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Government expenditure and equilibrium real exchange rates

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Abstract

Government expenditures (financed by lump-sum taxes) influence real exchange rates potentially via a *resource-withdrawal* channel and a *consumption-tilting* channel. Recent theoretical and empirical studies have considered only the effects of government spending through the resource-withdrawal channel. We solve for the theoretical relationships among the real exchange rate, relative private consumption, relative government consumption, and tradables and nontradables production in a two-country general equilibrium model and then estimate the model's structural equations. The results suggest that government expenditures influence real exchange rates *approximately equally* via the resource-withdrawal and consumption-tilting channels and that government and private consumption are *complements* in utility. © 2002 Elsevier Science Ltd. All rights reserved.

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1. Introduction

Economists have long investigated theoretically and empirically the relationship between government spending and real exchange rates, as recently highlighted in sections of Froot and Rogoff (1995), Obstfeld and Rogoff (1996), and Rogoff (1996). Frenkel and Razin (1996), however, noted that government spending influences the private sector and the real exchange rate potentially through two channels: the

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resource-withdrawal and *consumption-tilting* channels. Several recent empirical studies of the determinants of equilibrium real exchange rates have provided neoclassical models incorporating government spending amenable to econometric implementation. However, all have focused upon the resource-withdrawal channel for government spending, ignoring the potential relevance theoretically and empirically of complementarity vs. substitutability in utility of private and government consumption.

This paper introduces government expenditure into a model of two countries' representative consumers' behavior to investigate how differences in government expenditures between countries (alongside productivity differences) potentially can explain structural departures from purchasing power parity (PPP) and movements in equilibrium real exchange rates. Extending Frenkel and Razin, this paper generates closed-form theoretical solutions for the relationships among the real exchange rate, relative per capita private consumption, relative per capita government consumption, and relative per capita tradables and nontradables production in a general equilibrium model. Government expenditure influences consumers' utility explicitly as a potential substitute for or complement to private consumption and is allowed to fall on tradables as well as nontradables. Using data from the United Nations International Comparisons Program and OECD (1996), we estimate the model's structural and reduced-form equations.

The empirical results suggest three interesting conclusions. First, government expenditure influences equilibrium real exchange rates in the medium-run via both the resource-withdrawal and consumption-tilting channels. Second, the consumption-tilting effect is *approximately equal* to the resource-withdrawal effect. Third, government and private consumption are *complements* in utility. This suggests that a real appreciation of a country's currency associated with a larger government expenditure share reflects partially the shadow price and associated benefit of this complementarity.

The remainder of this paper is as follows. Section 2 provides a brief literature review. Section 3 describes the theoretical model. In Section 4, we discuss data constraints, stationarity issues, and parameter-identification issues associated with the econometric analogues to the intertemporal and intratemporal equilibrium conditions derived in Section 3. Section 5 presents empirical results from estimating these equilibrium conditions and finds plausible estimates of the elasticity of intertemporal substitution (approximately one-half), the relative share of nontradable goods in utility (approximately one-half), and the role of government expenditure in explaining relative consumption levels and real exchange rate movements. Section 6 concludes that government spending can influence equilibrium real exchange rates via both the resource-withdrawal *and* consumption-tilting channels and contrasts our results with associated estimates in the literature. Section 7 concludes.

2. Literature background

Frenkel and Razin (1996) summarize nicely the theoretical relationship between government spending and real exchange rates in an intertemporal, neoclassical framework.[1] In the context of a two-period, small open economy model, Frenkel and Razin note that government spending influences the private sector and the real exchange rate essentially through two channels: the *resource-withdrawal* and *consumption-tilting* channels. Regarding the first channel, the influence of a government expenditure increase is similar to that of a negative supply shock; the effect on private consumption and the real exchange rate will depend upon the proportion of government consumption falling on nontradables vs. that falling on tradables. Regarding the second channel, Frenkel and Razin point out that the effect of government expenditure on private consumption levels and the real exchange rate will depend upon the "characteristics of the utility function" (p. 165). They note the potential importance of complementarity vs. substitutability between private consumption and government consumption in utility, which dictates how the marginal rate of intertemporal substitution in utility is influenced by government expenditure levels.

Several recent studies of the determinants of equilibrium real exchange rates have provided neoclassical models amenable to econometric implementation—notably Froot and Rogoff (1991), Rogoff (1992), De Gregorio et al. (1994a, b), De Gregorio and Wolf (1994), and Chinn and Johnston (1996). Although each of these empirical studies incorporates government spending, all have focused upon the resource-withdrawal channel. De Gregorio et al. (1994a, b), and De Gregorio and Wolf (1994) present static models where government expenditure (financed by lump-sum taxes) falls exclusively on nontradables and the effect on the real exchange rate is entirely through the resource-withdrawal channel. They find a significant short-run (but not long-run) effect of government spending on the real exchange rate. In intertemporal neoclassical contexts, Froot and Rogoff (1991) and Chinn and Johnston (1996) find significant empirical effects of the share of government expenditures on real exchange rates. However, as in the static models above, government spending only works through resource withdrawal, though both models allow government spending on both tradables and nontradables.[2]

While broadly aimed at a better understanding of the potential distinction between the resource-withdrawal and consumption-tilting effects of government expenditure, this study sheds light additionally upon a related issue. Backus and Smith (1993) found weak empirical support for a monotonic relationship between the real exchange rate and relative consumption. Because changes in relative prices and relative consumptions are driven in their model by endowment shocks to tradables and nontradables, the authors qualified their weak empirical results by noting prominently that a "more general model"—that is, one admitting demand-side shocks"—might help sort out the issues. The present model employs changes in relative government

[1] Throughout we are concerned with government spending, not fiscal policy. Government spending is financed here by lump-sum taxes. Since Ricardian equivalence holds in our context, budget deficits related to the timing of taxation are not relevant here.

[2] Penati (1987) addressed government spending (potentially on tradables and nontradables) and equilibrium real exchange rates in an intertemporal neoclassical theoretical model, but similarly focused only upon the resource-withdrawal channel. This paper did not provide empirical investigation, but Penati (1986) did.

expenditures to consider such demand-side shocks' effects on the real exchange rate in a framework similar to Balvers and Bergstrand (1997).

3. A two-country equilibrium model with government expenditures

Frenkel and Razin (1996, sections 8.2–8.4) describe theoretically the relationship between government expenditure and the real exchange rate using an intertemporal neoclassical model. In the context of a small open economy, they outline the potential complementary roles for government spending influencing private consumption behavior via the resource-withdrawal and consumption-tilting channels. However, in their two-country extension (when either country can influence the world interest rate), they introduce government spending into the utility function in a "separable way," implying that the marginal rate of intertemporal substitution in private consumption "does not depend on the level of government spending" (p. 262).

In the following, we outline a two-country model where both the consumption-tilting and resource-withdrawal channels of government expenditure are introduced explicitly. The model is a direct extension of Balvers and Bergstrand (1997), amended to include a government sector. We first describe the model and then derive intertemporal and intratemporal equilibria for specific preferences. Estimable closed-form solutions are obtained for a nested utility function with constant relative risk aversion in consumption and with some symmetry conditions imposed on the preferences of the two representative consumers.

3.1. Description of the Model

Following Lucas (1978), we abstract from investment decisions by assuming endowment economies. Stochastic production processes for all goods are owned by the consumers and yield perishable outputs. Each country (foreign variables denoted by $*$) consists of a tradables production process, a nontradables production process, and one infinitely-lived representative consumer with a time-additive utility function. The tradables produced in both countries are perfect substitutes.[3]

In the home country and analogously abroad, the representative consumer maximizes the expected present discounted value of the stream of future utilities from private consumption and the consumption of publicly provided goods:

[3] Since the production processes are exogenous, the consumers take outputs as givens (and factors are fixed). The exogenous output structure suggests a model that explains medium-run, rather than long-run, real exchange rate behavior. As discussed in Obstfeld and Rogoff (1996, Chapter 4), long-run equilibrium real exchange rates should be determined by supply-side factors alone, as suggested by the Balassa-Samuelson hypothesis; relative government expenditures and other relative demand shocks should not matter with homothetic preferences in the long run. However, with exogenous outputs, relative government expenditures will affect equilibrium real exchange rates in the medium run. This is consistent with empirical studies such as De Gregorio et al. (1994b), where government expenditures had short-run, but not long-run, effects on real exchange rates

$$\mathrm{Max}E_0 \sum_{t=0}^{\infty} \beta^t u(c_t,g_t) \tag{1}$$

$$c_t = v(c_t^{\mathrm{T}},c_t^{\mathrm{N}}), \tag{2}$$

$$g_t = w(g_t^{\mathrm{T}},g_t^{\mathrm{N}}), \tag{3}$$

where E_0 represents the expectation conditional on information at time 0, β is the standard discount factor, and $u(\cdot)$ is a current-period utility function strictly concave in the consumption index c_t and dependent on g_t (per capita government expenditure). The consumption index at time t, c_t in Eq. (2), represents an optimally chosen basket of the tradable good, c_t^{T}, and the nontradable good, c_t^{N}. Preferences over both goods embodied in the consumption index are assumed to be homothetic in order that an exact price index may be defined. Without loss of generality, we then may apply a monotonic transformation such that $v(\cdot)$ in Eq. (2) is homogeneous of degree one and concave in the decision variables. In Eq. (3), $\mathrm{g_t}$ represents government consumption, which is produced from government purchases of tradables, g_t^{T}, and nontradables, g_t^{N}, via function $w(\cdot)$. Government consumption is exogenous to the individual consumer and is financed by lump-sum taxes.

To obtain closed-form solutions for the real exchange rate and relative consumption (which also make the model amenable to empirical research), we assume a constant relative risk aversion (CRRA) current-period utility function for consumption but with generally specified preferences for the government expenditures [Eq. (4)], and a Cobb-Douglas subutility function [Eq. (5)]:

$$u(c_t,g_t) = z(g_t) + h(g_t)(c_t)^{1-\sigma}/(1-\sigma), \;\; \sigma > 0 \tag{4}$$

$$c_t = (c_t^{\mathrm{T}})^{1/(1+\gamma)}(c_t^{\mathrm{N}})^{\gamma/(1+\gamma)}, \;\; \gamma > 0. \tag{5}$$

Letting subscripted variables denote first derivatives, assume $z_g(g) > 0$. Then government consumption (g) and private consumption (c) are complements whenever $h_g(g) > 0$ and substitutes whenever $h_g(g) < 0$. (Note that in the case of substitutes, it can be assured that $u_g(c,g) > 0$ only if $z_g(g) > 0$. The preferences of the two countries' representative consumers may differ in discount factors β and β^* *and* in functions $h(g)$ and $z(g)$ vs. $h*(g)$ and $z*(g)$, respectively, but parameters σ and γ are assumed identical across countries.

The nested structure of the utility function enables us to separate the representative consumer's decisions into an intratemporal decision concerning the distribution of overall consumption expenditure between the tradable and nontradable goods and an intertemporal decision concerning the demand for assets and the overall expenditure on current consumption. Thus, the two equilibrium conditions can be discussed separately.

Assume complete markets so that productivity shocks are insurable.[4] The Second

[4] Given the CRRA utility and perfect capital mobility assumptions, it can be shown that the existence of claims for all production sectors, representing shares in the stochastic endowment processes, is sufficient to effectively complete markets in our model.

Welfare Theorem now implies that the competitive equilibrium outcome will be equivalent to the Pareto optimal allocation as chosen by a social planner. This social planner maximizes the weighted average of the lifetime utilities of the two countries' representative consumers by choosing the distribution of the tradable good subject to the available quantity. Due to the time separability of the lifetime utility functions, the social planner solves the following decision problem for each period:

$$\max_{c_t^{\mathrm{T}},c_t^{\mathrm{T}*}} \{\beta^t[z(g_t) + h(g_t)(c_t)^{1-\sigma}/(1-\sigma)] + \alpha\beta^{*t}[z^*(g_t^*) + h^*(g_t^*)(c_t^*)^{1-\sigma}/(1-\sigma)]\} \tag{6}$$

subject to:

$$c_t = (c_t^{\mathrm{T}})^{1/(1+\gamma)}(c_t^{\mathrm{N}})^{\gamma/(1+\gamma)},\ c_t^* = (c_t^{\mathrm{T}*})^{1/(1+\gamma)}(c_t^{\mathrm{N}*})^{\gamma/(1+\gamma)}, \tag{7}$$

$$c_t^{\mathrm{N}} + g_t^{\mathrm{N}} = y_t^{\mathrm{N}},\ c_t^{\mathrm{N}*} + g_t^{\mathrm{N}*} = y_t^{\mathrm{N}*} \tag{8}$$

$$c_t^{\mathrm{T}} + g_t^{\mathrm{T}} + c_t^{\mathrm{T}*} + g_t^{\mathrm{T}*} = y_t^{\mathrm{T}} + y_t^{\mathrm{T}*}. \tag{9}$$

Let α denote the constant weight that the social planner places on the utility of the foreign consumer, $y_t^{\mathrm{N}}(y_t^{\mathrm{N}*})$ is per capita nontradables output in the home (foreign) country in period t, and $y_t^{\mathrm{T}}(y_t^{\mathrm{T}*})$ is per capita tradables output in the home (foreign) country in period t.[5] Eq. (8) represents the period market-clearing conditions for nontradable goods in the two countries. Eq. (9) represents the period market-clearing condition for tradables goods, reflecting that the countries' tradable goods are perfect substitutes. As stated earlier, closed-form solutions are obtained by constraining some parameters of the utility functions to be equal, but the representative consumers can still differ with respect to rates of time preference, initial endowments, benefits derived from government expenditure, and consumption opportunities related to the nontradable good.

Finally, we define the real exchange rate, x_t, conventionally as:

$$x_t = e_t p_t^*/p_t, \tag{10}$$

where e_t is the nominal exchange rate (expressed in units of domestic currency per

[5] We note two issues. First, α depends on initial relative wealth of the two consumers and the exogenous processes for government expenditures. Our formulation, in which the social planner chooses private expenditure but not public expenditure, may appear paradoxical. However, it corresponds to a market outcome for the standard case where individuals choose their consumption subject to exogenously determined government expenditure. Second, for simplicity we have modeled the social planner as maximizing a weighted average of lifetime utilities of representative consumers in each of two countries. Alternatively, to allow for differing populations, we could have constructed the model to have n ($n*$) identical consumers in the home (foreign) country. In this case, the first (second) RHS term in Eq. (6) would be scaled by n ($n*$) and, in Eq. (9), c_t^{T}, g_t^{T}, and y_t^{T}would each be scaled by n and $c_t^{\mathrm{T}*}$, $g_t^{\mathrm{T}*}$, and $y_t^{\mathrm{T}*}$ would each be scaled by $n*$. In the theoretical results that follow, Eqs. (10) through (17) would not be affected by this alternative approach. Our formulation allows quantity variables to be interpreted in per capita terms, consistent with the empirical work.

unit of foreign currency) and p_t (p_t^*) is the domestic (foreign) consumption-based price level.[6]

3.2. Intratemporal equilibrium

The separability of the lifetime utility function implies a budgeting process where the consumer in each period maximizes the (intratemporal) value of the consumption index, c_t, in Eq. (5), subject to an after-tax budget constraint, B_t, given by:

$$B_t = p_t^{\mathrm{T}} c_t^{\mathrm{T}} + p_t^{\mathrm{N}} c_t^{\mathrm{N}}, \tag{11}$$

where p_t^{T} (p_t^{N}) is the domestic consumer price for the tradable (nontradable), and similarly for the foreign consumer. Choosing consumption of the tradable good to maximize Eq. (5) subject to Eq. (11) yields:

$$\lambda_t = (1 + \gamma)^{-1}(c_t^{\mathrm{T}})^{-\gamma/(1+\gamma)}(c_t^{\mathrm{N}})^{\gamma/(1+\gamma)}/p_t^{\mathrm{T}}, \tag{12}$$

and similarly for the foreign country, where λ_t is the Lagrange multiplier.

Given homotheticity of the subutility function (c_t), one can rewrite Eq. (11) as:

$$B_t = p_t c_t. \tag{13}$$

Since the Lagrange multiplier equals the marginal benefit of a unit increase in the budget on the maximum subutility, we also have:

$$\lambda_t = \partial c_t / \partial B_t = 1/p_t. \tag{14}$$

Combining Eqs. (12) and (14), and similarly for the foreign country, using the market-clearing conditions for the nontradable goods, the law of one price for tradable goods ($p_t^{\mathrm{T}} = e_t p_t^{\mathrm{T}*}$), and the definition of the real exchange rate, yields the intratemporal equilibrium condition:

$$x_t = (c_t^*/c_t)^{\gamma}[(y_t^{\mathrm{N}*} - g_t^{\mathrm{N}*})/(y_t^{\mathrm{N}} - g_t^{\mathrm{N}})]^{-\gamma}. \tag{15}$$

Eq. (15) illustrates the expected *resource-withdrawal* effect of government spending discussed earlier. A rise in per capita foreign government purchases of nontradables causes a rise in the real exchange rate, that is, a real appreciation of the foreign currency.

3.3. Intertemporal equilibrium

In the intertemporal stage of the budgeting process, the social planner maximizes in each period Eq. (6), the weighted average of the two consumers' utilities, subject to the constraints in Eqs. (7), (8) and (9). This yields a set of first-order conditions that, by eliminating the Lagrange multiplier, results in:

[6] The tradable good may serve as the numeraire in both countries, implying e_t=1 in Eq. (10). If money is introduced explicitly—for instance, via a binding cash-in-advance constraint, m_t=$p_t c_t$ (and analogously abroad)—then the prices of the tradable goods will be determined within the model as will the nominal exchange rate.

$$\frac{(c_t^{\mathrm{T}})^{-\gamma/(1+\gamma)}(c_t^{\mathrm{N}})^{\gamma/(1+\gamma)}}{(c_t^{\mathrm{T}*})^{-\gamma/(1+\gamma)}(c_t^{\mathrm{N}*})^{\gamma/(1+\gamma)}} = \alpha\,\frac{\beta^{*t}h^*(g_t^*)(c_t^*)^{-\sigma}}{\beta^t h(g_t)(c_t)^{-\sigma}}. \tag{16}$$

From the intratemporal equilibrium, the left hand side of Eq. (16) equals x_t so that the intertemporal equilibrium condition is:

$$x_t = \alpha(\beta^*/\beta)^t[h^*(g_t^*)/h(g_t)](c_t^*/c_t)^{-\sigma}. \tag{17}$$

where α and $(\beta^*/\beta)^t$ can be interpreted in the model's context as initial and accumulated relative wealth, respectively, and σ can be interpreted as the coefficient of relative risk aversion (inverse of the elasticity of intertemporal substitution).[7] Eq. (17) illustrates the *consumption-tilting* effect of government spending discussed earlier and the dependence of the marginal rate of intertemporal substitution on relative government spending, omitted in recent studies of government spending and equilibrium real exchange rates. Higher per capita foreign government consumption may increase or decrease the real exchange rate, depending upon the complementarity or substitutability between government consumption and private consumption.

3.4. Interpretation

Fig. 1 demonstrates the intratemporal equilibrium locus, INTRA$_t$, in logarithmic form with slope γ. It has a positive slope since a higher desired foreign consumption level implies more demand so that the foreign price level and, accordingly, the real exchange rate needed for intratemporal equilibrium are higher. An increase in foreign government purchases of nontradables ($g_t^{\mathrm{N}*}$) withdraws resources from foreign nontradables production. The resulting excess demand drives up the relative price of nontradables abroad and the real exchange rate, lowering relative foreign private nontradables consumption expenditure (point A to point B in Fig. 1).

Fig. 1 also displays in logarithmic terms the negative relationship at time t between relative consumption and the real exchange rate along the intertemporal equilibrium locus, INTER$_t$, with slope $-\sigma$. A higher relative foreign price level (i.e., real exchange rate) currently provides relatively more incentive for the foreign consumer to defer consumption to the future so that equilibrium current consumption abroad declines relative to the home country. An increase in foreign government consumption (g_t^*) can shift the INTER$_t$ curve up or down depending upon whether government consumption complements or substitutes for private consumption. If government consumption complements private consumption (i.e., $h_g(g) > 0$), INTER$_t$ shifts up, raising the real exchange rate and relative consumption (point A to point C in Fig. 1). If government consumption substitutes for private consumption (i.e., $h_g(g) < 0$), the

[7] Note that the intertemporal equilibrium condition is neither stochastic nor explicitly dynamic. Although the model is derived in a stochastic setting, the assumption of complete markets implies optimal risk sharing so that *relative* prices and *relative* consumption levels are deterministic. Moreover, the dynamic Euler equation linking the marginal rates of intertemporal substitution of the representative consumers to intertemporal price ratios in a market economy can be converted to the static version of Eq. (17) when couched in the relative consumption allocations assigned by a social planner.

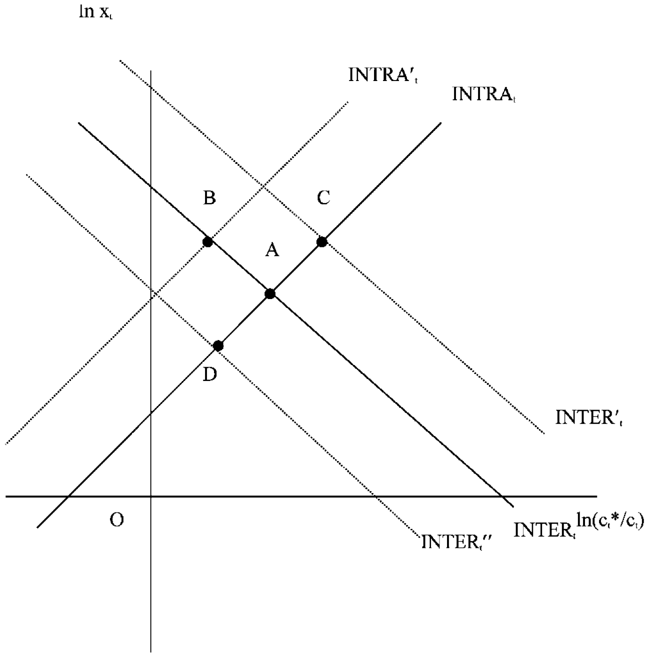

Fig. 1. Resource-withdrawal vs. consumption-tilting effects of relative government expenditures.

reverse happens with a rise in foreign government expenditures (point A to point D in Fig. 1). In Fig. 1, note the four possible quadrants for the equilibrium. Initial equilibrium point A assumes the foreign country has higher per capita consumption and price level than the home country.

The model can now be solved explicitly for the real exchange rate and relative consumption expenditure, as suggested by Fig. 1. For empirical purposes, assume $h(g) = g^{\delta}$ where $\delta > 0$ ($\delta < 0$) indicates government consumption complements (substitutes for) private consumption; analogously, let $h^*(g^*) = g^{*\delta}$. Combining Eqs. (15) and (17) then yields reduced forms:

$$x_t = \alpha^{\gamma/(\gamma+\sigma)}(\beta^*/\beta)^{[\gamma/(\gamma+\sigma)]t}[(y_t^{N*}-g_t^{N*})/(y_t^{N}-g_t^{N})]^{-\gamma\sigma/(\gamma+\sigma)}(g_t^*/g_t)^{\delta\gamma/(\gamma+\sigma)}, \qquad (18)$$

$$c_t^*/c_t = \alpha^{1/(\gamma+\sigma)}(\beta^*/\beta)^{[1/(\gamma+\sigma)]t}[(y_t^{N*}-g_t^{N*})/(y_t^{N}-g_t^{N})]^{\gamma/(\gamma+\sigma)}(g_t^*/g_t)^{\delta/(\gamma+\sigma)}. \qquad (19)$$

Initial wealth, rates of time preference, nontradables productivity *and* government expenditures all affect relative consumption and the real exchange rate.[8] Eq. (19),

[8] The reader familiar with the Balassa-Samuelson hypothesis may be concerned in Eq. (18) over the omission of relative tradables productivity levels. However, the productivity-differentials theory of Balassa and Samuelson is imbedded in our intratemporal equilibrium. In his analysis, Balassa (1964) assumes “invisibles and capital movements do not enter the balance of payments,” thus avoiding the relevance of the intertemporal equilibrium. It follows that the trade balance must be assumed equal to zero so that not only $y_t^{N} = c_t^{N}$ but also $y_t^{T} = c_t^{T}$. Combining Eqs. (7) and (15), assuming no government expenditures, with market clearing produces:

moreover, implies that consumption levels in the two countries need not be highly correlated, as long as relative nontradables productivity or relative government expenditure varies substantially.

4. Data and econometric issues

The reduction of the general equilibrium model to closed-form equilibrium conditions (15) and (17) and to reduced-form real exchange rate and relative consumption Eqs. (18) and (19), respectively, allows estimation. In this section, we discuss the choice of data and relevant econometric issues.

4.1. Data issues

Given the model describes medium-run behavior, a data set consistent with such behavior is optimal. Annual data may be most appropriate to evaluate a medium-run theoretical model. Higher frequency data, such as monthly or quarterly time series, would be more appropriate for explaining short-run real exchange rate behavior. Lower-frequency data, such as that averaged over five-year to fifteen-year intervals, is typically used for explaining long-run economic behavior. For a similar categorization, see De Gregorio et al. (1994b).

The econometric analogues to Eqs. (15), (17), (18) and (19) require for estimation data on relative consumption price levels, relative per capita private consumption expenditures, relative per capita government consumption expenditures, and relative per capita private nontradables consumption expenditures; in the context of the model, i.e., Eq. (8), $(y^{N*}-g^{N*})/(y^{N}-g^{N}) = c^{N*}/c^{N}$. Since our model has implications for both the time-series and cross-sectional behavior of real exchange rates, a panel data set is potentially useful. The Penn World Tables Version 5.6a, accessible at the National Bureau of Economic Research website (www.nber.org), provides annual time series from 1950 to 1992 on relative consumption price levels (or consumption-based real exchange rates), relative per capita private consumption expenditures, and relative per capita government consumption expenditures for over one-hundred countries with the United States as the numeraire (US=100), designed for pooled cross-section time-series analysis; this is the data of the United Nations (UN) International Comparisons Program (ICP), which follows closely the UN's System of National Accounts (SNA). The only shortcoming of this dataset for our estimation of the econometric analogues to Eqs. (15), (17), (18) and (19) is the absence of a decomposition of private consumption expenditures between tradables and nontradables.

Fortunately, the *OECD Annual National Accounts, Volume II* (1996) enables construction of relative shares of private consumption expenditures into nontradables

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$$x_t = [(y_t^{T*}/y_t^{N*})/(y_t^{T}/y_t^{N})]^{\gamma/(1+\gamma)}$$

which is the Balassa–Samuelson relationship.

and tradables using a similar categorization as in Kravis et al. (1982, Chapter 2) for the United States and ten other OECD countries over the years 1970 through 1990.[9] Specifically, the *OECD Annual National Accounts, Volume II* decomposes private final consumption expenditures *in the domestic market* of resident households into goods and services using the UN's SNA classifications. The eight classifications of final private consumption expenditures are each decomposed into commodities and services, the criterion distinguishing them being that a commodity (service) is storable (nonstorable). Kravis et al. (1982, Appendix Table 2-1) define "nontradables" in consumption as "all service categories" (p. 69). Using the OECD data and Penn World Tables, "c^N" is per capita private final consumption expenditures on services in the domestic market by resident households. Examples of these services from each classification include repair services, gross rents, electricity services, domestic services, certain transport services, public entertainment services, and restaurant and hotel services (all purchased, as noted above, in the *domestic* market). Since goods *and* services purchased abroad by home residents are essentially "tradable," we add these expenditures of home residents to private final consumption expenditures in the domestic market of resident households to arrive at "c." Hence "c" is per capita private final consumption expenditures of goods and services by resident households (purchased directly at home or abroad).

4.2. Stationarity issues

The log-linear versions of reduced-form Eqs. (18) and (19) potentially estimable by ordinary least squares (OLS) are:

$$\ln x_{it} = \sum_{j=1}^{10} \Phi_j^1 + \sum_{j=1}^{10} \{[\gamma/(\gamma+\sigma)][\ln(\beta_j/\beta)]\}\mathrm{trend}_{jt} \tag{20}$$
$$-\gamma[\sigma/(\gamma+\sigma)]\ln[(y_{it}^N - g_{it}^N)/(y_t^N - g_t^N)] + [\delta\gamma/(\gamma+\sigma)]\ln(g_{it}/g_t) + \varepsilon_{it}^1,$$

$$\ln(c_{it}/c_t) = \sum_{j=1}^{10} \Phi_j^2 + \sum_{j=1}^{10} \{[1/(\gamma+\sigma)][\ln(\beta_j/\beta)]\}\mathrm{trend}_{jt} \tag{21}$$
$$+ [\gamma+\sigma]\ln[(y_{it}^N - g_{it}^N)/(y_t^N - g_t^N)] + [\delta/(\gamma+\sigma)]\ln(g_{it}/g_t) + \varepsilon_{it}^2,$$

where x_{it} is the real exchange rate of country i relative to the United States in year t, c_{it} (c_t) is per capita consumption in country i (US), $y_{it}^N - g_{it}^N(y_t^N - g_t^N)$ represents per capita nontradables production net of government absorption in country i (US), and

[9] The ten other OECD countries are Austria, Canada, Denmark, Finland, France, Greece, Italy, Japan, Norway, and the United Kingdom.

g_{it} (g_t) is per capita government consumption in country i (US).[10] Variable trend$_{jt}$ is a time trend when j=i and is 0 for $j \neq i$. β_j/β is the discount rate in country j relative to the US discount rate when j=i. The εs are i.i.d. error terms and Φ_j^1 and Φ_j^2 are dummy variables assuming the value 1 when j=i and 0 when $j \neq i$, to reflect exogenous differences in initial wealth.[11]

Since the data set includes time series as well as cross section observations, one needs to address the issue of data stationarity over time. If the individual time series are stationary in log-levels, coefficient estimates in these regressions are consistent and the student t-distribution can be used to evaluate their statistical significance. However, if the individual time series are stationary in first-differences of their log-levels but the series are cointegrated, OLS coefficient estimates are consistent but standard *t*-statistics would have to be adjusted to evaluate the coefficients' statistical significance.

Although the time-series dimension of our data is fairly small by most time-series analyses standards, Levin and Lin (1992) show theoretically the increased power of testing for unit roots in short time series by pooling data cross-sectionally. In particular, they demonstrate that when the time dimension is of "moderate length (i.e., 25–100 periods)", pooling data cross-sectionally with only a small number of individuals can dramatically increase the power of the unit-root test. They note that with only 25 time-series periods but a panel of 10 units (similar to this study), the power of the test exceeds 90%. Three studies (Frankel and Rose, 1996; Oh, 1996; Wu, 1996) found evidence that real exchange rates are stationary in log-levels by pooling time-series data across country pairs.

Using the techniques in Levin and Lin (1992), we conducted similar tests of stationarity for our relative price level, relative (per capita) consumption level, and relative (per capita) government consumption level variables for the 37-year period 1953 to 1990, pooling data across all ten country pairs (relative to the United States).[12] The results of the tests for nonstationarity are presented in Tables 1–3. The tests indicated overwhelmingly that the null hypothesis of nonstationarity could

[10] In the context of the model,$(y_{it}^{\mathrm{N}}-g_{it}^{\mathrm{N}})/(y_t^{\mathrm{N}}-g_t^{\mathrm{N}})$ can be measured by $c_{it}^{\mathrm{N}}/c_t^{\mathrm{N}}$. However, due to the potential endogeneity of this RHS variable,$(c_{it}^{\mathrm{N}}c_t^{\mathrm{N}})$ is replaced by an instrument—$pv[(y_{it}^{\mathrm{N}}-g_{it}^{\mathrm{N}})/(y_t^{\mathrm{N}}-g_t^{\mathrm{N}})]$—created from its lagged value and a constant. Although a Hausman specification test did not indicate evidence of endogeneity, the results presented in this paper use the instrument to ensure that this RHS variable is predetermined. Results using $c_{it}^{\mathrm{N}}/c_t^{\mathrm{N}}$ (instead of the instrument) provide a similar fit and the coefficient estimates are not materially different. The latter results are omitted here for brevity, but are available upon request.

[11] For econometric convenience, we use an intercept and 9 dummy variables.

[12] As noted earlier, data was available for these three variables for the period 1950–1990; data was available for relative nontradables consumption levels only for the much shorter period 1970–1990. Since unit root tests for the shorter period for all variables would have had a time dimension (with necessary lags) of only 17 time periods, we considered this time dimension too short to benefit from the power of the Levin-Lin tests. Consequently, we chose to conduct the unit root tests for all the variables (except relative nontradables consumption) for the longer period (1953–1990). The results are for tests where the test statistics are asymptotically normally distributed. We did not include tests with individual fixed effects; these effects were statistically insignificant.

Table 1
Test for nonstationarity of relative consumption price levels using panel data and Levin–Lin (1992)

Model	ρ_0-1 estimate	t-statistic	10% (5%) Critical t	Null hypothesis of nonstationarity[a]
1a. $\Delta y_{it} = (\rho_0-1)y_{it-1} + \varepsilon_{it}$	−0.05	−3.74	−1.44 (−1.81)	Reject (Reject)
1b. $\Delta y_{it} = (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \varepsilon_{it}$	−0.06	−4.19	−1.44 (−1.81)	Reject (Reject)
1c. $\Delta y_{it} = (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \rho_2\Delta y_{it-2} + \varepsilon_{it}$	−0.05	−3.75	−1.44 (−1.81)	Reject (Reject)
2a. $\Delta y_{it} = \delta_0 + (\rho_0-1)y_{it-1} + \varepsilon_{it}$	−0.04	−2.53	−1.72 (−2.09)	Reject (Reject)
2b. $\Delta y_{it} = \delta_0 + (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \varepsilon_{it}$	−0.05	−3.55	−1.72 (−2.09)	Reject (Reject)
2c. $\Delta y_{it} = \delta_0 + (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \rho_2\Delta y_{it-2} + \varepsilon_{it}$	−0.05	−2.94	−1.72 (−2.09)	Reject (Reject)
3a. $\Delta y_{it} = \delta_0 + \delta_1 t + (\rho_0-1)y_{it-1} + \varepsilon_{it}$	−0.06	−3.53	−2.01 (−2.39)	Reject (Reject)
3b. $\Delta y_{it} = \delta_0 + \delta_1 t + (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \varepsilon_{it}$	−0.09	−5.10	−2.01 (−2.39)	Reject (Reject)
3c. $\Delta y_{it} = \delta_0 + \delta_1 t + (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \rho_2\Delta y_{it-2} + \varepsilon_{it}$	−0.08	−4.61	−2.01 (−2.39)	Reject (Reject)

[a] At the 10% (5%) significance level. Numeraire country is US.

Table 2
Tests for nonstationarity of relative per capita consumption levels using panel data and Levin–Lin (1992)

Model	ρ_0-1 estimate	t-statistic	10% (5%) Critical t	Null hypothesis of nonstationarity[a]
1a. $\Delta y_{it} = (\rho_0-1)y_{it-1} + \varepsilon_{it}$	−0.02	−11.94	−1.44 (−1.81)	Reject (Reject)
1b. $\Delta y_{it} = (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \varepsilon_{it}$	−0.02	−8.38	−1.44 (−1.81)	Reject (Reject)
1c. $\Delta y_{it} = (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \rho_2\Delta y_{it-2} + \varepsilon_{it}$	−0.02	−7.90	−1.44 (−1.81)	Reject (Reject)
2a. $\Delta y_{it} = \delta_0 + (\rho_0-1)y_{it-1} + \varepsilon_{it}$	−0.04	−8.96	−1.72 (−2.09)	Reject (Reject)
2b. $\Delta y_{it} = \delta_0 + (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \varepsilon_{it}$	−0.03	−6.56	−1.72 (−2.09)	Reject (Reject)
2c. $\Delta y_{it} = \delta_0 + (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \rho_2\Delta y_{it-2} + \varepsilon_{it}$	−0.03	−6.30	−1.72 (−2.09)	Reject (Reject)
3a. $\Delta y_{it} = \delta_0 + \delta_1 t + (\rho_0-1)y_{it-1} + \varepsilon_{it}$	−0.04	−7.72	−2.01 (−2.39)	Reject (Reject)
3b. $\Delta y_{it} = \delta_0 + \delta_1 t + (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \varepsilon_{it}$	−0.03	−5.98	−2.01 (−2.39)	Reject (Reject)
3c. $\Delta y_{it} = \delta_0 + \delta_1 t + (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \rho_2\Delta y_{it-2} + \varepsilon_{it}$	−0.03	−5.82	−2.01 (−2.39)	Reject (Reject)

[a] At the 10% (5%) significance level. Numeraire country is US.

Table 3
Tests for nonstationarity of relative per capita government consumption levels using panel data and Levin–Lin (1992)

Model		ρ_0-1 estimate	t-statistic	10% (5%) Critical t	Null hypothesis of nonstationarity[a]
1a.	$\Delta y_{it} = (\rho_0-1)y_{it-1} + \varepsilon_{it}$	−0.01	−2.93	−1.44 (−1.81)	Reject (Reject)
1b.	$\Delta y_{it} = (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \varepsilon_{it}$	−0.02	−9.08	−1.44 (−1.81)	Reject (Reject)
1c.	$\Delta y_{it} = (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \rho_2\Delta y_{it-2} + \varepsilon_{it}$	−0.02	−8.82	−1.44 (−1.81)	Reject (Reject)
2a.	$\Delta y_{it} = \delta_0 + (\rho_0-1)y_{it-1} + \varepsilon_{it}$	−0.01	−1.27	−1.72 (−2.09)	Not Reject (Not Reject)
2b.	$\Delta y_{it} = \delta_0 + (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \varepsilon_{it}$	−0.02	−4.38	−1.72 (−2.09)	Reject (Reject)
2c.	$\Delta y_{it} = \delta_0 + (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \rho_2\Delta y_{it-2} + \varepsilon_{it}$	−0.02	−4.38	−1.72 (−2.09)	Reject (Reject)
3a.	$\Delta y_{it} = \delta_0 + \delta_1 t + (\rho_0-1)y_{it-1} + \varepsilon_{it}$	−0.03	−3.44	−2.01 (−2.39)	Reject (Reject)
3b.	$\Delta y_{it} = \delta_0 + \delta_1 t + (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \varepsilon_{it}$	−0.01	−2.95	−2.01 (−2.39)	Reject (Reject)
3c.	$\Delta y_{it} = \delta_0 + \delta_1 t + (\rho_0-1)y_{it-1} + \rho_1\Delta y_{it-1} + \rho_2\Delta y_{it-2} + \varepsilon_{it}$	−0.01	−2.91	−2.01 (−2.39)	Reject (Reject)

[a] At the 10% (5%) significance level. Numeraire country is US.

be rejected in favor of the alternative hypothesis of stationarity for all three variables at the 5% significance level, with or without intercepts and time trends, consistent with the three studies cited earlier. Consequently, we are able to estimate specifications (20) and (21) in level form (as well as structural equations later) and compare t-statistics with conventional critical values.

Naturally, the use of panel data and the techniques of Levin and Lin to evaluate the time-series properties of variables has not developed without criticism. For instance, Karlsson and Lothgren (2000) provide an excellent methodological evaluation and simulation study of common panel unit-root tests. Given that the alternative hypothesis in the Levin-Lin test is that all series in the panel are stationary, Karlsson and Lothgren argue that the panel-test results can possibly be driven by only a few stationary series. However, they note that this outcome is likely in panels with a large time-series dimension ($T>100$). By contrast, their simulation study shows that for panels with a small time-series dimension (such as here), the “potential risk” is that the whole panel may be erroneously modeled as *nonstationary*, not stationary. Thus, the bias in the panel test is against finding stationarity. Moreover, in our data set, standard unit-root tests for each time series individually yield the usual small-T result that nonstationarity cannot be rejected, due to the usual low-power problem of such tests with a small time-series sample (T). Yet a thorough examination of the results of the individual-country tests reveals that the coefficient estimates are uniform in suggesting stationarity, but the Dickey–Fuller tests suffer conventionally from low power (results available from the authors on request). Our evidence suggests that the rejection of nonstationarity using the panel data is due to the power of the panel, rather than only a few stationary series driving the results.[13]

4.3. Parameter identification issues

An econometric issue not yet addressed is the identification of the parameters. Eqs. (20) and (21) are reduced forms from a system of equations that are *overidentified*.[14] Consequently, indirect least squares cannot be used efficiently to identify the parameter estimates, but two-stage least squares (2SLS) can be used. Two-stage least squares estimation of:

[13] As a further indication of the stationarity of our variables and that our results are not driven by only a few stationary series, we also conducted the stationarity tests for sub-samples of the panel. For instance, when the individual-country tests indicated only a few stationary series, we excluded these from the panel. The resulting panel tests using the remaining (small-T) time series still indicated stationarity, suggesting that our panel test results were indeed benefitting from increased power. In 90 individual Dickey–Fuller and ADF tests, 89 coefficient estimates of the lagged value of the log-level of the variable had the (negative) sign consistent with stationarity.

[14] For instance, dividing the coefficient for country 1’s time trend in Eq. (20), $\{[\gamma/(\gamma+\sigma)][\ln(\beta_1/\beta)]\}$, by the coefficient for this time trend in Eq. (21), $\{[1/(\gamma+\sigma)][\ln(\beta_1/\beta)]\}$, yields identification of γ. However, dividing the coefficient for country 2’s (or 3’s, etc.) time trend in Eq. (20), $\{[\gamma/(\gamma+\sigma)][\ln(\beta_2/\beta)]\}$, by the coefficient for this time trend in Eq. (21), $\{[1/(\gamma+\sigma)][\ln(\beta_2/\beta)]\}$, yields identification of γ also. Thus, using indirect least squares with the reduced forms cannot identify uniquely γ or any of the 10 country pairings’ relative discount rates.

$$\ln x_{it} = \sum_{j=1}^{10} \Phi_j^3 = \gamma\ln(c_{it}/c_t) - \gamma\ln[(y_{it}^{N} - g_{it}^{N})/(y_t^{N} - g_t^{N})] + \varepsilon_{it}^3, \tag{22}$$

$$\ln x_{it} = \sum_{j=1}^{10} \Phi_j^4 + \sum_{j=1}^{10} [\ln(\beta_j/\beta)]\text{trend}_{jt} - \sigma\ln(c_{it}/c_t) + \delta\ln(g_{it}/g_t) + \varepsilon_{it}^4, \tag{23}$$

can yield unique parameter estimates of the relative share of tradables in utility ($1/[1+\gamma]$), the relative discount rates (β_j/β), the elasticity of intertemporal substitution ($1/\sigma$), and degree of complementarity or substitutability of government consumption for private consumption (δ). Conventional 2SLS is applied. In the first stage, we estimate the reduced-form Eq. (21) to obtain the predicted values of $\ln(c_{it}/c_t)$. In the second stage, the predicted values of $\ln(c_{it}/c_t)$, denoted $\ln[pv(c_{it}/c_t)]$, are used as an "instrument" for $\ln(c_{it}/c_t)$. Hence, regressions (22) and (23) will use the predicted values of $\ln(c_{it}/c_t)$from the first stage.

Finally, OLS estimation of these equations (reduced-forms or second-stage of 2SLS) will lead to consistent estimates. However, in the presence of potential autocorrelation and/or heteroskedasticity of the error terms, the standard errors of the coefficient estimates may be inefficient. To account for this, we employ the heteroskedasticity- and autocorrelation-consistent covariance estimator of Newey and West (1987).

5. Empirical results

In this section, we present first the results of estimating reduced-form Eqs. (20) and (21). The estimate of reduced-form Eq. (21) generates the instrument for estimating structural Eqs. (22) and (23). The panel included 20 time-series observations (1971–1990) for each of 10 country pairs.

5.1. Reduced-form equations

Estimation of Eqs. (20) and (21) yields:

$$\begin{aligned}\ln x_{it} = &\underset{(-2.54)}{-0.7924} \underset{(-2.78)}{-1.2741}\ln(pv[(y_{it}^{N}-g_{it}^{N})/(y_t^{N}-g_t^{N})]) + \underset{(1.44)}{0.2898\ln(g_{it}/g_t)} \\ &+ \underset{(2.23)}{0.0139 \mathit{TrendAus}_{it}} - \underset{(0.33)}{0.0018 \mathit{TrendCan}_{it}} - \underset{(0.36)}{0.0026 \mathit{TrendDen}_{it}} \\ &+ \underset{(1.54)}{0.0154 \mathit{TrendFin}_{it}} + \underset{(1.18)}{0.0077 \mathit{TrendFra}_{it}} - \underset{(1.63)}{0.0119 \mathit{TrendGre}_{it}} \\ &+ \underset{(2.70)}{0.0170 \mathit{TrendIta}_{it}} + \underset{(4.80)}{0.0417 \mathit{TrendJpn}_{it}} + \underset{(1.31)}{0.0119 \mathit{TrendNor}_{it}}\end{aligned} \tag{24}$$

$$
\begin{aligned}
&+\underset{(2.48)}{0.0163}TrendUK_{it}+\underset{(2.72)}{0.3703}DVCan_{it}-\underset{(-0.58)}{0.1045}\ DVDen_{it}\\
&-\underset{(-1.07)}{0.2353}\ DVFin_{it}-\underset{(-0.43)}{0.0603}\ DVFra_{it}-\underset{(-1.77)}{0.8427}\ DVGre_{it}-\underset{(2.50)}{0.5430}DVIta_{it}\\
&-\underset{(-0.16)}{0.0276}\ DVJpn_{it}-\underset{(-1.34)}{0.3948}\ DVNor_{it}-\underset{(-2.91)}{0.3899}\ DVUK_{it}
\end{aligned}
$$

$R^2 = 0.68$, $AdjR^2 = 0.64$, $S.E.E. = 0.144$, $n = 200$

$$
\begin{aligned}
\ln(c_{it}/c_t) &= \underset{(-1.49)}{-0.0846}+\underset{(8.25)}{0.6244}\ln(pv[(y^{N}_{it}-g^{N}_{it})/(y^{N}_{t}-g^{N}_{t})])\\
&+\underset{(5.08)}{0.2164}\ln(g_{it}/g_t)-\underset{(1.79)}{0.0018}TrendAus_{it}-\underset{(-0.22)}{0.0003}\ TrendCan_{it}\\
&-\underset{(-7.47)}{0.0073}\ TrendDen_{it}+\underset{(0.09)}{0.0002}TrendFin_{it}-\underset{(-4.12)}{0.0041}\ TrendFra_{it}\\
&-\underset{(-0.28)}{0.0006}\ TrendGre_{it}+\underset{(4.64)}{0.0052}TrendIta_{it}+\underset{(0.57)}{0.0008}TrendJpn_{it}\\
&-\underset{(-1.75)}{0.0062}\ TrendNor_{it}+\underset{(3.18)}{0.0036}TrendUK_{it}+\underset{(4.82)}{0.1473}DVCan_{it}\\
&+\underset{(7.34)}{0.1970}DVDen_{it}+\underset{(2.55)}{0.0952}DVFin_{it}+\underset{(11.14)}{0.2158}\,DVFra_{it}\\
&+\underset{(2.70)}{0.2310}DVGre_{it}+\underset{(5.94)}{02140}DVIta_{it}+\underset{(1.81)}{0.0517}DVJpn_{it}+\underset{(4.71)}{0.2624}DVNor_{it}\\
&+\underset{(1.98)}{0.0510}DVUK_{it}
\end{aligned}
\tag{25}
$$

$R^2 = 0.99$, $AdjR^2 = 0.98$, $S.E.E. = 0.029$, $n = 200$

The values of t-statistics are in parentheses and S.E.E. is the standard error of the regression.

The results indicate that relative per capita nontradables output has the expected negative relationship with the real exchange rate in (24) and the expected positive relationship with relative per capita consumption in (25). Relative per capita government consumption has a positive relationship with both the real exchange rate and relative per capita private consumption, suggesting that government consumption *complements* private consumption via the consumption-tilting channel. However, as discussed earlier, indirect least squares cannot identify uniquely structural parameters because of overidentification. Eq. (25) is used in the first stage of 2SLS to generate the instrument, denoted $\ln[pv(c_{it}/c_t)]$, used in the next section.

5.2. Structural equations

Two-stage least squares estimation of Eqs. (22) and (23)—with the restriction that the slope coefficient estimates in intratemporal equilibrium condition (22) be equal but oppositely signed (consistent with the model)—yields:

$$\begin{aligned}
\ln x_{it} = {} & \underset{(-2.00)}{-0.1108} + \underset{(2.38)}{1.2360}\ln[pv(c_{it}/c_t)] - \underset{(-2.38)}{1.2360}\ \ln(pv[(y^{\mathrm{N}}_{it}-g^{\mathrm{N}}_{it})/(y^{\mathrm{N}}_{t}-g^{\mathrm{N}}_{t})]) \\
& - \underset{(-1.56)}{0.108}\ DVCan_{it} - \underset{(-0.80)}{0.1433}\ DVDen_{it} - \underset{(-0.39)}{0.0521}\ DVFin_{it} \\
& - \underset{(-1.34)}{0.1966}\ DVFra_{it} - \underset{(3.48)}{0.7349}DVGre_{it} - \underset{(-3.01)}{0.5518}\ DVIta_{it} + \underset{(3.24)}{0.2452}DVJpn_{it} \\
& - \underset{(-0.86)}{0.2055}\ DVNor_{it} - \underset{(-3.18)}{0.3444}\ DVUK_{it}
\end{aligned} \tag{26}$$

$R^2 = 0.57$, $AdjR^2 = 0.54$, $S.E.E. = 0.163$, $n = 200$

$$\begin{aligned}
\ln x_{it} = {} & \underset{(-2.61)}{-0.9652} - \underset{(-2.78)}{2.0406}\ \ln[pv(c_{it}/c_t)] + \underset{(3.12)}{0.7315}\ln(g_{it}/g_t) \\
& + \underset{(2.72)}{0.0175}TrendAus_{it} + \underset{(0.19)}{0.0012}TrendCan_{it} \\
& - \underset{(-1.73)}{0.0122}\ TrendDen_{it} + \underset{(1.94)}{0.0157}TrendFin_{it} + \underset{(-0.13)}{0.0007}\ TrendFra_{it} \\
& - \underset{(-1.68)}{0.0130}\ TrendGre_{it} + \underset{(3.12)}{0.0276}TransIta_{it} \\
& + \underset{(4.76)}{0.0434}TrendJpn_{it} - \underset{(-0.10)}{0.0007}\ TrendNor_{it} + \underset{(3.00)}{0.0238}TrendUK_{it} \\
& + \underset{(3.05)}{0.6709}DVCan_{it} + \underset{(1.78)}{0.2975}DVDen_{it} - \underset{(-0.25)}{0.0411}\ DVFin_{it} \\
& + \underset{(2.74)}{0.3801}DVFra_{it} - \underset{(-1.14)}{0.3713}\ DVGre_{it} - \underset{(-0.98)}{0.1064}\ DVIta_{it} + \underset{(0.50)}{0.0780}DVJpn_{it} \\
& + \underset{(1.05)}{0.1408}DVNor_{it} - \underset{(-2.27)}{0.2858}\ DVUK_{it}
\end{aligned} \tag{27}$$

The values of t-statistics are in parentheses.

The 2SLS estimation of the intratemporal and intertemporal equilibrium conditions suggests the following inferences regarding government expenditures, in the context of the model. Eq. (26) implies that an increase in home government expenditures on nontradables (g_t^N) creates excess demand for nontradables. This has the effect of withdrawing resources from private nontradables consumption and raising the relative price of (private consumption) nontradables to tradables, causing a real appreciation of the home currency (fall in x_{it}); this is the resource-withdrawal effect. Additionally, Eq. (27) implies that an increase in home per capita government expenditure raises the marginal utility of home private consumption, raising both home per capita private consumption and the home relative price of nontradables to tradables, causing a real appreciation of the home currency. The positive estimate of δ suggests that government and private consumption are *complements*; the estimate of δ (0.7315) is statistically significantly different from zero at the 1% level.

The 2SLS estimation of the intratemporal and intertemporal equilibrium conditions yields the following inferences of the model's other parameters. The estimated coefficient of relative risk aversion (or the inverse of the elasticity of intertemporal substitution), σ, equals 2.04. This estimate is consistent with a priori considerations in the closed-economy macrofinance literature; see Kocherlakota (1996). Most financial economists argue the coefficient of relative risk aversion should be 3 or less, despite finding much higher empirical estimates. Our estimate is statistically significantly different from zero at the 5% level.

The estimate of γ is 1.2; this coefficient is statistically significant at the 5% level. The equation was estimated also with the restriction relaxed. Using a χ^2 statistic, equality of the two coefficients in intratemporal condition (26), allowing for the sign difference, could not be rejected at the 5% significance level ($\chi^2 = 2.28$ compared to critical value $\chi^2[0.95;1] = 3.84$). This estimate of γ implies a share of nontradables in the period Cobb–Douglas utility function of 0.55, which is quite plausible.

The estimated values of relative discount rates, β_i/β, in Eq. (27) range between -1% for Denmark and Greece and 4% for Japan. The estimates suggest that the US rate of time preference was less than that of several other OECD countries. However, the results suggest that the United States was economically and statistically significantly *less* thrifty than Austria, Italy, Japan, and the United Kingdom over the period examined.[15]

[15] As a referee notes, the results may be sensitive to the choice of "denominator country." To evaluate the robustness of the results, we re-estimated the model (including stationarity tests) using Austria as the numeraire country for a sub-sample of only continental European countries, leaving a smaller sample of six European pairs. The results (available on request) generally supported the robustness of the model with the estimates of γ and δ being virtually identical between the full and partial samples, but the estimates of σ being different. In the full sample, the estimate of σ is approximately 2; in the sub-sample of continental European economies, the estimate of σ is approximately 0.4, with these two estimates consistent with recent estimates in Cashin and McDermott (2001).

6. Interpreting the empirical results: relative effects of government expenditure resource withdrawal vs. consumption tilting

The parameter estimates in Eqs. (26) and (27) allow us to estimate the relative importance of the resource-withdrawal vs. consumption-tilting effects of government expenditure on real exchange rates. Consider theoretical real exchange rate Eq. (20). Differentiating Eq. (20) with respect to changes in country i's endogenous variables yields:

$$\mathrm{dln}x_{it} = -[\gamma\sigma(\gamma+\sigma)]\mathrm{dln}(y_{it}^{\mathrm{N}}-g_{it}^{\mathrm{N}}) + [\delta\gamma/(\gamma+\sigma)]\mathrm{dln}(g_{it}).$$

Then, using Eq. (8), this equation becomes:

$$\mathrm{d}x_{it}/x_{it} = [\gamma\sigma/(\gamma+\sigma)][g_{it}^{\mathrm{N}}/c_{it}^{\mathrm{N}}]\mathrm{d}g_{it}^{\mathrm{N}}/g_{it}^{\mathrm{N}} + [\delta\gamma/(\gamma+\sigma)]\mathrm{d}g_{it}/g_{it}. \quad (28)$$

The first term on the RHS represents the resource-withdrawal effect of a 1% increase in country i's government expenditures on nontradables, but the second term on the RHS represents the consumption-tilting effect of a 1% increase in country i's *total* government expenditure. If government per capita consumption expenditure is fully on nontradables, so that $\mathrm{d}g_{it}^{\mathrm{N}} = \mathrm{d}g_{it}$, Eq. (28) becomes:

$$\mathrm{d}x_{it}/x_{it} = [\gamma\sigma/(\gamma+\sigma)][g_{it}^{\mathrm{N}}/c_{it}^{\mathrm{N}}]\mathrm{d}g_{it}^{\mathrm{N}}/g_{it}^{\mathrm{N}} + [\delta\gamma/(\gamma+\sigma)][g_{it}^{\mathrm{N}}/g_{it}]\mathrm{d}g_{it}^{\mathrm{N}}/g_{it}^{\mathrm{N}}. \quad (29)$$

To evaluate the contribution of the first RHS term, an estimate of $g_{it}^{\mathrm{N}}/c_{it}^{\mathrm{N}}$ is needed. To evaluate the second RHS term, an estimate of $g_{it}^{\mathrm{N}}/g_{it}$ is needed. The estimated means of $g_{it}^{\mathrm{N}}/c_{it}^{\mathrm{N}}$ and $g_{it}^{\mathrm{N}}/g_{it}$ would be appropriate; however, annual data on (and, consequently, the mean of) g_{it}^{N} is not available for any country in Penn World Tables Version 5.6a or in OECD *Annual National Accounts* (1996). Fortunately, Kravis et al. (1982) provides estimates of g_{it}^{N} for 7 of the 11 countries in our sample for one year, 1975. Consequently, this allows us to construct $g_{it}^{\mathrm{N}}/c_{it}^{\mathrm{N}}$and $g_{it}^{\mathrm{N}}/g_{it}$ for the year 1975 (near the middle of our time series).

Table 4 (columns 4 and 5) provides estimates of the sizes of the resource-with-

Table 4
Resource-withdrawal (RW) vs. consumption-tilting (CT) effects of a one percent increase in per capita nontradables government consumption expenditures

Country	$g_{1975}^{\mathrm{N}}/c_{1975}^{\mathrm{N}}$	$g_{1975}^{\mathrm{N}}/g_{1975}$	RW effect[a]	CT effect[a]
Austria	0.230	0.767	0.177	0.212
Denmark	0.580	0.827	0.447	0.229
France	0.306	0.654	0.236	0.181
Italy	0.300	0.614	0.231	0.170
Japan	0.132	0.793	0.101	0.220
UK	0.393	0.670	0.303	0.186
US	0.162	0.519	0.125	0.144
Mean	0.300	0.692	0.231	0.192

[a] Effect (in percentage) on the country's real exchange reate of a one percent increase in the country's per capita nontradables government consumption expenditures.

drawal (RW) and consumption-tilting (CT) effects on the real exchange rate of a 1% increase in per capita nontradables government consumption for each of seven of the eleven countries in our sample. Columns (4) and (5) indicate that the two effects are *approximately equal* in size. For some countries, the resource-withdrawal effect is larger than the consumption-tilting effect, and for other countries the reverse holds. The average resource-withdrawal effect for the seven countries is 0.231 and the average consumption-tilting effect is 0.192.

If instead a country's per capita government consumption increase of 1% is proportionate across nontradables *and* tradables, so that $\mathrm{d}g^{\mathrm{N}}_{it}/g^{\mathrm{N}}_{it} = \mathrm{d}g_{it}/g_{it}$ in Eq. (28) , the theoretical effect on the real exchange rate of consumption-tilting increases:

$$\mathrm{d}x_{it}/x_{it} = [\gamma\sigma/(\gamma + \sigma)][g^{\mathrm{N}}_{it}/c^{\mathrm{N}}_{it}]\mathrm{d}g_{it}/g_{it} + [\delta\gamma/(\gamma + \sigma)]\mathrm{d}g_{it}/g_{it}. \tag{30}$$

Table 5 (columns 2 and 3) provides estimates of the sizes of the resource-withdrawal and consumption-tilting effects for each of the seven countries; given the second RHS term in Eq. (30), the consumption-tilting effect is the same (0.277) for all countries for a 1% increase in country *i*'s per capita government consumption. As before, for some countries the resource-withdrawal effect is larger, and for other countries the consumption-tilting effect is larger. However, while also *approximately equal*, the consumption-tilting effect (0.277) is larger than average resource-withdrawal effect (0.231). On average, in the case of proportionate increases in US government tradables and nontradables consumption, a 1% increase in a country's per capita government consumption causes a real appreciation of its currency of roughly 0.5 of 1%; the source of the dollar's real appreciation is estimated to be attributable *approximately equally* to the consumption-tilting and resource-withdrawal channels.

Finally, we attempt to compare briefly the estimated effects of relative government spending increases in our study with some roughly comparable results from previous studies in the literature. First, since no previous study has attempted empirically to isolate the resource-withdrawal effect from the consumption-tilting effect, we cannot

Table 5
Resource-withdrawal (RW) vs. consumption-tilting (CT) effects of a one percent increase in per capita government consumption expenditures

Country	RW effect[a]	CT effect[a]
Austria	0.177	0.277
Denmark	0.447	0.277
France	0.236	0.277
Italy	0.231	0.277
Japan	0.101	0.277
UK	0.303	0.277
US	0.125	0.277
Mean	0.231	0.277

[a] Effect (in percentage) on the country's real exchange rate of a one percent increase in the country's per capita government consumption expenditures.

offer any direct comparisons; all previous studies have assumed that the effect of government expenditure on real exchange rates is attributed entirely to the "supply-side" effect of resource withdrawal. However, we can compare potentially our estimates of the "total" effect to those of other studies. Second, no previous study has attempted to estimate these two effects based upon estimation of structural equations derived from a general equilibrium model. Previous empirical efforts to estimate government expenditure's effect on equilibrium real exchange rates—which are few—typically estimate a reduced-form equation loosely related to the Balassa-Samuelson framework, but not tied to a specific general equilibrium model. Third, most previous studies in this literature have employed *levels* of governments' shares of GDP, and not logarithms; moreover, no study has used relative per capita government consumption expenditure or relative per capita nontradables government consumption expenditure. Thus, a comparison of our study's results to those in the literature is made with caution.

Three representative studies are Penati (1986), Froot and Rogoff (1991), and De Gregorio et al. (1994b). Froot and Rogoff (1991) and De Gregorio et al. (1994b) are studies that examine theoretically and empirically the effect of government expenditures on real exchange rates using panel data sets. Froot and Rogoff (1991) estimated a regression of the log of real exchange rates on the domestic and foreign levels of shares of government consumption expenditures in GDP and productivity levels using a panel of EMS countries for the period 1979–1989 (annual data). Froot and Rogoff (Table 8) found statistically significant coefficient estimates for the difference in two countries' shares of government consumption expenditure in GDP of 1.68 to 2.36. However, while consistent qualitatively with our results, we cannot make a quantitative comparison, since we used the *logarithms* of relative per capita government consumption. De Gregorio et al. (1994b) estimated a panel of first-differenced annual log values of relative prices of nontradables to tradables for 14 OECD countries for the period 1970–1985, regressing this variable in alternative specifications on productivity differentials, government expenditure's share of GDP, per capita income, and inflation rate changes. Like Froot and Rogoff (1991), De Gregorio, Giovannini and Wolf used the level of government expenditure's share, not the log-level. De Gregorio, Giovannini, and Wolf found coefficient estimates for government expenditure's share between 1.50 and 1.97, a range overlapping with the results in Froot and Rogoff (1991).

Penati (1986) estimated several alternative regressions using a panel of six EEC countries over the period 1970–1982, including such RHS variables as government's share of GDP, money supply growth, relative ex post real interest rates, and real oil prices using first-differenced logarithmic annual data. Unlike the previous two studies, Penati used the logarithm of government's share of GDP. Penati found that a one percentage point increase in government's expenditure share of GDP increased the relative price of nontradables to tradables (the real exchange rate) by 0.18–0.35%, depending on the specification; Penati's coefficient estimates were statistically significant. This compares with a slightly larger estimated mean total effect in our Table 5 of 0.51% , although our panel spanned 1970–1990 while Penati's panel spanned only the 1970s. We conclude that the estimates in our study of the "total" impact—

resource-withdrawal and consumption-tilting combined—of a 1% increase in government expenditures on real exchange rates are similar to comparable previous empirical efforts. Moreover, our model suggests that the resource-withdrawal and consumption-tilting effects contribute *approximately equally* to the effect of government consumption expenditure on real exchange rates.

7. Conclusions

The NBER Reporter noted not too long ago that:

> One of the recent areas of resurgent research in open economy macroeconomics has been the examination of real exchange rates. . . . Chinn and Johnston find that government spending and productivity trends help in the analysis of real exchange rates; their finding is confirmed by Canzoneri et al. (1999), and by De Gregorio and Wolf (Rose, 1997, pp. 1–2).

This paper has extended the intertemporal neoclassical framework in Frenkel and Razin (1996), which applies to a small open economy, to examine theoretically and empirically the “resource-withdrawal” vs. “consumption-tilting” effects of government expenditure on real exchange rates in a two-country setting. Our stochastic, dynamic, general equilibrium model illustrated the effect of government expenditure on private consumption decisions through both intratemporal and intertemporal channels.

Empirical evaluation of the model using panel data from the U.N. Income Comparisons Program and OECD (1996) provides evidence that a per capita government expenditure increase may be causing a real appreciation of a country’s currency via resource withdrawal in the medium run. Simultaneously, the same government spending increase may cause the country’s currency to appreciate in real terms because government consumption complements the utility from private consumption. Moreover, these effects are found in the context of plausible parameter estimates of the countries representative consumers’ elasticities of intertemporal substitution, relative rates of time preference, and shares of nontradables in utility. The empirical results suggest that the potential importance of the consumption-tilting channel should not be ignored since this channel has roughly an equal effect on the real exchange rate as the resource-withdrawal channel.

The extant literature on real exchange rates has traditionally perceived the real appreciation of a country’s currency as a cost necessarily incurred to benefit from the provision of public goods, interpreted typically as the opportunity cost of *foregone* private consumption (notably of nontradables private consumption). Our framework and empirical results suggest that the associated real appreciation of a country’s currency in response to increased public consumption can be interpreted partially as the shadow price of a benefit to the representative agent’s utility from the complementarity of government and private consumption. This suggests reconsidering the relative benefits of public expenditures and their implications for explaining departures from purchasing power parity.

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