic and assumed to be generated by

$$\Gamma_t = \rho \cos \lambda_c \Gamma_{t-1} + \rho \sin \lambda_c \Gamma^*_{t-1} + x_t$$

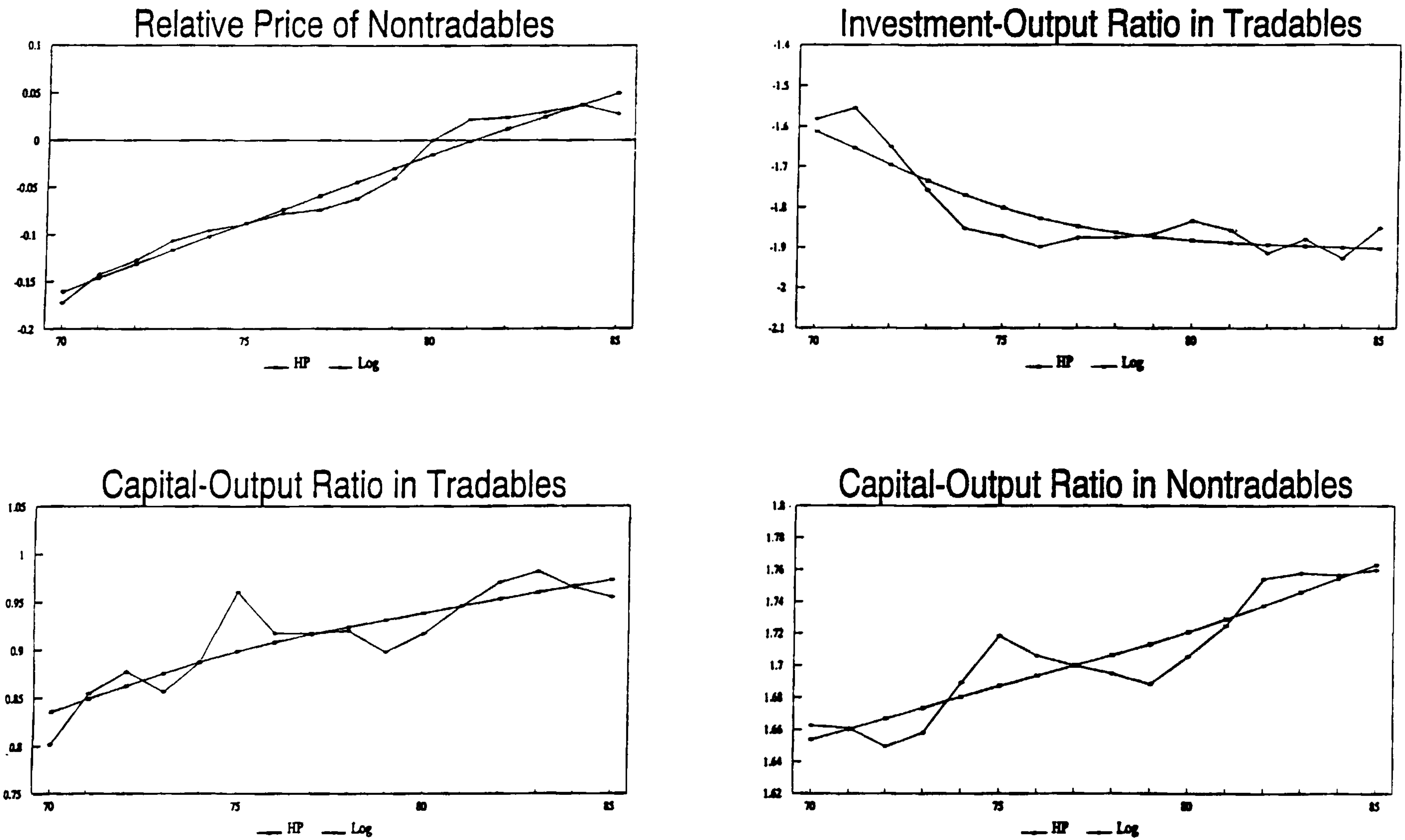
$$\Gamma^*_t = -\rho \sin \lambda_c \Gamma_{t-1} + \rho \cos \lambda_c \Gamma^*_{t-1} + x^*_t$$

where ρ is a dampening factor such that $0 \leq \rho \leq 1$, λ_c is the frequency of the cycle, and x_t and x^*_t are both normal and identically distributed disturbances with mean zero and variance σ^2_x. The random error term is also normal and identically distributed with mean zero and variance σ^2_e and all three components are assumed to be independent of each other.

If $\sigma^2_\xi=0$ the trend reduces to a random walk with drift. Furthermore if $\sigma^2_\eta=0$, the trend becomes deterministic, that is $u_t=\beta_t$. When $\sigma^2_\eta=0$, but $\sigma^2_\xi>0$ the trend component is relatively smooth. Therefore, whether the trend component is deterministic and well represented by a smooth process can be verified by testing whether $\sigma^2_\eta=0$. We carried out maximum likelihood estimation of the parameters of the structural model for each of the 14 countries for 4 variables (the real exchange rate, the relative price of nontradables, the investment-output ratio in tradables and the capital-output ratio in nontradables) to determine whether this restriction was supported by the data. The results indicate that the deterministic smooth trend assumption is supported by the data for 10 of the 14 countries for all

1/ The following discussion draws heavily on Harvey and Jaeger (1993).

Figure 1. Comparison of HP Filter and Raw (log)
Data for Germany



4 variables since $\tilde{\sigma}_\eta$=0. 1/ The remaining 4 countries had values of $\tilde{\sigma}_\eta$ that were small ranging from 1-4 but with values of $\tilde{\sigma}_e$=0. 2/ The fact that $\tilde{\sigma}_e$=0 suggests that even for these 4 countries the series decomposes into a smooth trend and cycle. Finally, plots of the trend component from estimates of the structural model for the 4 countries suggest that trends from the structural model have similar features to those from the HP filter. These results are consistent with Obstfeld (1993) who provides evidence of deterministic trends in real exchange rates for the United States and Japan. 3/

IV. Empirical Results

The empirical analysis is structured around two questions. First, do long-run relative labor productivities explain long-run relative nontradable prices? Addressing this question will enable us to evaluate the Balassa-Samuelson model as a theory of the determination of domestic relative prices. Second, do cross-country differences in long-run relative nontradable prices explain cross-country, long-run, real exchange rate differentials? Addressing the second question enables us to determine the extent to which the Balassa-Samuelson framework can be considered a theory of real exchange rates.

1. Evidence on the long-run relative price of nontradables

Having derived closed-form solutions for the long-run relative price of nontradables in Section II, our empirical strategy is to confront the theory with the data in the most parsimonious manner possible. In reassessing the Balassa-Samuelson model, we therefore purposefully refrain from adding additional right-hand-side variables not derived from the model to the regressions. The tests we carry out are joint tests of the theory and the assumption of Cobb-Douglas technologies.

The log-linear form of the nontradable price equations for country j derived in (27), (28) and (30) can be conveniently summarized for estimation as:

1/ Australia, Belgium, Canada, Germany, Italy, Japan, the Netherlands, Sweden, United Kingdom, United States.

2/ Denmark, Finland, France and Norway.

3/ These results are not reported here to conserve space (they would require 14 separate tables). The results are available on request from the authors.

$$p_{jt}^{NT} = \alpha_{0j} + \alpha_1 ky_{jt}^{T} + \alpha_2 ky_{jt}^{NT} + e_{jt} \qquad (I)$$

$$p_{jt}^{NT} = \gamma 0_j + \gamma_1 ky_{jt}^{T} + e_{jt} \qquad (II)$$

$$p_{jt}^{NT} = \eta 0_j + \eta 1\, iy_{jt}^{NT} + e_{jt} \qquad (III)$$

for j=1, 2, ..., M countries and t=1, 2, ..., T time periods, where p^{NT} is the log of the relative price of nontradables; ky^T is the log of the capital-output ratio in tradables; ky^{NT} is the log of the capital-output ratio in nontradables; iy^{NT} is the log of the investment-output ratio in tradables and e_{jt} are random disturbances. For easy reference these three specifications will henceforth be referred to as specification (I), (II) and (III) respectively.

The theory requires the coefficient on the capital-output ratio (α_2) in nontradables to be negative and the coefficient on the capital-output ratio (α_1) in the tradables sector to be positive in (I). With respect to (II) and (III) the theory does not impose constraints on the coefficient on the capital-output ratio in tradables (γ_1) or on the coefficient on the investment-output ratio in tradables (η_1). However, if $\alpha^T > \alpha^{NT}$, as data on labor income shares suggests, then both γ_1 and η_1 should be negative. 1/ Moreover, the model also implies that the cross equation restrictions $\gamma_1 = \eta_1 = \alpha_1 + \alpha_2$ should hold.

Table 1 provides least squares estimates of a pooled (total) regression of equations (I), (II) and (III). Equation (I) performs particularly well in several respects. First, the coefficients are statistically significant and of the correct sign. Second, $\alpha^T > \alpha^{NT}$ is implicit in the results although the implied shares $\alpha^T = 0.81$ and $\alpha^{AT} = 0.78$ are higher than direct measures suggest. Finally, equation (I) explains nearly one quarter of the variations in the relative price of nontradables.

In contrast, the results from estimating (II) and (III) are less favorable. The coefficient estimates of γ_1 and η_1 are not statistically different from zero and the explanatory power of the regressions is very low. However, the t-ratios for the null hypotheses that γ_1 and η_1 are not different from $\alpha_1 + \alpha_2 = -0.038$ are 3.2 and 4.1, respectively. Thus, although the data do not provide precise estimates of γ_1 and η_1, the cross equation restrictions $\gamma_1 = \eta_1 = \alpha_1 + \alpha_2$ cannot be rejected.

A possible reason for the poor performance of the pooled regressions (II, III) is that in performing least squares regressions with all M T

1/ See Kravis, Heston and Summers (1983) and Stockman and Tesar (1990). The latter noted that the labor share of tradable goods was greater than that for nontradables for 5 of 7 countries in their sample.

observations we have assumed that the intercept and slope coefficients take values common to all cross-sectional units. If this assumption is not valid, the pooled least squares estimates may lead to false inferences. To investigate whether the regression coefficients are the same for all countries we carry out several homogeneity tests. Our strategy is to determine whether the slopes and intercepts simultaneously are homogeneous among different countries at different times. Then we test if the regression slopes are collectively the same. We construct F tests of the above restrictions under the assumption that the errors e_{jt} are independently normally distributed over j and t with mean zero and variance σ^2_u.

Table 2 presents the results of tests for the homogeneity of regression slope coefficients and homogeneity of the regression intercept coefficients. In hypothesis 1 (same slopes, same intercepts) the F ratio is significant so we reject the hypothesis of complete homogeneity. Hypothesis 2 (same slopes but different intercepts) is also rejected, suggesting that the slope coefficients are also different across countries. We interpret the failure of these tests as suggesting that sectoral labor shares, which are the determinants of intercept and slope estimates in (I), (II) and (III), differ across countries or groups of countries. Later, we show how estimation performance improves if we group countries according to relative labor shares implicit in the intercept estimates.

Next, we decompose the pooled regression estimates into "within" and "between" components for two partitions of the sample. Panel A in Table 3 is for the full sample while panel B is based on subsamples for 1970-77 and 1978-85. The between component represents the output of an OLS regression based on the means of each country's time series, while the within component is the outcome of a fixed-effects model. By proceeding in this manner we can determine the contribution of each of the two components to the outcome of the total pooled regression.

The results of the decomposition are reported in Table 3. The weights (κ) on the between estimates indicates that almost 90 percent of the variation in the pooled estimates is due to heterogeneity across countries. Thus the favorable results obtained with the pooled regressions reported in Table 1, particularly for equation (I), can be viewed as reflecting mainly cross-country differences in trend behavior, rather than within country time series patterns. This result is robust to the specification of two subsamples. Moreover, coefficient estimates are generally stable for the sample break down examined.

To explain the differences in performance between (I), (II) and (III) recall that in deriving (II) and (III) we imposed the equilibrium condition that equates the marginal products of capital in the tradable and nontradable sectors. We also simplified this equality with the conditions required for balanced-growth in the model. Particularly the assumption that the domestic relative price of nontradables is constant in the long-run (at levels that differ across countries depending on total factor productivity growth). Therefore, our results may reflect the fact that

Table 1. Pooled (Total) Regression of Nontradable Price on the Investment-Output and Capital-Output Ratios

Variable	Estimated Coefficients (t-ratio) Equation (I)	Equation (II)	Equation (III)
ky^T	0.240 2/	0.075	--
	(4.7)	(1.3)	--
ky^{NT}	-0.278 2/	--	--
	(-7.9)	--	--
iy^T	--	--	0.009
	--	--	(0.8)
Intercept	0.149 2/	-0.048	0.059 1/
	(2.6)	(-0.8)	(1.7)
Adjusted R^2	0.225	0.003	-0.002
F-statistic	34.763	1.750	0.599
Log-likelihood	75.467	46.961	46.384

Notes: ky^T is the capital-output ratio in the tradable sector. ky^{NT} is the capital-output ratio in the nontradable sector. iy^T is the investment-output ratio in tradable sector.

1/ Statistically significant at the 10 percent level.
2/ Statistically significant at the 5 percent level.

Table 2. Covariance Tests for Homogeneity

	Equation (I)	Equation (II)	Equation (III)
Residual sum of squares under			
Hypothesis 1	13.212	8.514	8.559
Hypothesis 2	0.926	0.705	0.444
Degrees of freedom under			
Hypothesis 1 [N(T-K-1)]	221	222	222
Hypothesis 2 [N(T-1)-K]	208	209	209
F-statistics under			
Hypothesis 1	112.24 1/	749.58 1/	508.31 1/
(95 percent c.v.)	(1.5)	(1.7)	(1.7)
Hypothesis 2	5.28 1/	110.40 1/	38.49 1/
(95 percent c.v.)	(1.4)	(1.5)	(1.5)

Notes: Hypothesis 1: Homogeneous slope, homogeneous intercept. Hypothesis 2: Homogeneous slope, heterogeneous intercept.

1/ Null hypothesis can be rejected at the 5 percent significance level.

Table 3. Decomposition of Pooled Estimates to Within and Between Components

Estimates	Pool	Between (κ)	Within ($(1-\kappa)$)
Partition of sample = 14 countries			
α_1	0.240	0.176 (0.898)	0.596 (0.102)
α_2	-0.278	-0.290 (0.998)	-0.180 (0.002)
γ_1	0.075	0.025 (0.915)	0.610 (0.085)
η_1	0.009	0.013 (0.992)	-0.518 (0.008)
Partition of sample = 14 countries and 2 periods			
α_1	0.240	0.223 (0.974)	0.035 (0.026)
α_2	-0.278	-0.279 (0.997)	0.740 (0.003)
γ_1	0.075	0.070 (0.989)	0.601 (0.011)
η_1	0.009	0.0009 (0.999)	-0.466 (0.001)

Notes: Figures in parenthesis are the weights attached to the between and within estimates in producing the coefficient estimates of the pooled regression (i.e., pool=κ between + $(1-\kappa)$ within). Pool represents the pooled OLS estimates (see Table 1). The partition for the 2-period sample corresponds to 1970-78 and 1979-85.

these requirements are too demanding on this fragile dataset. To explore this hypothesis further, we examine the cross equation restriction $\gamma_1=(\alpha_1+\alpha_2)$ by estimating (I) and (II) using Zellner's seemingly unrelated regression technique. The Wald statistic reported in Table 4 states that we cannot reject the restriction. Failure to reject the restriction should be interpreted with caution as the t-ratios are small, implying the standard errors are large, and therefore that the test has low power. A possible interpretation of these results is that there is some degree of sectoral capital mobility but that it is less than perfect. Measurement errors in the capital stock may be another reason for the poor performance of (II).

We next attempt to determine whether there are any cross-country patterns related to productivity that can be exploited for estimation. To do this we use parameter restrictions related to the differential of total factor productivity growth from the Balassa-Samuelson model given in (5). In particular, recall that in steady-state, balanced-growth equilibrium, productivity growth in the tradables sector will be faster than that in the nontradables sector if $\alpha T>\alpha NT$. However, note that the intercept of (I) is (α^T/α^{NT}). This is a measure of the magnitude of the differential in productivity growth. Following this observation we use the parameter estimates from (I) to group countries by the degree to which they behave consistently with the Balassa-Samuelson hypothesis.

The individual country estimates reveal a group of countries for which the intercept is greater than 1, another group with intercepts less than 1 and an intermediate group with intercepts close to 1. We therefore classified the countries as *low-Balassa*, *medium-Balassa* or *high-Balassa* with 4 countries in the low-Balassa group: Denmark, Finland, Germany and United States; 6 countries in the medium-Balassa group: Australia, Belgium, England, France, Norway and Sweden; and 4 countries in the high-Balassa group: Canada, Italy, Japan, and the Netherlands. 1/

After grouping the countries by this criterion we estimate a fixed-effects model for equation (III). The results reported in Table 5 are striking. The explanatory power of the regression improves remarkably from the *low-Balassa* to the *high-Balassa* countries. The coefficients on the investment-output ratio for all countries are of the correct sign and statistically significant. 2/

Having established that (I) and (III) are reasonable empirical representations of the Balassa-Samuelson model we address some robustness issues. So far the entire analysis has been carried out with pooled and fixed-effects models. Fixed-effects is the appropriate statistical model

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1/ This grouping is admittedly arbitrary being based on casual observations of the productivity differential. It is, however, consistent with the literature that typically uses Japan as an example of a *high--* Balasa country (Marston 1987, Obstfeld 1993).

2/ Correcting for serial correlation did not change the pattern or the significance of the coefficient estimates.

Table 4. Seemingly Unrelated Regression of (I) and (II): Test of Cross Equation Restriction

Variable	Estimated Coefficients (t-Statistic) Equation (I)	Equation (II)
ky^T	0.030	0.031
	(0.9)	(0.9)
ky^{NT}	0.001	--
	(0.3)	--
Intercept	-0.001	-0.001
	(-0.025)	(-0.002)
Test of restriction $\gamma_1 = \alpha_1+\alpha_2$:		
Wald test statistic	0.045	
Significance Level 1/	0.832	

1/ Restriction is rejected with this significance level.

Table 5. Fixed-Effects Regression of Nontradable Price on Investment-Output Ratio by Groups Based on Model Restriction

Variable	Estimated Coefficients (t-ratio) Low-Balassa	Medium-Balassa	High-Balassa
iy^T	-0.778 1/ (-1.9)	-0.019 1/ (-2.3)	-0.675 2/ (-19.6)
Intercept	-0.034 (-1.3)	-0.093 2/ (-3.3)	-1.022 2/ (-16.2)
Adjusted R^2	0.012	0.042	0.863
F-statistic	1.80	5.19	384.6
Log-likelihood	98.5	101.8	47.6

Notes: Low-Balassa group: Denmark, Finland, Germany, and United States. Medium-Balassa group: Australia, Belgium, England, Norway and Sweden. High-Balassa group: Canada, Italy, Japan, and the Netherlands.

1/ Statistically significant at the 5 percent level.
2/ Statistically significant at the 1 percent level.

when the cross-section of countries represents the entire universe of interest. However, recall that we use data for 14 of the 24 OECD countries. This may raise some doubt as to the appropriateness of the fixed-effects model in the present circumstances. If one views the country-specific effects as randomly distributed across cross-sectional units then the appropriate methodology is a random-effects model.

We estimate a random-effects model by adopting the following component structure for the disturbances: $e_{jt}=\xi_j+\nu_{jt}$, where ξ_j are the country specific effects, and ν_{jt} are idiosyncratic shocks. If the right-hand-side variable is uncorrelated with both e_{jt} and ν_{jt} and if ν_{jt} is uncorrelated across time, then the standard variance components, generalized least squares (GLS) estimates are appropriate.

The results of the random-effects model estimated using GLS are reported in Table 6. While (III) performs well with coefficients that are statistically significant and of the correct sign, (I) and (II) yield wrong sign coefficients. To alleviate concerns about whether fixed or random-effects is the appropriate model we apply the Hausman specification test (Hausman 1978). The test resoundingly rejects the random-effects specification suggesting that the fixed-effects estimates are robust.

In section 3 we established the appropriateness of the smooth deterministic trend assumption imposed by the HP-filter. To verify that our empirical results are robust to the HP-filtering procedure we also carried out the entire estimation using the linear-trend filter. The result of estimating (III), presented in Table 7, shows there is little difference between the two procedures. 1/

In short, our results suggest that the Balassa-Samuelson proposition that relative marginal products of labor explain domestic relative prices in the long-run is well supported by the data in the total and fixed-effects models of equations (I) and (III). Furthermore, our results are not sensitive to the use of the HP-filter.

2. Evidence on the long-run real exchange rate

The evidence provided above supports the appropriateness of the Balassa-Samuelson model as a theory explaining long-run, cross-country differences in domestic relative prices. The next issue we address is the extent to which these differences can explain differences in long-run real

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1/ Results of estimating (I) and (II) with the linear-trend filter yield qualitative similar results to estimates reported above with the HP-filter.

Table 6. Random-Effects Regression of Nontradables Relative Price on Investment-Output and Capital-Output Ratios

Variable	Estimated Coefficients (t ratio) Equation (I)	Equation (II)	Equation (III)
ky^T	0.141 1/	0.579 2/	--
	(2.5)	(10.7)	--
ky^{NT}	0.573 2/	--	--
	(8.5)	--	--
iy^T	--	--	-0.369 2/
	--	--	(-14.8)
Intercept	-0.906 2/	-0.604 2/	-0.976 2/
	(-10.5)	(-7.6)	(-11.4)
Hausman-statistic (Fixed vs. random-effects)	53.97	5.86	92.52

1/ Statistically significant at the 5 percent level.
2/ Statistically significant at the 1 percent level.

Table 7. Comparison of Linear-Trend Filter and HP-Filter Fixed-Effects Regression of Relative Price of Nontradables on Investment-Output Ratio of Tradables

Variable	Estimated Coefficients (t-ratio) HP-Filter		Linear-Trend 1/	
iy^T		-0.518 2/ (-17.7)		-0.547 2/ (-20.0)
Group Dummies				
Australia	-2.25	(-17.3)	-2.40	(-19.7)
Belgium	-2.25	(-18.3)	-2.37	(-20.1)
Canada	-0.66	(-15.7)	-0.70	(-17.9)
Denmark	-2.25	(-17.7)	-2.38	(-20.1)
Finland	-1.88	(-17.4)	-2.00	(-20.0)
France	-0.91	(-18.5)	-0.96	(-21.4)
Germany	-0.99	(-18.3)	-1.04	(-20.8)
Italy	-0.85	(-15.7)	-0.91	(-18.6)
Japan	-0.87	(-19.3)	-0.93	(-22.6)
Netherlands	-0.59	(-8.2)	-0.67	(-10.0)
Norway	-1.96	(-17.6)	-2.06	(-20.0)
Sweden	-1.84	(-18.2)	-1.94	(-20.9)
United Kingdom	-1.15	(-17.9)	-1.21	(-20.5)
United States	-0.88	(-17.2)	-0.94	(-19.5)
Adjusted R^2		0.945		0.955
F-statistic		599.7		612.3
Log-likelihood		46.38		51.46

1/ Linear-trend filter values are the predicted values from a regression on a constant and a linear function of time.
2/ Statistically significant at the 1 percent level.

exchange rates. We focus on a log-linear version of (32). Assuming $\Omega^{*}_{t}=1$ yields the following testable equation: 1/

$$s_{jt}=\delta_{0j}+\delta_{1}p^{NT}_{jt}+e_{jt} \qquad (IV)$$

for j=1, 2, ..., M countries, and t=1, 2, ..., T time periods, where p^{NT} is the log of the relative price of nontradables, s is the log of the real exchange rate and e_{jt} are random disturbances.

Due to data limitations we use two separate real exchange rate series The IMF's CPI-based real exchange rates series for all 14 countries, which covers only part of our sample period (1975-85), and a GDP deflator-based real exchange rate series for the full sample period but for only 8 of the 14 countries from Micosi and Milesi (1993). As in the previous analysis we extracted the long-run growth component from the data by using the HP-filter.

Table 8 presents least squares estimates of a simple pooled linear regression of the CPI-based real exchange rates on both actual measures of relative prices, i.e, (IVa), and the predicted relative prices estimated from (III), i.e (IVb), for all 14 countries for the period 1975-85. As expected, from (IVa), a higher relative price of nontradables is positively associated with the real exchange rate. The coefficient estimates on the relative price of nontradables are positive though insignificantly different from zero. In (IVb) the coefficient is statistically significant at the 10 percent level in a one-tailed test. Moreover, note that the explanatory power of both the actual and the predicted nontradables price specifications is extremely low.

We also estimate a fixed-effects regression to examine the cross-country properties of this specification. The results reported in Table 9 show that the explanatory power is very high, but this is because within country intercepts are very good at tracking HP trends. The total coefficients δ_1 are still statistically insignificant, although with correct signs, and the between means coefficients, which illustrate the cross-sectional properties of the data, are insignificant and have incorrect signs.

1/ The more general case in which $0<\Omega^{*}_{t}\leq 1$ yields

$$s_{jt} = \lambda_{0i}+\lambda_{1}p^{NT}_{jt}+\sum_{j=1}^{k}\lambda_{1+j}p^{NT}_{jt}+e_{jt}$$

where the k's are the home country's trading partners and the null hypothesis is that $\lambda_1>0, \lambda_{1+j}<0$ for all j>1.

Table 8. Pooled (Total) Regression of CPI-Based Real Effective Exchange Rates on the Relative Price of Nontradables

$$s_{it}^{hp} = \delta_{0i} + \delta_1 p_{it}^{NT} + e_{it} \qquad (IVa)$$

$$s_{it}^{hp} = \delta_{0i} + \delta_1 \hat{p}_{it}^{NT} + e_{it} \qquad (IVb)$$

Variable	Estimated Coefficients	
	Equation (IVa)	Equation (IVb)
δ_1	0.274 (1.25)	0.315 (1.45)
δ_0	-1.169 (-1.15)	-1.378 (-1.37)
R^2	0.03	0.02

Notes: p^{NT} is the HP-filtered log relative price of nontradables, $\hat{p}^{NT}$ is the predicted value of p^{NT} from (III), and s^{hp} is the HP-filtered log real exchange rate. t-ratios are in parenthesis.

Table 9. Decomposition of Pooled Regression of CPI-Based Real Exchange Rates on the Relative Price of Nontradables

$$s_{it}^{hp} = \delta_{0i} + \delta_1 p_{it}^{NT} + e_{it} \qquad (IVa)$$

$$s_{it}^{hp} = \delta_{0i} + \delta_1 \hat{p}_{it}^{NT} + e_{it} \qquad (IVb)$$

Variable	Estimated Coefficients Equation (IVa)	Estimated Coefficients Equation (IVb)
Total (δ_1)	0.274 (1.25)	0.315 (1.45)
Between	-0.098 [0.40]	0.046 [0.20]
Within	-0.073 [0.60]	-0.080 [0.80]
Group Dummies (Num. of obs.)		
Australia	0.43 (1.5)	0.41 (1.6)
Belgium	0.39 (1.4)	0.37 (1.5)
Canada	0.37 (1.4)	0.40 (1.6)
Denmark	0.35 (1.2)	0.41 (1.6)
Finland	0.39 (1.4)	0.36 (1.5)
Germany	0.35 (1.3)	0.36 (1.4)
Italy	0.49 (1.8) 1/	0.49 (2.0) 1/
Japan	0.38 (1.4)	0.32 (1.3)
Netherlands	1.06 (3.8) 1/	1.06 (4.2) 1/
Norway	0.39 (1.4)	0.50 (1.9) 1/
Sweden	0.33 (1.2)	0.37 (1.4)
United Kingdom	0.36 (1.3)	0.41 (1.6)
United States	0.36 (1.4)	0.35 (1.5)
R^2	0.976	0.980

Notes: p^{NT} is the HP-filtered log relative price of nontradables. $\hat{p}^{NT}$ is the predicted value of p^{NT} from specification (III), and S^{hp} is the HP-filtered log real exchange rate. Weights, (κ), are in square brackets, see notes to Table 3 for details. t-ratios are in parenthesis.

1/ Statistically significant at the 5 percent level.

Tables 10 and 11 repeat the previous exercise with GDP deflator-based real exchange rates. Table 10 reports results for least squares estimates of a simple pooled regression using both actual relative prices and our predicted relative prices from (III) to explain the GDP deflator-based real exchange rates. None of the coefficient estimates are statistically significant and the explanatory power is still very low. With the fixed-effects regression (Table 11), the results remain poor. The explanatory power improves considerably for the same reason as above, but the coefficients on the price of nontradables have incorrect signs with one of them being statistically significant. Table 11 also reports estimates for the within and between regressions. These results indicate that unlike nontradable prices (see Table 3) in which much of the variation in the pooled OLS estimates is due to heterogeneity across country units, much of the variability in GDP deflator-based real exchange rates is due to "within" country factors. It appears that while the panel structure of the data was helpful in explaining the relative price of nontradables, it is less helpful in explaining long-run real exchange differentials.

Finally, aware of the limitations of our dataset, and the fact that our decomposition of tradables and nontradables is at best a rough approximation, we attempt to determine to what extent the inability of our relative price measure to explain real exchange rate behavior can be attributed to measurement errors. One, albeit limited way, to address this question is to use better quality data on tradables and nontradables from Kravis, Heston and Summers [1982] (KHS). We take the following data: from KHS: (i) the prices of tradable and nontradable goods from Table 6-12 and (ii) the measure of the GDP-based real exchange rates (the exchange rate deviation index) from their Table 1-2, both for 1975 and for 34 countries. We use these data to estimate a least squares regression of the log real exchange rate on the log of the ratio of nontradable and tradable prices. The results of this regression (a slope coefficient with correct sign and a t-statistic of 6.48, and an R^2 of 0.65) suggest that there may be a relationship between real exchange rates and the relative price of nontradables.

In conclusion, the results of the empirical tests of the second Balassa-Samuelson proposition suggest that, while international differences in the long-run relative price of nontradables reflect differences in sectoral marginal products of labor as predicted by the theory, these differences explain only a small fraction of long-run deviations from PPP based on aggregate price indexes. One interpretation of this evidence is to cast doubt on the validity of long-run PPP for tradables. However, significant measurement error, as suggested by the estimates obtained from the Kravis-Heston-Summers data, may also account for our findings. Furthermore, the tests we conducted embody nested hypotheses regarding the balanced-growth neoclassical framework and constant-elasticity utility and production functions.

Table 10. Pooled Regression of GDP Deflator-Based Real Effective Exchange Rates on the Relative Price of Nontradables

$$s_{it}^{hp} = \delta_{0i} + \delta_1 p_{it}^{NT} + e_{it} \qquad (IVa)$$

$$s_{it}^{hp} = \delta_{0i} + \delta_1 \hat{p}_{it}^{NT} + e_{it} \qquad (IVb)$$

Variable	Estimated Coefficients	
	Equation (IVa)	Equation (IVb)
δ_1	-0.921	-0.578
	(-0.6)	(-0.4)
δ_0	1.475	0.954
	(0.7)	(0.4)
R_2	0.04	0.02

Notes: p^{NT} is the HP-filtered log relative price of nontradables, $\hat{p}^{NT}$ is the predicted value of p^{NT} from specification (III), and s^{hp} is the HP-filtered log real exchange rate. t ratios are in parenthesis.

1/ Statistically significant at the 5 percent level.

Table 11. Decomposition of Pooled Regression of GDP Deflator-Based Real Exchange Rate on the Relative Price of Nontradables

$$s_{it}^{hp} = \delta o_i + \delta_1 p_{it}^{NT} + e_{it} \qquad (IVa)$$

$$s_{it}^{hp} = \delta o_i + \delta_1 \hat{p}_{it}^{NT} + e_{it} \qquad (IVb)$$

Variable	Estimated Coefficients: Equation (IVa)	Estimated Coefficients: Equation (IVb)
Total (δ_1)	-0.921	-0.578
	(-0.6)	(-0.4)
Between	-0.009	0.004
	[0.44]	[0.30]
Within	-1.610	-0.817
	[0.56]	[0.70]
Group Dummies		
Belgium	2.38 (2.5) 1/	1.17 (1.5)
Denmark	2.46 (2.6) 1/	1.23 (1.6)
France	2.40 (2.6) 1/	1.19 (1.5)
Germany	2.41 (2.6) 1/	1.19 (1.5)
Italy	2.56 (2.7) 1/	1.33 (1.7)
Netherlands	3.12 (3.3) 1/	1.91 (2.4) 1/
United Kingdom	2.38 (2.4) 1/	1.19 (1.5)
R^2	0.918	0.954

Notes: Weights, (κ), are in square brackets, see notes to Table 3 for details. p^{NT} is the HP-filtered log relative price of nontradables. $\hat{p}^{NT}$ is the predicted value of p^{NT} from specification (III), and S^{hp} is the HP-filtered log real exchange rate. t-ratios are in parenthesis.

1/ Statistically significant at the 5 percent level.

V. Concluding Remarks

In celebration of thirty years of the Balassa-Samuelson model, we have attempted to provide an appraisal of the static theory of Balassa (1964) and Samuelson (1964) by embedding it in an explicitly dynamic general equilibrium setting. Our appraisal of this celebrated model followed three stages. First, we derived two of the Balassa-Samuelson propositions as long-run, balanced-growth, implications of a two-country intertemporal equilibrium model. Second, we identified restrictions imposed on the cross-sectional, low-frequency behavior of the data implied by our model and thus derived testable predictions. Third, we constructed a cross-country sectoral database from existing OECD data and conducted econometric tests of the predictions of our model using panel data methods.

The empirical analysis suggests that the Balassa-Samuelson proposition that cross-country differences in long-run domestic relative prices of nontradables are determined by differences in the ratio of long-run sectoral marginal products of labor cannot be rejected by the data. However, we also found that long-run relative prices (as measured in the data or as predicted by our regressions) are of little help in explaining long-run, cross-country differences in the level of real exchange rates measured with CPI- or GDP deflator-based exchange rates. Thus, while the Balassa-Samuelson general equilibrium model performs well as a theory of relative prices, it seems to be unable to account for trend deviations from PPP. This statement echoes Paul Samuelson's quotation that prefaces the paper.

We conclude by pointing out some limitations of our work. On the empirical side, further work is required to develop a higher quality sectoral database covering a longer period and for a larger panel of countries. On the theoretical side, while we have succeeded in extending the static model to a dynamic setting, the simple deterministic neoclassical growth framework restricts our analysis to balanced-growth paths. Furthermore, an important assumption in our model is that Harrod-neutral technological progress expands at a constant rate. This assumption enables us to get a clear separation between trend growth and cycles and motivates the use of the HP-filter. However, such a clear separation fails if technological progress is stochastic or in models of endogenous growth. In a recent paper, Asea and Sturzenegger (1994) develop and test a Balassa-Samuelson type model based on an endogenous growth framework. Work along the lines carried out in this paper of developing robust general equilibrium restrictions that can be tested with the data will enhance our understanding of the enduring empirical regularities observed by Bela Balassa (1964) and Paul Samuelson (1964).

APPENDIX I

The Hodrick and Prescott [1090] filter is a two-sided filter that removes a trend that resembles a smooth curve drawn through the data. The HP filter defines a trend $\{\tau_t\}$ for a series $\{y_t\}$ as the solution to the following optimization problem:

$$\min_{\{\tau_t\}} \sum_{t=1}^{T} (y_t - \tau_t)^2 + \lambda \sum_{t=2}^{T-1} \left[(\tau_{t+1} - \tau_t) - \tau_t - \tau_{t-1})\right]^2$$

where λ is a parameter which penalizes changes in the trend component. The larger the value of λ the smoother the trend component. We chose λ= 400 which is consistent with other studies that use annual data. We also expėrimented with λ= 100 this value gave us no noticeable difference in results. All results reported in the text are for λ= 400. We used the RATS version 4.02 procedure HPFILTER.SRC to compute the trend and checked our results against a routine written in GAUSS.

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