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I thank my parents for their support during my studies and I dedicate this thesis to them.

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CHAPTER I

INTRODUCTION

Models of the international economy which assume complete asset markets predict that consumption co-moves closely in different countries as this structure of asset markets allows agents in different countries to 'pool' the country-specific risks which they face (see Scheinkman (1984), Leme (1984)). Examples in this class of models include the recent international Real Business Cycle models of, among others, Backus, Kehoe & Kydland (1989), Baxter & Crucini (1989), Stockman & Tesar (1991).

The first essay in this thesis (chapter II) tests the implications for the trend behavior of consumption of models of the international economy which assume complete asset markets. In a world where consumptions and real bilateral real exchange rates of different countries follow unit root processes, such models predict that (under certain assumptions about preferences) consumptions and bilateral real exchange rates are cointegrated for any given pair of countries. The paper presents statistical tests which suggest that data on consumptions and real exchange rates for the US, Japan, France, Britain, Italy and Canada during the period 1971-1987 are inconsistent with this prediction.

These findings suggest that models with limitations on international asset markets might be needed to explain the international covariation of consumption.

The second essay in this thesis (chapter III) presents a Real Business Cycle model in which limitations on international capital markets exist in the sense that only debt contracts are available for international capital flows. Simulations of the model suggest that it can explain the low cross-country correlations observed in detrended consumption data, and that for 'realistic' cross-country correlations of the exogenous shocks.

The second essay argues also that a model which allows for *additive* technology shocks is better able to explain the observed positive correlations of investment and output across countries than standard business cycle theories in which *multiplicative* shocks to total factor productivity are the only source of economic fluctuations. One possible interpretation of the *additive* shocks is as shocks to government consumption.

CHAPTER II

THE STRUCTURE OF INTERNATIONAL CAPITAL MARKETS AND COMMON TRENDS IN INTERNATIONAL CONSUMPTION DATA: AN EMPIRICAL ANALYSIS

1. Introduction

Models of the international economy which assume complete asset markets predict that consumption co-moves closely in different countries as this structure of asset markets allows agents in different countries to 'pool' the country-specific risks which they face (see Scheinkman (1984), Leme (1984)). Examples in this class of models include the recent international Real Business Cycle models of, among others, Backus, Kehoe & Kydland (1989), Baxter & Crucini (1989), Stockman & Tesar (1991).[1]
Previous empirical evaluations of these models have typically focused on their high-frequency implications (see , e.g., Backus, Kehoe & Kydland (1989) who 'calibrate' their model of international business cycles by comparing its predictions to detrended data). In contrast the present paper tests long-run implications of these models, by focusing on their implications for the trend behavior of consumption in different countries.

Unit root tests suggest that consumption in the US, Japan, France, Britain, Italy and Canada during the period 1971-1987 can be described by unit root processes. Under certain assumptions about preferences, models with complete asset markets predict that in a setting where (log) consumptions and bilateral real exchange rates follow unit root processes, these

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[1] See these papers and chapter III for further references.

variables are (stochastically) cointegrated for any given pair of countries, i.e., there exists a linear combination of these series which is trend stationary (see Ogaki (1988)). The paper presents results based on two types of cointegration tests (Park (1991), Phillips & Ouliaris (1990)) which suggest that consumption behavior in the countries mentioned earlier is inconsistent with this prediction.

These results suggest that models with incomplete asset markets might be needed to explain the international covariation of consumption. The paper considers a structure of incomplete asset market in which consumptions in different countries fail to be cointegrated. What makes asset markets incomplete in this economy is the fact that only debt instruments can be used for international borrowing and lending. When rates of change of consumption and of real exchange rates are log-normally distributed and when agents have iso-elastic preferences, this model implies a linear restriction on conditional expectations of future rates of consumption and real exchange rates in different countries. Various tests of this restriction are discussed. Simulations of an international Real Business Cycle model with the asset market incompleteness which was just described suggest that it can explain the low cross-country correlations observed in detrended consumption data (see chapter III). Generalized Method of Moments based tests of a version of the model with incomplete asset markets which assumes a single consumption good fail to reject the model. Tests which are based on a method due to Velu et al. (1986) reject the single good version of the model, but they are more supportive of a version of the model in which countries consume heterogeneous (country-specific) consumption goods.

Section 2 of the paper presents a model of the international economy with complete asset markets and it describes its testable implications. Section 3 describes statistical methods used to test the predictions of that model. Section 4 describes the data used for the tests and section 5 presents the empirical results. Section 6 summarizes the findings obtained in sections 2-5. Section 7 presents a model with an alternative (incomplete) asset markets structure and it tests implications of this asset structure. Tables with summary statistics and empirical results are presented in the appendix.

2. Complete International Asset Markets and Comovements in International Consumption

In the presence of complete asset markets, the behavior of the world economy can be characterized as the solution to a social planning problem which consists in maximizing a weighted sum of the utility levels of the (representative) consumers of the different countries subject to world resource and technology constraints (see , e.g., Scheinkman (1984), Leme (1984), Backus, Kehoe & Kydland (1989), Baxter & Crucini (1989), Stockman & Tesar (1990), Yi (1990)).[2]

Consider a world with 'I' countries indexed by i=1,..,I. Following Backus, Kehoe & Kydland (1989), assume that each country is inhabited by a (representative) agent who is infinitely lived and whose intertemporal preferences can be represented by a time separable utility function of the

[2] Applications of the same idea in a context which involves individual consumers in a closed economy setting can be found in Cochrane (1991), Mace (1991) and Townsend (1989).

form $V_t^i=E_t\sum_{\tau=0}^{\tau=\infty}(\beta^i)^\tau \upsilon_{t+\tau}^i$; here E_t denotes expectations conditional on informations available in period t; $0<\beta^i\equiv 1/(1+b^i)<1$, where b^i is i's subjective rate of time preference and '$\upsilon_{t+\tau}^i$' is the agent's instantaneous utility function in period $t+\tau$.

The social planning problem mentioned earlier consists in maximizing the term $\sum_{i=1}^{i=I}\lambda^i V_0^i$, where λ^i is a 'welfare weight' attached to country i. The welfare weights are time invariant. They reflect the distribution of wealth between the different countries.[3]

2.1. *A World with a Single Consumption Good*

Assume that--as in the Real Business Cycle models of, among others, Backus, Kehoe & Kydland (1989) and Baxter & Crucini (1989)--there exists a unique consumption good which can costlessly be shipped between countries, and let C_t and c_t^i denote total world-wide consumption and the consumption of country 'i' in period t respectively.

Let C_t^* denote the optimal aggregate world consumption implied by the solution of the social planning problem. Given optimal world consumption, the optimal consumptions of countries i=1,...,n can be characterized by solving the following problem:

Max $\sum_{i=1}^{i=n}\lambda^i V_0^i$ by choice of $c_t^1,\ldots,c_t^n$ s.t. $\sum_{i=1}^{i=n}c_t^i\leq C_t^*$ for all periods $t\geq 0$.

[3] To any competitive equilibrium in the world economy with complete asset markets there corresponds a set of welfare weights which is such that the solution of the social planning problem is identical to the competitive equilibrium. Countries whose wealth (evaluated at the prices which obtain in a given competitive equilibrium) is large are given large welfare weights in the social planning problem.

Throughout this paper, it is assumed that the period t utility function of country i is additively separable in consumption and all other goods (denoted by 'Z') which affect i's well-being and that it depends on consumptions in period t only (departures from these assumptions are discussed below):

$$\upsilon_t^i=u^i(c_t^i)+\varphi^i(Z_t^i) \ , \qquad (2.1)$$

where u^i is assumed to be an increasing and concave function.

For this specification of preferences, the solution to the social planning problem requires the following condition to be satisfied:

$$\lambda^i*(\beta^i)^t*u^{i\prime}(c_t^i)=\lambda^j*(\beta^j)^t*u^{j\prime}(c_t^j) \ . \qquad (2.2)$$

(2.2) imposes strong restrictions on the behavior of consumption in countries i and j: it implies that country j's consumption can be expressed as an increasing function of country i's consumption.[4]

To obtain testable implications from (2.2), I follow the Real Business Cycle literature and adopt a constant elasticity specification of the period utility functions:

$$u^k=A^k*(1/\sigma^k)*c^{\sigma^k} \ , \text{ with } A^k>0, \ \sigma^k<1 \quad \text{for } k=i,j \qquad (2.3)$$

With these preferences, condition (2.2) becomes

$$\lambda^i*(\beta^i)^t*A^i*(c_t^i)^{(\sigma^i-1)}=\lambda^j*(\beta^j)^t*A^j*(c_t^j)^{(\sigma^j-1)} . \qquad (2.4)$$

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[4]This follows from the concavity of υ^i and υ^j. Recall that the weights λ^i and λ^j are time invariant.

Taking logs of (2.4) yields the expression

$$(\sigma^i-1)*\ln(c_t^i)=K^{i,j}+\ln(\beta^j/\beta^i)*t+(\sigma^j-1)*\ln(c_t^j), \qquad (2.5)$$

where $K^{i,j}\equiv\ln[(\lambda^j*A^j)/(\lambda^i*A^i)]$.

(2.5) forms the basis for the first set of tests to be presented below.

Statistical tests discussed below suggest that consumption series in the sample of countries considered in this paper can be modeled as unit root processes.

(2.5) implies that there exists a linear combination of $\ln(c^i)$ and $\ln(c^j)$ which *exactly* equals a deterministic trend. From an empirical point of view however it makes more sense, to interpret (2.5) as a long-run 'equilibrium condition' rather than as a condition which each period is exactly satisfied in the data.[5] Given the evidence according to which (log) consumptions follows unit roots processes, I interpret condition (2.5) as predicting that $\ln(c^i)$ and $\ln(c^j)$ are 'stochastically cointegrated', i.e., that there exists a linear combination of these variables which is trend stationary.[6] Ogaki (1988) defines random

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[5]For example because of measurement errors in consumption data. Random taste shocks are another reason why the deterministic restriction on the consumptions of different countries specified in (2.5) may fail to hold. To model these shocks, assume for example that A^i and A^j are random numbers rather than constant parameters, as was assumed up to now. Then $K^{i,j}$ is random. Under the identifying assumption that $K^{i,j}$ is stationary, condition (2.5) predicts that a linear combination of the consumptions of countries i and j is trend stationary.

[6]A variable is trend stationary if it can be represented as the sum of a deterministic trend and a covariance stationary random variable.

variables with unit roots for which there exists a linear combination which is trend stationary as "stochastically cointegrated" . He contrasts this with the concept of "deterministic cointegration" which requires the existence of a linear combination which is covariance stationary.[7]

The tests presented below are tests of the joint hypothesis of complete markets and the preference specifications (2.3). It should however be noted that - up to a (log) linear approximation - linear restrictions similar to (2.5) follow from the risk-sharing condition (2.2) for more general specifications of the instantanous utility function 'u'. Under certain conditions, restrictions on the behavior of consumption in different countries which are simliar to (2.5) can also be obtained if the assumption that life-time utility functions are time-separable is dropped.[8]

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[7]See Campbell & Perron (1991) for further discussions of these concepts.

[8]Assume for example that $u^k_t = u^k(c^k_t + a*c^k_{t-1})$ (where a is a constant), i.e., that the period t instantaneous utility of country k depends on consumptions in periods t and t-1. The solution to the social planning problem now requires the following condition to be satisfied for all periods $t \geq 0$:

$$\lambda^i*\Big((\beta^i)^t*u^{i\prime}(c^i_t+a*c^i_{t-1})+(\beta^i)^{t+1}*a*E_t u^{i\prime}(c^i_{t+1}+a*c^i_t)\Big)=$$
$$\lambda^j*\Big((\beta^j)^t*u^{j\prime}(c^j_t+a*c^j_{t-1})+(\beta^j)^{t+1}*a*E_t u^{j\prime}(c^j_{t+1}+a*c^j_t)\Big).$$

This condition is satisfied if

$$\lambda^i*(\beta^i)^\tau*u^{i\prime}(c^i_\tau+a*c^i_{\tau-1})=\lambda^j*(\beta^j)^\tau*u^{j\prime}(c^j_\tau+a*c^j_{\tau-1}) \qquad \text{(N.1)}$$

holds in all periods.

(to obtain this condition it is important that the parameter 'a' is the same for countries i and j). For the constant elasticity specification (2.3), this condition implies a linear restriction on the behavior of $\ln(c^i_t+a*c^i_{t-1})$ and $\ln(c^j_t+a*c^j_{t-1})$. A more convenient restriction can be

The constant elasticity specification (2.3) is a standard feature of many model in macroeconomics and finance. Existing Real Business Cycle models use it, as it is the only preference specification which - in a model with infinitely lived agents - yields steady state growth paths for which consumption growth rates and real interest rates are constant.

In this context it seems important to note that the Real Business Cycle literature typically allows for preferences in which consumption and work effort interact in a non-separable way; when labor is immobile internationally, non-separabilities between consumption and work effort reduce the international correlation of consumption.[9] This literature assumes however that there exists a steady state level of work effort which

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obtained if instead of a constant elasticity utility function an exponential utility function is used: $u^k(c_t^k+a*c_{t-1}^k)=A^k*\exp(B^k*(c_t^k+a*c_{t-1}^k))$ for k=i,j, where A^k and B^k are constants. Rewriting (N.1) for this utility function and taking logs of the resulting expression yields:

$$B^i*(c_t^i+a*c_{t-1}^i)=\kappa^{i,j}+\ln(\beta^j/\beta^i)*t+B^j*(c_t^j+a*c_{t-1}^j), \tag{N.2}$$

where $\kappa^{i,j}\equiv\ln[(\lambda^j*A^j*B^j)/(\lambda^i*A^j*B^j)]$. Assume that $\{c_t^i\}$ and $\{c_t^j\}$ follow unit root processes, i.e. that $c_t^k=\mu^k+c_{t-1}^k+\varepsilon_t^k$ for k=i,j where ε_t^k is a mean zero variable which is covariance stationary. Substituting c_{t-1}^i and c_{t-1}^j from these expressions into (N.2) yields:

$$B^i*c_{t-1}^i=\hat{\kappa}^{i,j}+B^j*c_{t-1}^j+\ln(\beta^i/\beta^j)*t+\eta_t, \tag{N.3}$$

where $\hat{\kappa}^{i,j}\equiv\kappa^{i,j}+B^i*\mu^i-B^j*\mu^j$ and $\eta_t\equiv B^j*\varepsilon_t^j-B^i*\varepsilon_t^i$ is a covariance stationary random variable. According to (N.3) consumptions in countries i and j are cointegrated.

[9]See Devereux, Gregory & Smith (1991) for detailed discussions of this point.

is constant.[10] Under this assumption, the restriction for the trend behavior of consumption in different countries which are implied by (2.5) continue to hold.[11]

In addition to the setting with a single consumption good, I also consider economies in which countries differ in the consumption goods which they consume. This is motivated by the large variations in real exchange rates experienced by the countries in the sample. Note for example that on average the price of a unit of non-durable Japanese consumption in terms of US consumption has risen at the rate of 5.4% p.a. during the period 1971:1-88:1.[12] Table 2 formally tests the hypothesis that unconditional means of the log growth rates of real exchange rates are zero. For most country pairs in the sample there is strong evidence against this hypothesis.

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[10] See , e.g., King, Plosser & Rebelo (1988).

[11] This can easily be illustrated using the standard utility function used in the Real Business Cycle literature (see, e.g., King, Plossser & Rebelo (1988), Rotemberg & Woodford (1989)): $u=A*(c*\exp[-\psi(L)])^{\sigma}$ where 'L' denotes the number of hours worked and ψ is an increasing function. For this utility function, condition (2.5) becomes:
$(\sigma^{i}-1)*\ln(c_t^{i})-\sigma^{i}*\psi^{i}(L_t^{i}) =K^{i,j}+\ln(\beta^{j}/\beta^{i})*t+(\sigma^{j}-1)*\ln(c_t^{j})-\sigma^{j}*\psi^{j}(L_t^{j})$.
If $\{L_t^{i}\}$ and $\{L_t^{j}\}$ are covariance stationary, then the prediction that $\ln(c_t^{i})$ and $\ln(c_t^{j})$ are stochastically cointegrated continues to hold (actually, the prediction continues to hold provided $\psi^{i}(L_t^{i})$ and $\psi^{j}(L_t^{j})$ that are trend stationary).

[12] Real exchange rates defined in terms of aggregate consumption have behaved similarly during the sample period.

2.2. *A World with one Country-Specific Good*

I next consider a model in which each country consumes a single good which is distinct from the goods consumed by other countries. In what follows, I refer to this framework as the 'one country-specific good model'.[13] In this setting, the assumption of complete asset markets implies that

$$\lambda^{i}*(\beta^{i})^{t}*u^{i,}=\lambda^{j}*(\beta^{j})^{t}*u^{j,}*R_{t}^{i,j} \tag{2.6}$$

holds where $R_t^{i,j}$ is the relative price of the consumption goods of countries i and j (i.e. $R^{i,j}$ is the bilateral real exchange rate of these countries in terms of their respective consumption goods: let p_t^i and p_t^j denote the prices of c_t^i and c_t^j respectively in terms of some numéraire; then $R_t^{i,j} \equiv p^i/p^j$).

To understand why (2.6) is implied by the assumption of complete asset markets, note that the marginal rate of substitution between c_t^i and c_t^j from the point of view of the social planner[14] is $[\lambda^{i}*(\beta^{i})^{t}*u^{i,}]$ / $[\lambda^{j}*(\beta^{j})^{t}*u^{j,}]$. (2.6) is the condition that the planner's marginal rate of substitution between c_t^i and c_t^j is equated to the relative price of these two goods.[15]

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[13] A more realistic specification would assume that each country consumes *two* goods: a tradable good (which is consumed by more than one country) and a non-tradable good ('home' good). The framework with one country-specific consumption good is adopted because available statistical sources do not disaggregate national consumption into home and imported goods.

[14] Recall that the planner's objective function is $\sum_{\tau=0}^{\tau=\infty}\lambda^{i}*V_0^i$.

[15] A more intuitive justification of condition (2.6) is to imagine that there exists a material input (such as oil) which is freely traded between countries and which is used by all countries as an input into the

With the constant elasticity preferences defined in (2.3), condition (2.6) implies that

$$(\sigma^{i}-1)*\ln(c_t^{i})=K^{i,j}+\ln(\beta^{j}/\beta^{i})*t+(\sigma^{j}-1)*\ln(c_t^{j})+\ln(R^{i,j}) \quad , \tag{2.7}$$

where as before $K^{i,j}\equiv\ln[(\lambda^{i}*A^{j})/(\lambda^{i}*A^{i})]$.

As shown below, bilateral real exchange rates in the sample of countries can be described by unit root processes. I therefore interpret (2.7) as predicting that (log) consumptions and (log) bilateral real exchange rates are stochastically cointegrated for these countries.

2.3. *A World with Two Country-Specific Goods*

I also consider an extension of the previous model in which each country consumes two country-specific goods: assume that country i's period t utility function is

$$u_t^{i}=u^{i}(nd_t^{i},s_t^{i}), \tag{2.8}$$

where nd_t^{i} and s_t^{i} are the two goods consumed by i. In the empirical part of the paper these two goods will be interpreted as non-durables and services

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production of their country-specific consumption good. Denote this good by 'z'. At the optimum the social planner is indifferent between allocating an additional unit of the tradable input to country i and allocating that unit to country j. Allocating an additional unit to country k=i,j allows to increase that country's production (and consumption) of its country-specific consumption good by $\partial c^{k}/\partial z$ (the marginal product of the material input). Hence the solution to the social planning problem has the property that

$\lambda^{i}*(\beta^{i})^{t}*(\partial c^{i}/\partial z)*(\partial u^{i}(c^{i})/\partial c^{i})=\lambda^{j}*(\beta^{j})^{t}*(\partial c^{j}/\partial z)*(\partial u^{j}(c^{j})/\partial c^{j})$.

Under perfect competition, $\partial c^{k}/\partial z=1/p^{k}$ holds where p^{k} is the price of country k's consumption good in terms of the material input; we therefore have that $\lambda^{i}*(\beta^{i})^{t}*(\partial u^{i}(c^{i})/\partial c^{i})=\lambda^{j}*(\beta^{j})^{t}*(p^{i}/p^{k})*(\partial u^{j}(c^{j})/\partial c^{j})$, i.e. (2.6) holds.

consumptions respectively. (2.8) is motivated by Stockman & Tesar (1990) who present a Real Business Cycle model which disaggregates private consumption into tradables and non-tradables (services);[16] they argue that this goods markets structure helps the model explain the low cross-country correlations which are observed in detrended consumption data.

In the world with two country specific consumption goods, the solution of the social planning problem satisfies the following conditions:

$$\lambda^i*(\beta^i)^t*u^i_{nd}(nd^i_t,s^i_t)=\lambda^j*(\beta^j)^t*u^j_{nd}(nd^j_t,s^j_t)*RND^{i,j}_t \ , \qquad (2.9\ a)$$

and $$\lambda^i*(\beta^i)^t*u^i_s(nd^i_t,s^i_t)=\lambda^j*(\beta^j)^t*u^j_s(nd^j_t,s^j_t)*RS^{i,j}_t \ , \qquad (2.9\ b)$$

where $RND^{i,j}$ ($RS^{i,j}_t$) is the relative price of the nd (s) goods consumed by countries i and j.

The tests of (2.9 a) and (2.9 b) use a constant elasticity utility function:

$$u^k(nd^k,s^k)=A^k*(nd^k)^{\sigma^k}*(s^k)^{\mu^k} \ . \qquad (2.3')$$

This utility function is increasing and concave in nd and s iff $\sigma+\mu<1$ and $\sigma*\mu>0$ hold.

Writing (2.9 a) and (2.9 b) using the constant elasticity function specified in (2.3') and taking logs yields the following cointegrating relations:

[16] Empirical research in international economics has frequently assumed that services are non-tradable: see for example Kravis, Heston & Summers (1982), Kravis & Lipsey (1983, 1988).

$$(\sigma^i-1)*\ln(nd_t^i)+\mu^i*\ln(s_t^i)= \qquad (2.10\ a)$$
$$K^{i,j}+\ln(\beta^j/\beta^i)*t+(\sigma^j-1)*\ln(nd_t^j)+\mu^j*\ln(s_t^j)+\ln(RND_t^{i,j});$$

$$\sigma^i*\ln(nd_t^i)+(\mu^i-1)*\ln(s_t^i)= \qquad (2.10\ b)$$
$$K^{i,j}+\ln(\beta^j/\beta^i)*t+\sigma^j*\ln(nd_t^j)+(\mu^j-1)*\ln(s_t^j)+\ln(RS_t^{i,j}),$$

where as before $K^{i,j}\equiv\ln[(\lambda^i*A^j)/(\lambda^i*A^i)]$.

Hence the model with two country-specific consumption goods implies that log consumptions of non-durables and services in countries i and j and the bilateral log real exchange rates between these countries are cointegrated.

Empirically, tests of (2.10 a) and (2.10 b) yield very similar results. The discussion in the rest of the paper focuses on condition (2.10 a).

To conclude this section, I restate the cointegrating relations which will be tested below:

(i) The model with a single consumption good:

$$(\sigma^i-1)*\ln(c_t^i)=K^{i,j}+\ln(\beta^j/\beta^i)*t+(\sigma^j-1)*\ln(c_t^j). \qquad (2.5)$$

(ii) The model with one country-specific consumption good:

$$(\sigma^i-1)*\ln(c_t^i)=K^{i,j}+\ln(\beta^j/\beta^i)*t+(\sigma^j-1)*\ln(c_t^j)+\ln(R^{i,j}) . \qquad (2.7)$$

(iii) The model with two country-specific consumption goods:

$$(\sigma^i-1)*\ln(nd_t^i)+\mu^i*\ln(s_t^i)= \qquad (2.10\ a)$$
$$K^{i,j}+\ln(\beta^j/\beta^i)*t+(\sigma^j-1)*\ln(nd_t^j)+\mu^j*\ln(s_t^j)+\ln(RND_t^{i,j}),$$

The next two sections discusses the statistical methods and the data which will be used for the tests of these cointegrating relations.

3. Statistical methods

I assume that all variables can be represented as sums of deterministic trends and mean zero stochastic components as in the following expression:[17]

$$x_t=a+b*t+Z_t \ , \tag{3.1}$$

where Z_t is a mean zero random variable. Following Campbell & Perron (1991), I assume that Z_t follows an ARMA process: $A(L)Z_t=B(L)*\varepsilon_t$, where A(L) and B(L) are polynomials in the lag operator L, while ε_t is iid. The series $\{x_t\}$ has a unit root if the autoregressive polynomial of Z_t has one unit root, while all other roots are strictly outside the unit circle (see Campbell & Perron (1991)). If this condition is satisfied, $(1-L)*x_t$ is covariance stationary with mean b.

3.1. *Unit Root Tests*

Several unit root tests have been discussed in the econometrics literature.[18] The test used here is the Augmented Dickey-Fuller (ADF) test which consists in estimating the following model by OLS:

$$\Delta x_t=\alpha+\beta*t+\phi*x_{t-1}+\textstyle\sum_{s=1}^{k}\varphi_k*\Delta x_{t-k}+\varepsilon_t. \tag{3.2}$$

$\phi=0$ holds under the null hypothesis that the stochastic component of 'Z' in

[17] Campbell & Perron (1991) discuss time series with more general types of deterministic components, including ones characterized by variations in intercepts, slopes etc.

[18] See Campbell & Perron (1991) and Cochrane (1991) for references.

(3.1) follows an ARIMA(k+1,1,0) process.[19] The linear trend $\alpha+\beta^{*}t$ is included in (3.2), because the deterministic part of x_t is assumed to be a linear time trend (see (3.1)).

Under the hypothesis that $\phi=0$, the distribution of the studentized value of the OLS estimate of ϕ is non-standard; critical values for this distribution are tabulated in Fuller (1976).

Results from unit root tests are discussed in section 5.1.

3.2. *Cointegration Tests*

I use two methods for testing for cointegration: the 'spurious regression' test proposed by Park (1990) and the 'residual based' test of Phillips & Ouliaris (1990).

Park's method allows to test the null hypothesis that a set of variables is cointegrated. This is an attractive feature as it allows to directly test the cointegrating relations implied by the models discussed in the last section. In contrast, the Phillips & Ouliaris test (like all other cointegration tests currently available in the econometrics literature)[20] tests the hypothesis that a set of variables is *not* cointegrated. Phillips & Ouliaris tests are reported because failure to reject the no-cointegration hypothesis would provide useful information on the model

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[19] As ARIMA(p,1,q) processes can be approximated by ARIMA(g,1,0) processes for suitable choices of 'g', the Dickey-Fuller testing procedure can also be applied when Z_t has a moving average component (see Said & Dickey (1984)), although using the Dickey-Fuller procedure can be problematic if the MA parameters are large (see Schwert (1987)).

[20] See Campbell & Perron (1991) for an overview.

which is tested.

3.2.1. **Park's (1990) test**[21]

Assume that the q+1 variables x_t^0, $x_t^1, \ldots, x_t^q$ all have unit roots. If they are cointegrated, the residual in the following cointegrating regression is stationary:

$$x_t^0 = D + F*t + \sum_{s=1}^{s=q} \psi_s * x_t^s + \eta_t \quad . \tag{3.3}$$

Park's method considers a variant of this cointegrating regression which obtains when the 'x' variables in (3.3) are transformed by adding to them certain stationary random variables (see Park (1990), p.117). Denote the transformed 'x' variables by $\hat{x}$. Park (1991) refers to the cointegrating regression which obtains when in (3.3), the x variables are replaced by $\hat{x}$ as a 'canonical cointegrating regression'. To apply Park's method, 'superfluous' regressors, such as high order time polynomials,[22] are added to the canonical cointegrating regression and one tests whether these superfluous regressors enter significantly in the regression.

For example, let $z_t = (t^2, t^3, \ldots, t^p)$ and let γ be a column vector of coefficients which is conformable with z_t. To apply Park's method, one can add the term $z_t * \gamma$ to the canonical cointegrating regression:

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[21] See Campbell & Perron (1991) and and Fisher & Park (1990) for useful discussions of this test.

[22] Park (1990) shows that his test can also be performed by adding superfluous random variables with unit roots (such as computer generated random walks) to the cointegrating regression.

$$\hat{x}_t^0 = D + F*t + \sum_{s=1}^{s=n} \psi_s * \hat{x}_t^s + z_t * \gamma + \hat{\eta}_t \tag{3.4}$$

Park's test exploits the fact that if x_t^0, x_t^1,..., x_t^q are stochastically cointegrated, then the Wald test statistic for the hypothesis $\gamma=0$ converges to a stable limiting distribution. Otherwise, the Wald statistic diverges. The Park test statistic for the test of the null of cointegration is a transformation of the Wald test statistic.

As the outcome of the Park test can depend on which of the 'x' variables is used on the left-hand side of the cointegrating regression (3.3), test results will be reported for all possible choices for the left-hand side variable in that equation.

Park (1991) shows how to conduct tests of hypotheses concerning the coefficients of the cointegrating regression. This is useful, as the coefficients of the cointegrating relations (2.5), (2.7) and (2.10 a), are functions of preference parameters. Park's method therefore allows to test whether the preference parameters implied by the estimated coefficients of the cointegrating regressions are consistent with well-behaved utility functions.

To compute Park's test statistics for the test of the null of cointegration, it is necessary to correct for serial correlation in the residual in the cointegrating relation (3.3) and in the first differences of the 'x' variables included on the right-hand side of (3.3) (see Park (1990), p.117) . I use the Newey & West (1987) method and 10 autocorrelations for that purpose.[23]

The Park tests reported below use t^2, t^3 and t^4 as 'superfluous' regressors in (3.4).

3.2.2. **The Phillips & Ouliaris (1990) method**

The fact which underlies the 'residuals based' method is that if the n+1 variables $x_t^0, x_t^1, \ldots, x_t^q$ all have unit roots, and if they are *not* stochastically cointegrated, then the residual in the cointegraing regression (3.3) is non-stationary; hence unit root tests can be applied to that regression residual in order to test the null hypothesis that $x_t^0, x_t^1, \ldots, x_t^q$ are *not* cointegrated. Phillips & Ouliaris (1990) use Phillips' (1987) $\hat{Z}_\alpha$ and $\hat{Z}_t$ unit root test statistics for this purpose. In the work presented below, I use both of these tests.

To compute the $\hat{Z}_t$ and $\hat{Z}_\alpha$ test statistics, one runs the regression: $\hat{\eta}_t=\alpha^*\hat{\eta}_{t-1}+\hat{k}_t$, where $\hat{\eta}_t$ is the regression residual obtained by fitting (3.3). Under the null hypothesis that the variables x_t^0, $x_t^1, \ldots, x_t^n$ are not cointegrated, $\alpha=1$ holds. The $\hat{Z}_\alpha$ test statistic is a transformation of the expression $T^*(\hat{\alpha}-1)$ (where $\hat{\alpha}$ is the OLS estimate of α and T is the sample length) while the $\hat{Z}_t$ test is a transformation of the t-test statistic for a test of the hypothesis that $\alpha=1$.

To compute the Phillips & Ouliaris test statistics, it is necessary to correct for serial correlation in the residual $\hat{k}_t$ (see p.171 in Phillips

[23]Table 7 reports the results of Monte Carlo simulations which suggest that for the available sample length (69 periods) it might be appropriate to correct for 10 autocorrelations.

& Ouliaris (1990)). I use the Newey & West method and 10 autocorrelations for this purpose.

As outcomes of the Phillips & Ouliaris test can depend on which 'x' variable is used on the left-hand side in the cointegrating regression (3.3), test results will be reported for all possible choices for the left-hand side variable in that equation.

4. The Data

This paper uses quarterly data on private consumption in the US, Japan, France, Britain, Italy and Canada (henceforth I refer to these countries as 'G6' countries)[24] from the OECD Quarterly National Accounts (QNA) database. The consumption series for the US, France, Italy and Canada are supplied in seasonally adjusted form by the QNA. Series for Japan and the UK are provided in seasonally unadjusted form; I used the 'esmooth' command in the econometrics package RATS to seasonally adjust them. All consumption figures used in the empirical work are expressed in per capita terms.[25] The consumption data are available for the period 1970:1-88:1 (71:1-88:1 for Italy). The database contains non-durable and services consumption series in current prices (in units of the respective national currencies) as well as in constant prices; by dividing the former by the latter, I constructed price indices for non-durables and for services. Exchange rate data are

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[24]The G6 equals the more familiar G7 without Germany. Germany is not included in the sample because the OECD quarterly national accounts database does not provide data on German consumption of non-durables and services.

[25]Using population figures from the International Financial Statistics database.

taken from the International Financial Statistics database.
Descriptive statistics on consumption and real exchange rates for the G6 countries are presented in tables 1-4.

I tested the cointegrating relations implied by the single consumption good model (see 2.5) and by the model with one country-specific consumption good (see 2.7) for a variety of consumption measures: (i) non-durables, (ii) non-durables plus services, (iii) non-durables plus services and government consumption, (iv) total private consumption expenditures (including expenditures on durables), (v) total private consumption expenditures plus government consumption expenditures.
It appears that the tests results using these different consumption measures are quite similar.
For the tests of the single good model and of the model with one country-specific good, results are reported below for a consumption measures consisting of non-durables plus services, i.e. the cointegrating relations (2.5) and (2.7) are tested using the sum of non-durables and services consumption in country i as a measure for the consumption good 'c'. The tests of the model with two country-specific consumption goods identify these two goods with non-durables and services respectively.

5. Empirical Results

5.1. *Results From Unit Root Tests*

Tables 5 and 6 present the results of unit root tests for consumptions and bilateral real exchange rates. These tests were conducted for the following values of the lag parameter 'k' in (3.2): k=0,1,2,3,4,5,6.

Table 5 presents unit root tests for consumption. For countries other than Japan we see that there is little evidence at the 10% level against the unit root hypothesis. For the Japanese series, the table provides strong evidence against the unit root hypothesis; it appears however that for larger values of k there is little evidence against the unit root hypothesis. Because of this, I do not exclude Japan from the sample.

Table 6 presents tests of the hypothesis that the bilateral real exchange rates (in logs) of the G6 countries follow unit root processes. The tests are for real exchange rates in terms of non-durable consumption goods (test results for real exchange rates in terms of services are similar and are therefore not been presented in the appendix). Even at the 50% level, there is little evidence against the hypothesis that bilateral real exchange rates follow unit root processes.

5.2 *Results From Cointegration Tests*

5.2.1. **Park tests**

Table 8 reports results for the Park test.

For each country pair, table 8 reports p-values of 2 Park test statistics for the single good model (corresponding to two different choices for the left-hand side variable in the cointegrating regression (3.3)) and hence there is a total of 30 test statistics for the single good model.[26] 19 of the 30 test statistics reject the null of cointegration at the 20% level and 12 test statistics reject the null at the 10% level.

Concavity of the utility functions of countries i and j implies that $(\sigma^i-1)/(\sigma^j-1)>0$. With few exceptions, this inequality is satisfied by the data.[27]

The test results threfore cast doubt on the single good model.[28] They support the findings of Neusser (1991), who argues that a consumption measure consisting of the sum of private and government consumption fails to be deterministically cointegrated for Austria, Canada, Germany, Japan, Britain and the US.[29]

Table 8 also reports test results for the model with one country-specific consumption good. A total of 45 test statistics is now reported (for each country pair, three test statistics are computed). The cointegrating

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[26]There are 15 country pairs in the sample.

[27]A '(P)' in the table indicates a statistically significant (at the 5% level) violation of the condition $(\sigma^i-1)/(\sigma^j-1)>0$.

[28]A potential qualification of this suggested conclusion from Park's test is the fact that the for many country pairs the choice of the left-hand side variable in the the cointegrating regression affects the outcome of the test: there are 9 country pairs where one of the two test statistics rejects the null of cointegration (at the 20% level), while the the other test statistic fails to reject it. The dependence of the test outcome on the choice of the left-hand side variable is somewhat worrisome because asymptotically that choice does not affect the distribution of Park's test statistic (the asymptotic distribution is used to calculate the p-values reported in table 8).

[29]Neusser uses a cointegration test due to Johansen (1989).

relations implied by the model with one country-specific consumption good are rejected (at the 20% level) by 30 of the 45 test statistics (at the 10% level, there are 20 rejections).

The preference parameters recovered from the cointegrating regressions frequently violate concavity. If we eliminate the cases where for at least one of the countries included in a given country pair, we reject the hypothesis that $\sigma \leq 1$ at the 5% level, we are left with 12 cases where at the 10% level the Park test fails to reject the null of cointegration.

Finally, Table 8 reports test results for the model with two country-specific consumption goods. For each country pair, five test statistics are now computed (corresponding to five different choices of the left-hand side variable in (3.3)) which gives a total of 75 test statistics. At the 20% level, the null-hypothesis of cointegration is rejected for 40 of the 75 test statistics. At the 10% level there are 20 rejections. Statistically significant violations of the conditions $\sigma^i+\mu^i<1$, $\sigma^i*\mu^i>0$, $\sigma^j+\mu^j<1$, $\sigma^j*\mu^j>0$ [30] occur in approximately two-thirds of the 75 cases considered in table 8 for the model with two country-specific consumption goods.

[30] $\sigma+\mu<1$ and $\sigma*\mu>0$ are necessary and sufficient conditions under which the utility function in the model with two country-specific goods is concave and increasing in both goods.

5.2.2. **Phillips & Ouliaris tests**

Phillips & Ouliaris test results are in table 10.

The Phillips & Ouliaris method tests the null hypothesis that a set of variables is *not* cointegrated. The test results reported in table 10 are consistent with the hypothesis that the three cointegrating relations derived in section 2 do *not* hold: at the 20% significance level, approximately 80% of the test statistics reported in table 10 for the cointegrating relations (2.5), (2.7) and (2.10 a) fail to reject the hypothesis that the cointegrating relations stated at the end of section 1 do *not* hold.[31]

6. Summary of Empirical Results

Cointegration tests were used to test a model of the international economy which assumes complete asset markets. A version of that model which assumes a single consumption good and iso-elastic utility functions predicts that log consumptions in different countries are cointegrated.

In a world in which each country consumes a country specific consumption good, the log consumptions of different countries and their log bilateral real exchange rates are predicted to be stochastically cointegrated. Finally, a model was considered in which each country consumes two

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[31]The proportions of test statistics which yield rejections at the 20% level differ somewhat for the three cointegrating relations. The proportions of rejections by the $\hat{Z}_{\alpha}$ statistic of the null hypothesis of no cointegration are (at the 20% level): 0.20, 0.11 and 0.18 for the single good model, the model with one country-specific good and the model with two country-specific goods respectively. The corresponding proportions of rejections yielded by the Z_t statistic are 0.23, 0.13 and 0.28 respectively.

country-specific goods. In such a setting, a model with complete asset markets implies that (with isoelastic preferences) the consumption of the two country-specific goods by different countries and their bilateral real exchange rates are stochastically cointegrated.

Tests using methods developed by Park (1990) and Phillips & Ouliaris (1990) cast strong doubt on these predictions.

The results of this paper suggest that the failure of the single good model cannot be explained (at least not by a model with complete asset markets) by the large variations in real exchange rates which occured during the sample period.

While rejections of the implications of the complete asset markets model tested in this paper could be due to the fact that preferences or other aspects of the model are misspecified, the rejections cast strong doubt on the international Real Business Cycle model, as the preference specification and other features of the model tested in this paper are the ones commonly used in RBC models.

The findings presented in this section suggest that models with limitations on asset markets (see chapter III and Conze & Scheinkman(1991)) might be needed to gain a better understanding of the international covariation of consumption.

7. An Economy in Which Only Debt Contracts Can Be Used For International Borrowing and Lending

Chapter III of this thesis considers an economy with incomplete asset markets, in which only debt contracts can be used in international asset markets. As will be shown below, such a structure implies (in general) that consumption fails to be cointegrated in different countries.

Section 7.1. discusses testable implications for the model and it describes the statistical methods which will be used to test these implications. 7.2 presents the empirical results. Conclusion for section 7 are in 7.3.

7.1. *Testable Implications of the Debt Model*

7.1.1. A World With a Single Good

A one good world is considered first. The asset market structure is assumed to be the same as in chapter III and I initially assume a world with a single good. In period t, the only type of transaction between different countries consists in unconditional borrowing and lending at the real rate r_t using one period bonds: if country i makes a loan of A_t^i units of the consumption good in period t, then that country gets back $(1+r_t)*A_t^i$ units in period t+1.

Hence i's budget constraint in period t is:

$$c_t^i+I_t^i+A_t^i=(1+r_{t-1})*A_{t-1}^i+y_t^i, \tag{7.1}$$

where y_t^i is i's (gross) output in period t, while I_t^i is its net investment in physical capital between periods t and t+1.

The following constraint is imposed in order to rule out Ponzi schemes:

$$-Z\le A_t^i \quad \text{for all } t, \tag{7.2}$$

where Z is a large positive number.

Optimal behavior of country k implies that the following Euler condition is satisfied:

$$E_t\beta^k*(1+r_t)\{u_c^k(c_{t+1}^k)/u_c^k(c_t^k)\}=1 \ . \tag{7.3}$$

Hence expected intertemporal marginal rates of substitution are equated between countries:

$$\beta^i*E_t\{u_c^i(c_{t+1}^i)/u_c^i(c_t^i)\}=\beta^j*E_t\{u_c^j(c_{t+1}^j)/u_c^j(c_t^j)\} \quad \text{for } i\neq j. \tag{7.4}$$

For the iso-elastic utility function specified in (2.3), condition (7.4) can be restated as:

$$\beta^i*E_t\{(c_{t+1}^i/c_t^i)^{\sigma^i-1}\}=\beta^j*E_t\{(c_{t+1}^j/c_t^j)^{\sigma^j-1}\} \quad \text{for } i\neq j. \tag{7.4'}$$

In order to facilitate empirical testing of this condition, I follow Obstfeld (1989) and assume that consumption growth rates are jointly log-normally distributed and conditionally homoskedastic:[32] under the stated assumptions it follows from (7.4') that

$$(\sigma^i-1)*E_t\Delta\ln(c_{t+1}^i)=\mu^{i,j}+(\sigma^j-1)*E_t\Delta\ln(c_{t+1}^j) \ , \tag{7.5}$$

where $\mu^{i,j}$ is a constant.[33]

Hence we see that with log-normal and conditionally homoskedastic consumption growth and isoelastic utility functions, the condition that expected intertemporal rates of substitution are equated between countries implies that conditional expectations of growth rates of consumption between periods t and t+1 are perfectly correlated.

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[32] See also Hansen & Singleton (1983) for an example of the use of the assumptions of log-normality in the analysis of consumers' Euler equations.

[33] $\mu^{i,j}$ depends on the rates of time preference of countries i and j and on variances of consumption growth in these countries.

Note that condition (7.4) holds in any single good model in which there exists a riskless asset which can freely be traded by the residents of different countries. Hence it also holds with complete asset markets.

The key difference between incomplete and complete asset markets is that in the latter intertemporal marginal rates of substitution are equated ex post as well (and not merely in expected value), which implies that with isoelastic preferences, ex post growth rates of consumption are perfectly positively correlated between countries.[34] In general, consumption growth rates fail to be perfectly correlated between countries when asset markets are incomplete: it follows from (7.5) that

$$(\sigma^{i}-1)*\Delta\ln(c^{i}_{t+1})=\mu^{i,j}+(\sigma^{j}-1)*\Delta\ln(c^{j}_{t+1})+\eta_{t+1}, \tag{7.5'}$$

where $\eta_{t+1}\equiv(\sigma^{i}-1)*\varepsilon^{i}_{t+1}-(\sigma^{j}-1)*\varepsilon^{j}_{t+1}$, with $\varepsilon^{k}_{t+1}=\Delta\ln(c^{k}_{t+1})-E_t\Delta\ln(c^{k}_{t+1})$ for $k=i,j$. η_{t+1} is a linear combination of the forecast errors made in forecasting consumption growth between periods t and t+1 and therefore η_{t+1} is a serially uncorrelated random variable with mean zero.

Taking partial sums of (7.5') for periods $t=1,2,\ldots,T$, we get

$$(\sigma^{i}-1)*\ln(c^{i}_{T})=\vartheta+\mu^{i,j}*T+(\sigma^{j}-1)*\ln(c^{j}_{T})+H_T , \tag{7.5''}$$

where $H_T\equiv\sum_{t=1}^{t=T}\eta_t$ and $\vartheta\equiv(\sigma^{i}-1)*\ln(c^{i}_{0})-(\sigma^{j}-1)*\ln(c^{j}_{0})$. In general, η_T is non-zero when asset markets are incomplete because nothing guarantees that in the absence of complete asset markets, unexpected consumption growth

[34] To see why this is so, note that taking first differences of (2.5) yields: $(\sigma^{i}-1)*\Delta\ln(c^{i}_{t})=\ln(\beta^{i}/\beta^{j})+(\sigma^{j}-1)*\Delta\ln(c^{j}_{t})$.

rates are perfectly correlated across countries [35] which implies that H_T follows a random walk and hence that log consumptions in countries i and j fail to be cointegrated.

7.1.2. **A World With One Country-Specific Good**

I next consider a model in which (as in section 2.2) each country consumes one country specific consumption good which is distinct from the consumption goods consumed by other countries. Assume that the residents of different countries can invest in risk-free bonds which are denominated in the country-specific goods of the different countries in the world. As in section 2, let $R^{i,j}$ denote the price of country i's good in terms of country j's good. Optimal behavior of countries i and j now implies that the following Euler conditions are satisfied:

$$E_t\beta^k*(1+r_t^k)*\{u_c^k(c_{t+1}^k)/u_c^k(c_t^k)\}=1 \quad \text{for } k=i,j \text{ and} \tag{7.6}$$

$$E_t\beta^h*(1+r_t^k)*\{[R_t^{k,h}/R_{t+1}^{k,h}]*u_c^h(c_{t+1}^h)/u_c^h(c_t^h)\}=1 \quad h,k:i,j;\ h\neq k,$$

where r_t^k is the real one period interest rate in terms of country k's good. It follows from (7.6) that for countries i and j:

$$E_t\beta^i*\{u_c^i(c_{t+1}^i)/u_c^i(c_t^i)\}=E_t\beta^j*\{[R_t^{i,j}/R_{t+1}^{i,j}]*u_c^j(c_{t+1}^j)/u_c^j(c_t^j)\}.$$

Assuming iso-elastic preferences, this condition implies that

$$E_t\beta^i*\{(c_{t+1}^i/c_t^i)^{\sigma^i-1}\}=E_t\beta^j*\{[R_t^{i,j}/R_{t+1}^{i,j}]*(c_{t+1}^j/c_t^j)^{\sigma^j-1}\}.$$

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[35] Note however that there exist special (and rather unlikely) circumstances where even with incomplete asset markets, unexpected consumption growth can be perfectly correlated between countries. This is for example the case if the outputs of all countries are perfectly correlated.

Assuming that rates of change of consumption and of real exchange rates are log-normally distributed and conditionally homoskedastic yields an expression of the following form:

$$(\sigma^i-1)*E_t\Delta\ln(c^i_{t+1})=\tilde{\mu}^{i,j}+E_t\Delta\ln(R^{i,j}_{t+1})+(\sigma^j-1)*E_t\Delta\ln(c^j_{t+1}), \tag{7.7}$$

The model with one country-specific consumption good therefore yields a linear restriction involving the conditional expectations of the rates of change of the consumptions of countries i and j and of their real exchange rate.[36]

The restrictions on the rates of change of consumptions and real exchange rates stated in (7.5) and (7.7) can compactly be expressed as:

$$\gamma'*X_{t+1}=\eta_{t+1}; \quad E_t\eta_{t+1}=0 \tag{7.8}$$

where γ is an nx1 vector of coefficients, while X_{t+1} is an nx1 vector containing the log growth rates of the consumptions and real exchange rates of countries i and j; for the model with a single consumption good, we have $X_{t+1}\equiv(\Delta\ln(c^i_{t+1}),\ \Delta\ln(c^j_{t+1}))'$ and $\gamma\equiv((\sigma^i-1),\ -(\sigma^j-1))'$; for the model with one country-specific consumption good, we have $X_{t+1}\equiv(\Delta\ln(c^i_{t+1}),\ \Delta\ln(c^j_{t+1}),\ \Delta\ln(R^{i,j}_{t+1}))'$ and $\gamma\equiv((\sigma^i-1),\ -(\sigma^j-1),\ -1)'$).

[36] Obstfeld (1989) derives (7.7) for a world in which *nominal* risk-free assets denominated in different national currencies are traded between the residents of these countries.

In what follows, Z_t denotes a kx1 (with k>n) vector of instruments which are contained in the period t information set and which therefore are orthogonal to the random variable η_{t+1}:

$$E_t\eta_{t+1}{}^*Z_t=E_t\gamma'{}^*X_{t+1}{}^*Z_t=0 \qquad (7.8')$$

This orthogonality condition implies a restriction on the coefficients in linear regressions of the elements of the vector X_{t+1} on Z_t: consider the regression equation

$$X^m_{t+1}=\nu^m+\psi^{m,}{}^*Z_t+\omega^m_{t+1}\ , \qquad \text{for } m=1,..,n \qquad (7.9)$$

where X^m is the m^{th} element of the vector X and ψ^h is a kx1 vector of coefficients, while ν^m is an intercept.

The n equations stated in (7.9) can be written compactly as:

$$X_{t+1}=\nu+\Psi'{}^*Z_t+\omega_{t+1}, \qquad (7.10)$$

where $\nu\equiv(\nu^1,\dots,\nu^n)'$, $\omega_{t+1}\equiv(\omega^1_{t+1},..,\omega^n_{t+1})'$, while $\Psi\equiv(\psi^1,..,\psi^n)$ is a matrix of dimension kxn.

The orthogonality condition (7.8') implies the following restriction on the matrix Ψ:

$$\Psi^*\gamma=0\ . \qquad (7.11)$$

Hence the model of incomplete asset markets implies that the matrix Ψ is not of full rank:

$$\mathrm{rank}(\Psi)<n. \qquad (7.12)$$

A major difficulty in testing the model with incomplete asset markets is to find good instruments for future consumption growth. Our ability to accurately forecast future realizations of a variable depends on how strong the serial correlation of that variable is (at least when past realizations are used to make the forecasts). Table 4 reports the first 10 autocorrelations for non-durables and services consumption growth in the countries of the sample. While the serial correlation of consumption growth rates is relatively strong in Japan, it is quite weak for the other countries, which suggests that finding good instruments for consumption growth in these other countries might be difficult.

7.2 Test Results

The following tests of the model with incomplete asset markets are considered:[37]

(1.) The single good model implies that expectations of future consumption growth in two countries i and j conditional on a given set of instruments are perfectly correlated.

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[37]See Cumby & Huizinga (1991) for discussions of tests of condition (7.11) for the special case where n=2 (i.e. where the vector X consists of two elements). Alternatives to the tests considered in this paper can be found in recent work by Obstfeld (1989) and Barrionuevo (1991). Obstfeld jointly estimates the following condition (which follows from (7.7)) for a sample of countries which comprises the US, Japan and Germany, using three stage least squares:
$\Delta\ln(R^{i,j}_{t+1})=-\tilde{\mu}^{i,j}+(\sigma^{i}-1)*\Delta\ln(c^{i}_{t+1})-(\sigma^{j}-1)*\Delta\ln(c^{j}_{t+1})+\tilde{\eta}^{i,j}_{t+1}$, where $\tilde{\eta}^{i,j}_{t+1}$ is a linear combination of the forecast errors made in forecasting the first difference of log consumptions and bilateral real exchange rates for countries i and j. Obstfeld argues that quarterly data for the period 1973:1-85:2 support (7.7).
Barrionuevo (1991) uses a Generalized Method of Moments framework to test the Euler conditions (7.6), using consumption and interest rate data for a sample of industrialized countries.

For each country pair i,j, table 11 reports sample correlation coefficents of fitted values of $\Delta\ln(c_t^i)$ and $\Delta\ln(c_t^j)$ which are obtained by regressing these variables on a constant and on lagged consumption growth rates in countries i and j. The table also reports cross-country correlations of actual consumption growth rates. For 11 of the 15 country pairs, the cross-country correlations of fitted consumption growth rates are larger than those of actual consumption growth rates. It appears however that the standard deviations of the cross-country correlations of fitted consumption growth rates are large: for 13 country pairs one fails to reject the hypothesis that the cross-country correlation of conditional expectations of future consumption growth is zero. This most likely reflects the low predictive power of the instruments used to predict future consumption growth in table 11.[38]

(2.) When the set of instruments used to test the orthogonality condition consists of lagged values of X_{t+1} (i.e. when $Z_t=(X'_t,X'_{t-1},\ldots,X'_{t-h})'$) a method due to Velu, Reinsel & Wichern (1986) [39] can be used to test the null-hypothesis that rank(Ψ)=n-1 against the alternative that rank(Ψ)=n.[40]

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[38]For each of the two regressions which are considered in table 11 for a given country pair, the table shows adjusted R^2s. The adjusted R^2s are quite low for many of the reported regressions.
The standard deviations for cross-country correlations of fitted consumption growth rates are calculated using the method presented in Cumby & Huizinga (1991).

[39]See Neusser (1991) for an interesting application of this test.

[40](7.2) requires that rank(Ψ)<n, not necessarily that rank(Ψ)=n-1. Velu et al. actually present n test statistics, which I will denote by

Table 13 presents p-values for tests of the hypothesis that rank(Ψ)=n-1 for the the single good model and for the model with one country-specific consumption good. These tests are conducted for h=0,1,2,3,4 , where h is the maximal lag of X_t used in the set of instruments.

For each model, table 13 therefore reports p-values for 75 Velu et al. test statistics.[41] The results cast doubt on the single good model: at the 10% level, the hypothesis rank(Ψ)=n-1 is rejected by 39 of the 75 Velu et al. test statistics; at the 20% level there are 54 rejections.

Note that if rank(Ψ)=n-1, then--in the single good model--there exists a unique value $(\sigma^i-1)/(\sigma^j-1)$ which satisfies the condition that $\Psi*\gamma=0$, where $\gamma\equiv(\sigma^i-1,\sigma^j-1)$. Velu et al. describe a method for obtaining a regression estimate $\hat{\Psi}$ of Ψ which satisfies the restriction that rank($\hat{\Psi}$)=n-1. Using $\hat{\Psi}*\gamma=0$ then allows to obtain an estimate of $(\sigma^i-1)/(\sigma^j-1)$. Table 12 reports estimates of $(\sigma^i-1)/(\sigma^j-1)$ which were obtained in this way. For the different values of the lag parameter 'h' considered in the table, the condition $(\sigma^i-1)/(\sigma^j-1)>0$ is typically rejected for more than two-thirds of the country pairs. It should however be noted that in most cases, the quantity $(\sigma^i-1)/(\sigma^j-1)$ is not estimated precisely; this most likely

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$v_0,\ldots,v_{n-1}$, where v_i tests the hypothesis that rank(Ψ)=i against the alternative that rank(Ψ)=n. For each country pair and value of h, I calculated these n test statistics. It appears that whenever v_j for j<n-1 rejects the hypothesis (at the 10% or the 20% level) that rank(Ψ)=j, then the statistic v_{n-1} rejects the hypothesis rank(Ψ)=n-1.

[41]Five values of 'h' are considered for 15 country pairs.

reflects the poor predictive power of lagged consumption growth for future consumption growth. In most of the cases where the estimate of $(\sigma^i-1)/(\sigma^j-1)$ is negative, we fail to reject (at conventional significance levels) the hypothesis that this quantity is positive.

Table 13 reports Velu et al. tests for the model with one country-specific good. The test results seem more favorable for the model than the test results obtained for the single good model: at the 10% level, the hypothesis rank(Ψ)=n-1 is rejected by 28 of the 75 Velu et al. test statistics; at the 20% level there are 34 rejections. It seems noteworthy, that most of the rejections (at the 10% level) of the model with one country-specific consumption good are accounted for by 5 country pairs: US-France, US-Italy, Japan-Italy, France-Italy and France-Canada. For the remaining country pairs there are only very few rejections of the restriction rank(Ψ)=n-1. It seems interesting that for the country pair US-Japan (the country pair which experienced the largest change in bilateral real exchange rates during the sample period), the condition rank(Ψ)=n-1 is rejected for the single good model, but not for the model with one country-specific consumption good.

Using an estimate $\hat{\Psi}$ of the matrix Ψ which satisfies the restriction that rank(Ψ)=n-1 allows to obtain estimates of the prefereence parameters σ^i and σ^j for a given country pair. For all country pairs in the sample, table 13 reports estimates of preference parameters which obtain for h=4. Most of the estimated preference parameters are consistent with concave utility functions.

(3.) Tests of the single good model were also conducted using the Generalized Method of Moments,[42] exploiting the orthogonality condition $E\gamma' * X_{t+1} * Z_t = 0$ (see (7.8')). Table 14 presents results of these tests. For a given country pair i,j, the set of instruments used for the test consists of lagged growth rates of non-durables and services consumption and of lagged real interest rates in countries i and j (all instruments are lagged two and three periods).

Hansen's (1982) J-statistic shows that there is no country pair where the orthogonality condition (7.8') is rejected at the 10% significance level.[43]

7.3 *Summary*

This section has considered a model with incomplete asset markets in which only risk-free real debt contracts can be used for international borrowing and lending. In a setting with a single consumption good, this structure of asset markets implies (with iso-elastic preferences and log-normally distributed consumption growth rates) that conditional expectations of future growth rates of consumption are perfectly correlated across countries.

A setting is also considered in which each country consumes one country-specific consumption good. For this seeting, the structure of

[42] See Hansen (1982) and Cumby, Huizinga & Obstfeld (1983).

[43] I experimented with other sets of instruments (including lagged rates of change of GNP and of real exchange rates). GMM failed to reject the single good model for these alternative sets of instruments.

incomplete asset markets implies a linear restriction on conditional expectations of future rates of change of consumption and real exchange rates.

GMM based tests of the single good model with incomplete asset markets fail to reject that model. Tests based on the Velu et a. (1986) method however suggest that the single good model is rejected.

Velu et al. tests of the model with one country-specific good show that it is strongly rejected for five of the 15 country pairs in the sample. There is hardly any evidence against the model for the remaining country pairs.

CHAPTER III

WORLD BUSINESS CYCLES AND INCOMPLETE INTERNATIONAL ASSET MARKETS

1. Introduction

Standard neoclassical models of international economic fluctuations predict that private consuption is strongly correlated across countries (Scheinkman (1984), Backus, Kehoe & Kydland (1989), Crucini (1989), Baxter & Crucini (1989)). This prediction is incompatible with the low cross country-correlations of consumption which are observed empirically. While existing international theories assume complete international asset markets, this paper develops a model with incomplete international asset markets and it shows that in such a model the cross-country correlation of consumption can be markedly smaller than in models with complete asset markets.

The main reason why previous international models have assumed complete asset markets is presumably tractability: with complete asset markets, competitive equilibria are Pareto optimal and hence they can be determined by solving a social planning problem. Simple numerical methods for solving such a problem are available (c.f. Rebelo (1989 a,b); King, Plosser & Rebelo (1988); Backus, Kehoe & Kydland (1989)). With incomplete asset markets, the equivalence between competitive equilibra and Pareto optima breaks down. I present a method for solving a stochastic dynamic general equilibrium model with incomplete asset markets which is (almost) as simple as the methods used by previous researchers to solve models with complete asset markets.

The paper also argues that a model with *additive* technology shocks which do not affect marginal factor productivities is better able to explain the observed positive correlations of investment and output across countries than standard models in which *multiplicative* shocks to total (and marginal) factor productivities are the source of fluctuations.[1] The equalization of the expected marginal products of capital in different countries which takes place in an integrated world capital market implies that when labor supply elasticities are low, then in the absence of shocks to the marginal product of capital schedule, capital and output are positively correlated across countries. A positive additive shock to the technology of one of the countries induces increases in the capital stocks of all countries in the world (provided the shock leads to a fall in the world interest rate). In contrast, a persistent positive multiplicative shock to the technology of one of the countries increases the marginal product of capital in that country; hence it induces an increase in investment and output in that country and - unless there are strong technological 'spill-overs' between the two countries - it leads to a fall in the other country's investment and output; unless multiplicative technology shocks are strongly correlated across countries, investment tends to be negatively correlated across countries.

In comparison, even additive technology shocks which are not correlated

[1] Specifically, I consider production functions of the type $y=\theta*f(K)-g$, where θ and g are random variables representing *multiplicative* and *additive* technology shocks respectively, while K denotes physical capital.

across countries lead to positive cross-country correlations of investment (and output), provided that labor supply elasticities are low.

One possible interpretation of the additive technology shocks is as government consumption which is financed through lump-sum taxes. My solution method for the model with incomplete asset markets consists in approximating the agents' optimal decision rules by linear functions. Using these linear decision rules and imposing market clearing allows one to solve for an 'approximate' equilibrium.

Section 2 presents the two country model with incomplete asset markets and discusses the method used to numerically solve this model. Section 3 compares the behavior of the model to the behavior which would obtain in the presence of complete asset markets. Stylized facts about international business cycles are described in section 4. I use simulations to explore the properties of the model; simulation results are presented in section 5. Section 6 summarizes the findings obtained in this chapter.

2. The Model

2.1 *Preferences and Technologies*

I consider a world consisting of two countries (i=1,2); each country is inhabited by one consumer (the model can easily be extended to allow for different population sizes in the two countries). There exists a unique good in this world, which is produced and consumed by both countries. This good can also be used as an investment good.

Country i's intertemporal preferences are represented by

$$V_t^i=u^i(c_t^i,L_t^i,\tau_t^i)+\beta(c_t^i,L_t^i,\tau_t^i)*E_t\{V_{t+1}^i\}; \tag{1}$$

here V_t^i is i's expected lifetime utility in period t; c_t^i is i's consumption at date t, L_t^i denotes the number of hours worked by country i and τ_t^i is a random variable which represents an exogenous taste shock. E_t denotes expectations conditional on all informations available in period t; $0<\beta(c_t^i,L_t^i,\tau_t^i)\equiv 1/(1+b(c_t^i,L_t^i,\tau_t^i))<1$, where $b(c_t^i,L_t^i,\tau_t^i)$ is the agent's subjective rate of time preference between periods t and t+1;

I assume that the subjective rate of time preference between periods t and t+1 is a function of c_t^i, L_t^i and τ_t^i.[2] This assumption allows to guarantee the existence of a steady state which is unique (at least locally) as well as the existence of an equilibrium which involves stationary fluctuations around that steady state. This makes it possible to apply standard techniques used in business cycle analysis to solve for this equilibrium and to explore its properties using simulations.

Note that in the special case (which is usually considered in closed economy macro-economics) where β is constant, (1) gives the 'standard' intertemporal utility function $V_t=E_t\sum_{\tau=0}^{\tau=\infty}\beta^{\tau}*u_{t+\tau}$.

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[2] The idea that the rate of time preference depends on consumption has a long history (for an early discussion, see Fisher (1930); Obstfeld (1989) gives a recent discussion). Examples of its use in the international economics literature are Calvo & Findlay (1978), Obstfeld (1981 a,b), Mendoza (1989 a,b). Other examples can be found in Beals & Koopmans (1969), Lucas & Stokey (1984), Epstein (1987).

The period utility function u is increasing in consumption and decreasing in labor and it is strictly concave in both arguments; the function β is assumed to have properties which guarantee that V_t is increasing and concave in current and future consumption.

Country i has a technology which is described by:

$$y_t^i=\theta_t^i f^i(K_t^i,L_t^i)-g_t^i; \tag{2}$$

$f^i()$ is a function with the usual neoclassical properties. K_t^i is country i's capital stock in period t, while θ_t^i and g_t^i are stochastic shocks which are assumed to be stationary (the model abstracts from growth). In standard business cycle theories, *multiplicative* technology shocks (θ_t^i) are the only source of economic fluctuations. Note that while θ_t^i affects the marginal product of capital ($\theta_t^i * f^{i\prime}(K_t^i)$), the *additive* shock g_t^i does not. This is why - as simulations of the model show - the multiplicative shock has a much stronger impact on investment than the additive shock. g_t^i can either be interpreted as a technology shock or as government consumption which is financed by a lump-sum tax. In that case, the output of country 'i' is given by

$$q_t^i\equiv\theta_t^i f^i(K_t^i,L_t^i) \tag{3}$$

and y_t^i (as defined in (2)) is the output of country i which is available to the private sector. Note that the L_t^i variable which appears in the production function and in the utility function (1) is the same: all labor supplied by country i has to be used for production in that same country (i.e. labor cannot move internationally).

The law of motion of the capital stock in country 'i' is:

$$K^i_{t+1}=(1-d)*K^i_t+I^i_t \ ; \tag{4}$$

here d is a fixed rate of depreciation, while I^i_t is gross investment made in period t. Hence country i's capital stock in period t+1 is 'predetermined' in t: country i decides in t what its capital stock in the following period will be.

2.2 Unconditional Borrowing and Lending

In period t, country i can make (or receive) one period loans at the real rate r_t: if 'i' makes a loan of A^i_t units of the good in period t, it gets back $(1+r_t)*A^i_t$ units in period t+1 (the model can be extended to allow for loans of different maturities). $A^i_t>0$ means that the country is a lender and $A^i_t<0$ means that it is a borrower in period t.

Country i's period t budget constraint is:

$$c^i_t+K^i_{t+1}+A^i_t=\theta^i_t f^i(K^i_t,L^i_t)-g^i_t+(1-d)*K^i_t+(1+r_{t-1})*A^i_{t-1} \ . \tag{5 a}$$

In order to rule out Ponzi schemes, the following constraint is imposed:[3]

$$-Z \le A^i_t \quad \text{for all } t \ . \tag{5 b}$$

Z is a large positive number.

The real rate r_t is set in period t and hence it can depend on informations available in period t, but not on informations which become available in

[3] See Sargent (1987), p.364, for an example of the use of this constraint.

t+1. Hence there are no insurance markets in the present model.[4]

The decision problem of a resident of country i is to maximize the lifetime utility defined in (1) subject to the constraints that (5 a) and (5 b) hold in all periods.

The equilibria on which I focus involve small fluctuations of all variables of the model around a steady state. In these equilibria, the constraint (5 b) never binds and therefore the solution to the decision problem of a resident of country i satisfies the following familiar Euler equations:[5]

$$u^i_{1,t}+\beta^i_{1,t}*E_tV^i_{t+1}=(1+r_t)*\beta^i_t*E_t\{u^i_{1,t+1}+\beta^i_{1,t+1}*E_{t+1}V^i_{t+2}\} \quad \text{for } i=1,2\ . \tag{6}$$

$$u^i_{1,t}+\beta^i_{1,t}E_tV^i_{t+1}=\beta^i_tE_t\left\{[\theta^i_{t+1}f^i_{1,t+1}+(1-d)]*[u^i_{1,t+1}+\beta^i_{1,t+1}E_{t+1}V^i_{t+2}]\right\} \quad \text{for } i=1,2. \tag{7}$$

$u^i_{s,t}$ and $f^i_{s,t}$ denote the derivatives of $u^i(c^i_t,L^i_t,\tau^i_t)$ and $f^i(K^i_t,L^i_t)$ with

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[4] Risk-averse agents whose incomes fluctuate in an unpredictable way due to exogenous factors, have an incentive to issue or purchase assets which provide them with insurance against these fluctuations. The absence of such insurance can be justified by assuming that the income of a given country (or the exogenous factors which lead to fluctuations in that income) cannot be observed by agents in the other country (or only at a prohibitive cost).

[5] To motivate (6), note that country i behaves optimally, its expected lifetime utility does not change if it reduces its consumption in period t by an infinitesimal amount ε in order to buy a bond which has a return r_t in period t+1 and if it uses the proceeds from this investment to increase his consumption at date t+1 by $(1+r_t)*\varepsilon$. A similar reasoning shows that when the agent behaves optimally, his expected lifetime utility does not change if he reduces his consumption in period t by an infinitesimal amount ε in order to increase his capital stock in t+1 by ε and if he consumes the proceeds from this additional investment (so that the capital stock in t+2 and in subsequent periods is unchanged). This explains (7).

respect to the s^{th} argument of these functions.

In addition we have that in every period (and for all realizations of the exogenous shocks) country i equates its marginal rate of substitution between labor and consumption to the marginal product of labor and therefore:

$$\{u^i_{2,t}+\beta^i_{2,t}*E_tV^i_{t+1}\}+[\theta^i_t*f^i_{2,t}]*\{u^i_{1,t}+\beta^i_{1,t}*E_tV^i_{t+1}\}=0. \tag{8}$$

Given exogenous processes $\{\theta^1_t,\theta^2_t,g^1_t,g^2_t,\tau^1_t,\tau^2_t\}$,[6] an *equilibrium with incomplete asset markets* is a set of stochastic processes for the endogenous variables of the model $\{c^1_t,c^2_t,L^1_t,L^2_t,V^1_t,V^2_t,K^1_t,K^2_t,A^1_t,A^2_t,r_t\}$ which satisfy (1), (5 a), (6), (7) and (8) as well as the condition that the loan market clears:[7]

$$A^1_t+A^2_t=0 \qquad \text{for all } t\ . \tag{9}$$

A *steady state with incomplete asset markets* is an *equilibrium with incomplete asset markets* in which all exogenous and endogenous variables are constant. Given constant values $(\theta^1,\theta^2,g^1,g^2,\tau^1,\tau^2)$ for the exogenous shocks, a steady state is thus a solution of (1), (5 a), (6), (7), (8) and (9) in which the endogenous variables of the model are constant.

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[6] And assuming that (5 b) never binds.

[7] By Walras' law, market clearing in the loan market implies that the goods market clears as well.

Given constant values $(\theta^1,\theta^2,g^1,g^2,\tau^1,\tau^2)$ for the exogenous shocks, a steady state with incomplete asset markets can thus be described by a vector $(c^1,c^2,L^1,L^2,V^1,V^2,K^1,K^2,A^1,A^2,r)$ which satisfies the following equations:

$$V^i= u^i/(1-\beta^i) \qquad \text{for } i=1,2; \tag{10 a}$$

$$c^i=\theta^i*f^i-d*K^i-g^i+r*A^i \qquad \text{for } i=1,2; \tag{10 b}$$

$$(1+r)*\beta^i=1 \qquad \text{for } i=1,2; \tag{10 c}$$

$$(u^i_2+\beta^i_2*V^i)+[\theta^i*f^i_2]*(u^i_1+\beta^i_1*V^i)=0 \qquad \text{for } i=1,2; \tag{10 d}$$

$$r+d=\theta^i*f^i_1 \qquad \text{for } i=1,2; \tag{10 e}$$

$$A^1+A^2=0 \ . \tag{10 f}$$

Hence a steady state is the solution to a system of 11 equations in 11 unknowns. Note that if β is constant there does not exist a unique steady state because then the two equations in (10 c) reduce to $(1+r)*\beta=1$.

While working on this paper, I learned about recent work by Conze & Scheinkman (1990) and Stockman & Tesar (1990) which gets results which (with respect to explaining low cross-country correlations of consumption) are similar to those which I present. Conze & Scheinkman (1990) too consider a model with incomplete international asset markets. In their paper,[8] the only way in which agents can protect themselves against fluctuations in their labor productivity is by holding a durable asset ('money'). Agents must hold a positive amount of this asset, i.e. there is no borrowing at all in Conze & Scheinkman (1990).
An important difference between their work and my paper is that my solution method makes it possible to consider a model which is less stylized than their's; for example, my model allows for physical capital,

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[8] As in Scheinkman & Weiss (1986).

several types of exogenous shocks etc.

Stockman & Tesar (1990) present a model with complete international asset markets in which non-traded consumption goods and country-specific taste shocks break the close cross-country correlation of consumption.[9]

As a first pass, and in order to focus on the effects of limitations on asset markets, the present paper simply abstracts from non-traded consumption goods.

2.3. *The Solution Method*

An approximate solution to the model can be obtained by deriving a linear approximation to the equations listed in the definition of the equilibrium and by solving the resulting system of expectational difference equations. For the case where the exogenous shocks have small variances, and where there would exist a locally unique steady state for constant values of the exogenous variables equal to the unconditional means of the random shocks θ_t^1, θ_t^2, g_t^1, g_t^2, τ_t^1 and τ_t^2 Woodford (1986)[10] provides a rigorous justification for conducting the linear approximation around this steady state.

Hence I assume that when the exogenous shocks are equal to their unconditional means, a (locally) unique steady state exists ; I expand the equations listed in the above definition of an equilibrium in Taylor series around this steady state and I keep the linear terms of these expansions.

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[9] Their empirical analysis shows that allowing for non-traded goods in a model with complete asset markets is not sufficient to reduce the cross-country correlation of consumption to the correlations which are observed empirically; this is why Stockman & Tesar resort to country specific taste shocks.

[10] See also the discussion in Rotemberg & Woodford (1989).

In what follows $\Delta\hat{x}_t$ denotes the percentage deviation of a variable x_t from its steady state.[11]

Given exogenous forcing processes for $\{\Delta\hat{\theta}^1_t, \Delta\hat{\theta}^2_t, \Delta\hat{g}^1_t, \Delta\hat{g}^2_t, \Delta\hat{\tau}^1_t, \Delta\hat{\tau}^2_t\}$, *an approximate equilibrium with incomplete asset markets* is a set of stochastic processes for the endogenous variables $\{\Delta\hat{c}^1_t, \Delta\hat{c}^2_t, \Delta\hat{L}^1_t, \Delta\hat{L}^2_t, \Delta\hat{V}^1_t, \Delta\hat{V}^2_t, \Delta\hat{K}^1_t, \Delta\hat{K}^2_t, \Delta\hat{A}^1_t, \Delta\hat{A}^2_t, \Delta\hat{r}_t\}$ which satisfy the linearized versions of the equations listed in the definition of an equilibrium.

In what follows, I focus on 'approximate equilibria'.

Keeping the linear terms of the Taylor expansion of the equations listed in the definition of an equilibrium, we obtain (after a few substitutions) a system of equations which can be written as:

$$E_t h_{t+1} = G_0 * h_t + G_1 * z_t + G_2 * E_t z_{t+1} \ , \tag{11}$$

where $h_t \equiv (\Delta\hat{r}_{t-1}, \Delta\hat{K}^1_t, \Delta\hat{K}^2_t, \Delta\hat{A}^1_t, \Delta\hat{V}^1_t, \Delta\hat{L}^1_t, \Delta\hat{c}^1_t, \Delta\hat{V}^2_t, \Delta\hat{L}^2_t, \Delta\hat{c}_t)'$ and $z_t \equiv (\Delta\hat{\theta}^1_t, \Delta\hat{\theta}^2_t, \Delta\hat{g}^1_t, \Delta\hat{g}^2_t, \Delta\hat{\tau}^1_t, \Delta\hat{\tau}^2_t)'$. G_0 is a 10x10 matrix, while G_1 and G_2 are 10x6 matrices.

The first four elements of the vector h_t are 'predetermined' at date t (i.e. they are known at t-1), while the remaining elements are 'non-predetermined'. As shown by Blanchard & Kahn (1980), a necessary and

[11] With the following exception(s): in order to allow for cases where steady state asset holdings are zero and where $g^i=0$, $\Delta\hat{A}^i_t$ and $\Delta\hat{g}^i_t$ are defined by dividing the differences between A^i_t and g^i_t and their respective steady states by y^i.

sufficient condition for a model like (11) to have a unique stationary solution is that the number of eigenvalues of the matrix G_0 outside the unit circle equals the number of non-predetermined variables.

For a simple version of the model where labor supplies are fixed, $\partial\beta/\partial c<0$ can be shown to be a necessary condition for guaranteeing the existence of a unique stationary solution.[12]

The approximation method used here is equivalent to a method which consists in substituting agent i's budget constraint (5a) into his lifetime utility function (1) and expanding the resulting function in a second order Taylor expansion in all its arguments (for all dates and states) around their respective steady state values; maximizing the quadratic function which this expansion yields, gives a set of first-order conditions which is equivalent to (6)-(8).

Christiano (1990) shows that for a one sector neoclassical growth model this linear-quadratic approximation method performs strikingly well.[13]

Linearizing (6) and (7), we get that

$$\Delta r_t=f^i_1{*}E_t\Delta\theta^i_{t+1}+\theta^i{*}f^i_{1,1}{*}\Delta K^i_{t+1}+\theta^i{*}f^i_{1,2}{*}E_t\Delta L^i_{t+1} \quad . \tag{12}$$

This equation says that each country equates the expected marginal product of its capital stock to the risk-free interest rate. When labor supplies are variable or when the model is subjected to *multiplicative* stochastic

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[12] In the simulations described below, I assume that $\beta(c,L,\tau)=\beta(u(c,L,\tau))$. Numerical 'experiments' with many different parameter values suggest that unless $\partial\beta/\partial u<0$ holds, the Blanchard & Kahn condition for the existence of a unique stationary solution is not likely to be satisfied.

[13] See Taylor & Uhlig (1990) for an overview over different methods for solving the neoclassical growth model.

technology shocks, the marginal product of capital is random; hence (12) shows that the risk premium associated with this randomness is neglected when the linearization method is used. Note however that by making the variance of the exogenous shocks sufficiently small, the risk premium can be made arbitrarily small.[14]

The solution of (11) can easily be determined using Blanchard & Kahn (1980). It has the property that - as ΔA^1_{t-1} is a predetermined variable at date t - the decisions which the two countries make in equilibrium depend on ΔA^1_{t-1} (i.e. on the distribution of financial assets between the two countries). That country i's date t consumption and labor supply choices depend on ΔA^i_{t-1} is not surprising, as these choices clearly depend on i's wealth. As the marginal product of capital depends on the labor input, this implies that physical investment decisions in the two countries are also affected by the distribution of financial wealth between the two countries.

An important special case in which (in the approximate equilibrium) the distribution of financial wealth does not affect the behavior of world-wide aggregates of consumption, labor supplies and the world capital stock

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[14] To see this , let $m_{t,t+1}$ denote the marginal rate of substitution between consumption at dates t and t+1 ; from (6) and (7) we get that $E_t\theta^i_{t+1}*f'(K^i_{t+1})=(r_t+d)-cov(\theta^i_{t+1}*f'(K^i_{t+1}),m^i_{t,t+1})/E_tm^i_{t,t+1}$.

Using the Cauchy-Schwartz inequality, we get

$|E_t\theta^i_{t+1}*f'(K^i_{t+1})-(r_t+d)|\leq\{Var(\theta^i_{t+1}*f'(K^i_{t+1}))*Var(m^i_{t,t+1})\}^{0.5}/E_tm^i_{t,t+1}$.

By making the variances of the exogenous shocks sufficiently small, we can make the variances on the right-hand side of this inequality arbitrarily small.

occurs in the symmetric case when the two countries have identical preferences and technologies [15]: when this condition is satisfied, the coefficients of the linearized versions of the equilibrium conditions (1), (5a), (6), (7) and (8) are the same for i=1,2. Summing each of these linearized equations over i=1,2 gives a system of equations in the world capital stock, world consumption and world labor supply (the net asset positions of the two countries cancel out when the sum over i=1,2 is taken). This system uniquely determines the world-wide agregates of these variables.

3. Comparison of International Business Cycles With Complete and With Incomplete International Asset Markets

In the presence of complete asset markets, agents can at each date t trade in claims to units of consumption at date t+s, conditional on each possible 'history' of the realizations of the exogenous shocks between dates t and t+s. Let $x_t \equiv (\theta_t^1, \theta_t^2, g_t^1, g_t^2, \tau_t^1, \tau_t^2)$ and let $X^t \equiv \{\ldots, x_{t-2}, x_{t-1}, x_t\}$ denote the history of the shocks which have affected the two countries up until date t. In what follows, $q_t(X^s)$ with $s \geq t$ denotes the price at date t (in terms of output at t) of a claim to one unit of output to be delivered at date s if the history X^s occurs and $w_t^i(X^s)$ is the number of such claims which country i holds at the end of period t. Given X^t, the date t budget of country 'i' is thus:

[15] What is needed is identical u, β and f functions for the two countries and identical steady state values of the exogenous shocks affecting the two countries (and hence identical steady state values of the decision variables of the two countries).

$$c_t^i+K_{t+1}^i+\sum_{s=t+1}^{\infty}\int q_t(X^s)w_t^i(X^s)dX^s=\theta_t^i f^i(K_t^i,L_t^i)-g_t^i+w_{t-1}^i(X^t)+\sum_{s=t+1}^{\infty}\int q_t(X^s)w_{t-1}^i(X^s)dX^s. \qquad (13)$$

The decision problem of country 'i' is to maximize its expected life-time utility subject to the restriction that this budget constraint holds in all dates and for all possible histories.

In the presence of complete date and state contingent asset markets, we can characterize equilibria by using the equivalence between competitive equilibria and Pareto optima. Pareto optima can be found by maximizing a weighted sum of the expected lifetime utility levels of the residents of countries 1 and 2; therefore we can characterize equilibria by imagining that in some 'initial' time period t=0, the following problem is solved:

$$\text{Maximize} \qquad W_0\equiv\lambda*V_0^1+(1-\lambda)*V_0^2 \quad [16] \qquad (14)$$

subject to the constraint that the world resource constraint holds in all periods $t\geq0$:

$$c_t^1+c_t^2+K_{t+1}^1+K_{t+1}^2=\theta_t^1 f(K_t^1,L_t^1)+\theta_t^2 f(K_t^2,L_t^2)-g_t^1-g_t^2+(1-d)*K_t^1+(1-d)*K_t^2 \quad . \qquad (15)$$

Differentiating the objective function of the social planner with respect to the consumptions of the two countries gives the following condition:

$$B_{t-1}*\lambda*\{u_{1,t}^1+\beta_{1,t}^1*E_tV_{t+1}^1\}=(1-\lambda)*\{u_{1,t}^2+\beta_{1,t}^2E_tV_{t+1}^2\}, \qquad (16)$$

[16] λ and $(1-\lambda)$ are 'welfare' weight attached to the residents of the two countries. These weights reflect the wealth of the two countries (see Backus, Kehoe & Kydland (1989) and Rebelo (1988 a,b)).

where $B_{t-1} \equiv B_{t-2} * \beta^1(c^1_{t-1}, L^1_{t-1}, \tau^1_{t-1}) / \beta^2(c^2_{t-1}, L^2_{t-1}, \tau^2_{t-1})$. (17)

Hence we see that with complete asset markets, weighted marginal utilities of consumption of the two countries are equated in period t, where the weights attached to the marginal utilities reflect the 'welfare weights' $(\lambda, 1-\lambda)$ in the above social planning problem, as well as the consumptions of the countries *before* date t.

Note that in the standard case where β is constant (16) becomes: $\mu * u^1_{1,t} = u^2_{1,t}$, where μ is a constant. If the utility function u is separable in consumption and its other arguments, then $\mu * u^1_{1,t} = u^2_{1,t}$ implies that (locally) consumption is perfectly correlated across countries.

(16) and (17) imply that with complete markets, the marginal rates of substitution of the two countries between consumption at dates t and t+1 are equated, and that for all possible realizations of the exogenous shocks in the two periods. In contrast to this, with incomplete markets, marginal rates of substitution between t and t+1 are in general not equated across the two countries, although they are in expected value: (6) implies

that $1/(1+r_t) = \beta^1_t * [E_t\{u^1_{1,t+1} + \beta^1_{1,t+1} * E_{t+1} V^1_{t+2}\}] / [u^1_{1,t} + \beta^1_{1,t} * E_t V^1_{t+1}] =$

$$\beta^2_t * [E_t\{u^2_{1,t+1} + \beta^2_{1,t+1} * E_{t+1} V^2_{t+2}\}] / [u^2_{1,t} + \beta^2_{1,t} * E_t V^2_{t+1}] .$$

Equation (7) and (8) are clearly valid first-order condition for the Pareto problem.

Given a welfare weight λ and exogenous shock processes $\{\theta^1_t, \theta^2_t, g^1_t, g^2_t, \tau^1_t, \tau^2_t\}$ an *equilibrium with complete asset markets* is a set of stochastic processes for the endogenous variables $\{c^1_t, c^2_t, L^1_t, L^2_t, V^1_t, V^2_t, K^1_t, K^2_t\}$ which satisfy (1), (7), (8), (15), (16) and (17).

One can easily show that the model with complete asset markets has the same steady state consumptions, labor supplies and capital stocks (and hence the same steady state interest rate) as the model with incomplete asset markets. Linearizing the equations listed in the definition of an equilibrium with complete asset markets around a steady state gives (after several substitutions) a system of equations of the form:

$$E_t e_{t+1}=H_0*e_t+H_1*z_t+H_2*E_t z_{t+1} \,, \tag{18}$$

where $e_t \equiv (\Delta\hat{B}_{t-1}, \Delta\hat{K}^1_t, \Delta\hat{K}^2_t, \Delta\hat{V}^1_t, \Delta\hat{L}^1_t, \Delta\hat{c}^1_t, \Delta\hat{V}^2_t, \Delta\hat{L}^2_t, \Delta\hat{c}^2_t)'$ and $z_t \equiv (\Delta\hat{\theta}^1_t, \Delta\hat{\theta}^2_t, \Delta\hat{g}^1_t, \Delta\hat{g}^2_t, \Delta\hat{\tau}^1_t, \Delta\hat{\tau}^2_t)'$. H_0 is a 9x9 matrix, while G_1 and G_2 are 9x6 matrices.

Linearizing (7), (16) and (17) shows that in an approximate equilibrium with complete asset markets, the expected marginal products of capital are equated across countries (as is the case with incomplete markets, see (12)):

$$\begin{aligned} f^1_1*E_t\Delta\theta^1_{t+1}+\theta^1*f^1_{11}*\Delta K^1_{t+1}+\theta^1*f^1_{12}*E_t\Delta L^1_{t+1} &= \\ f^2_1*E_t\Delta\theta^2_{t+1}+\theta^2*f^2_{11}*\Delta K^2_{t+1}&+\theta^2*f^2_{12}*E_t\Delta L^2_{t+1}. \end{aligned} \tag{19}$$

As the structure of asset markets affects the behavior of consumption and labor, the (in)-completeness of asset markets affects (in general) the behavior of capital in the two countries (because the marginal product of capital depends on labor).

It can however be shown that with identical preferences and technologies the (in)completeness of asset markets does not affect the behavior of worldwide aggregates of consumption, labor supply and capital in the approximate equilibrium. With fixed labor supplies, the condition that expected marginal products of capital are equated across countries implies that the capital stock in a given country can be expressed as a function of the world capital stock (and of expected future multiplicative technology shocks). Hence we get the result that with identical preferences and technologies and with fixed labor supplies, the approximate equilibrium has the property that the behavior of capital (and hence of output) in each of the two countries is the same with complete and with incomplete asset markets.

4. Stylized Facts

I evaluate the model with incomplete asset markets by comparing its predictions to the key stylized facts which characterize business cycles in industrialized countries (the following 'stylized facts' describe properties of detrended data).[17]

As documented in tables 15-17 for quarterly data, output, private consumption and investment are positively correlated across countries; consumption seems to be less correlated across countries than output. The averages of the cross-country correlations of quarterly growth rates of total private consumption, GNP and fixed investment in a sample countries consisting of the US, Japan, France, Britain, Italy and Canada during the period 1971:2-88:1 are 0.22, 0.26 and 0.24 respectively. When log

[17] See e.g. Backus & Kehoe (1989 a,b); Backus, Kehoe & Kydland (1989); Stockman & Tesar (1990) for statistical descriptions of salient features of international business cycles.

consumptions, log GNP and log investment are detrended using a linear time trend, the corresponding average cross-country correlations are 0.29, 0.40 and 0.42 respectively.

Within a given country the correlation of consumption and hours worked is low. Table 18 reports correlations between detrended consumption and hours for the US, Japan, France, Britain, Italy and Canada. The average of the within-country correlations of consumption and hours is 0.21.

Additional stylized facts which are of interest in evaluating the model developed in this paper are:

Within a given country, investment and consumption are procyclical; net exports are countercyclical; consumption tends to be less volatile than output, while investment is more volatile than output. For the US economy Backus, Kehoe & Kydland (1989) report that the percentage standard deviations of Hodrick-Prescott (1980) detrended quarterly output, consumption (non-durables and services) and fixed investment are 1.74, 0.86 and 5.51 respectively (in %).

The discussion in the next section centers mostly on the behavior of consumption, as the effect of the incompleteness of asset markets is strongest for that variable, and on the correlations of investment of output across countries, because the international correlation of investment and output differs significantly according to whether multiplicative or additive shocks are the source of fluctuations.

5. Simulations of the Model

5.1. *The Parameters of the Model*

I use simulations in order to investigate the properties of the model. These simulations focus on the symmetric case where the two countries have the same utility and productions functions, where the steady state values of the shocks which affect these countries are the same and where the steady state values of the net asset positions of the two countries are zero.

5.1.1. **Preference and Technologies**

The production function f(K,L) is assumed to be Cobb-Douglas:

$$f(K^i,L^i)=(K^i)^{\eta}*(L^i)^{1-\eta} \text{ with } 0\leq\eta\leq 1 \quad . \tag{20}$$

Concerning the utility function u, a standard Real Business Cycle specification is adopted:[18]

$$u(c,L)=(1/(1-\sigma))*(c*e^{-\psi(L)})^{1-\sigma}, \qquad \text{where } \sigma>0,\ \sigma\neq 1 \tag{21}$$

and ψ is an increasing function. One easily verifies that the concavity of u in c and L requires that $\nu_2+(\sigma-1)*\nu_1>0$ and $\sigma*\nu_2+(\sigma-1)*\nu_1>0$, where $\nu_1\equiv\psi'(L)*L$ and $\nu_2\equiv\psi''*L/\psi'$. Note that the steady state condition (10 d) implies (for the production function specified in (21)) that $\nu_1=(1-\eta)/(c/y)$; where $c/y = 1 - (g/y) - d*K/y = 1 - \gamma - d*(\alpha/r+d)$.

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[18] See King, Plosser & Rebelo (1990); Rotemberg & Woodford (1989).

ν_2 can be interpreted as a labor supply parameter:[19] holding constant country i's marginal instantaneous utility of consumption in period t $(u_c(c_t,L_t,\tau_t))$ and holding fixed E_tV_{t+1}, the elasticity of country i's labor supply with respect to its real wage rate (its marginal product of labor) is $els \equiv \sigma/(\sigma*\nu_2+(\sigma-1)*\nu_1)$.

Holding constant the life-time utilities and the labor supplies of a country in periods t and t+1, $(1/\sigma)$ is its intertemporal elasticity of substitution between consumption in t and t+1. σ also affects the way in which consumption and hours worked interact in the instantaneous utility function u: the elasticity of u_1 with respect to hours worked is $-(1-\sigma)*\nu_1$. By assumption ν_1 is positive. Hence consumption and hours are substitutes in the instantaneous utility function $u(c,L,\tau)$ if $\sigma<1$ and they are complements when $\sigma>1$. In the simulations, I focus on cases where consumption and hours are complements.

The function $\beta(c,L,\tau)$ is assumed to be of the form

$$\beta(c,L,\tau)=\beta(u(c,L,\tau)) \text{ with } \partial^2\beta/\partial u^2=0. \tag{22}$$

The elements of the matrices G_0, G_1, G_2 and H_0, H_1 and H_2 which determine the equilibrium behavior of the endogenous variables (see (11), (18)) can be expressed as functions of: r (the steady state interest rate), d (the depreciation rate of physical capital), η (the elasticity of output with respect to capital), the ratio $\gamma \equiv g/y$ (this share too is assumed to be the

[19] See the discussion in Rotemberg & Woodford (1989).

same for both countries), σ, $\varepsilon_{\beta,u}\equiv(\partial\beta/\partial u)*(u/\beta)$ and $\nu_2\equiv\psi''*L/\psi'$.

To simulate the model, one thus has to choose values for r, d, η, σ, $\varepsilon_{\beta,u}$ and ν_2; in addition a stochastic process of the exogenous shocks has to be specified.

η (the elasticity of output with respect to capital) is set equal to 0.35. In all simulations, $(\partial\beta/\partial u)*(u/\beta)=-0.10$ is assumed. The baseline choice for σ is: $\sigma=2$ (Backus et al. (1989) use $\sigma=1.35$ as their baseline value). The remaining preference and technology parameters are chosen with the aim of comparing the predictions of the model to quarterly data; this motivates the following baseline choices for the steady state interest rate and the depreciation rate of capital: r=0.01 and d=0.025.

5.1.2. The Stochastic Processes Followed by the Exogenous Shocks

In the simulations, I abstract from taste shocks and I study the effects of one type of shock at a time, i.e. the model is either subjected to multiplicative or to additive shocks. These shocks are assumed to follow bivariate AR(1) processes:

$$\begin{aligned} z_t^{\theta}&=RHO_{\theta}*z_{t-1}^{\theta}+\varepsilon_t^{\theta}, \quad \text{and} \\ z_t^{g}&=RHO_{g}*z_{t-1}^{g}+\varepsilon_t^{g} \end{aligned} \tag{23}$$

where $z_t^{\theta}=(\Delta\theta_t^1,\ \Delta\theta_t^2)'$ and $z_t^{g}=(\Delta g_t^1,\ \Delta g_t^2)'$ while RHO_{θ} and RHO_g are 2x2 matrices, while ε_t^{θ} and ε_t^{g} are vectors of white noises with covariance matrices $VAR_{\theta}\equiv E\varepsilon_t^{\theta}\varepsilon_t^{\theta\prime}$ and $VAR_g\equiv E\varepsilon_t^{g}\varepsilon_t^{g\prime}$ respectively.

The effect of the incompleteness of asset markets on the cross-country correlation of consumption is likely to be affected by how strongly these shocks are correlated between countries: in the extreme case where the shocks affecting different countries are perfectly correlated, there is no scope for risk-sharing between these countries and hence it does not matter whether asset markets are complete or incomplete.
Hence it is important to estimate empirically how strong the correlation of the 'shocks' affecting different countries is.

Backus et al. (1989) present a model with multiplicative productivity shocks. They assume that the shocks affecting the two countries in their model follow an AR(1) process. Backus et al. fit an AR(1) process to detrended quarterly log Solow residuals for the US and an aggregate of three large European countries (Germany, France and Britain). Their estimates of RHO_{θ} and V_{θ} (see (23)) are:

$$RHO_{\theta}=\begin{bmatrix} .93 & .04 \\ .09 & .92 \end{bmatrix} \text{ and } V_{\theta}=\begin{bmatrix} .0068^{2} & .00001 \\ .00001 & .0078^{2} \end{bmatrix}.$$

The correlation between the technology shocks affecting the two countries which is implied by these RHO_{θ} and V_{θ} matrices is 0.85. A correlation of 0.85 is much larger than the cross-country correlations which are documented below.

Unfortunately, Backus et al. do not say what data they used for the estimation of the Solow residuals; their paper also fails to provide standard errors for the estimates of RHO_{θ} and V_{θ}. It would be useful to know these standard errors, because the cross-country correlations implied by the estimates of RHO_{θ} and V_{θ} provided by Backus et al. are extremely sensitive to the values of the elements of the RHO_{θ} matrix. It appears for example that when the off-diagonal elements of the RHO_{θ} matrix reported by Backus et al. are set equal to zero, the cross-country correlation of the technology shocks drops from 0.85 to 0.18, which is much more consistent with the cross-country correlations reported in the appendix.

Costello (1989) presents estimates of cross-country correlations of Solow residuals which differ strongly from those implied by the Backus et al. estimates of RHO_θ and V_θ. She studies productivity growth in five two-digit SIC industries in six of the G7 countries. Her paper suggests that the cross-country correlations of detrended Solow residuals in these industries are close to zero.

Table 19 presents estimates of the correlation of detrended aggregate quarterly Solow residuals in the the US, Japan, France, Italy and Canada. These residuals are calculated using the formula $\ln(\theta) = \ln(y) - \eta^*\ln(K) - (1-\eta)^*\ln(L)$, assuming that $\eta=0.35$ (y, K and L are GNP, the stock of physical capital and hours worked respectively).

The GNP data are taken from the OECD Quarterly National Accounts.

Two measures of labor inputs are used: (i) for the US, Japan and France data on total hours worked in the non-agricultural sector are used. For the US, data on total employee hours are taken from Citibase (series LPMHU). For Japan and France, an index of hours worked was constructed by multiplying the series 'employment in non-agricultural establishments' and 'weekly hours of work in the non-agricultural sector' from the Bulletin of Labour Statistics published by the International Labour Office (ILO).

(ii) The second labor measure is the employment data provided by the International Financial Statistics (IFS). This measure is used for the US, Japan, France, Italy and Canada. The IFS employment series measure total numbers of employees and workers (and not total *hours*).

Capital stocks were estimated using data on gross capital formation from the International Financial Statistics, assuming a depreciation rate of 10%

p.a. .[20]

Table 19 presents cross-country correlations for first differences of log Solow residuals and for linearly detrended log Solow residuals.

A two-sided test of the null-hypothesis that cross-country correlations of Solow residuals are zero rejects that hypothesis at the 10% level for half of the estimated correlation coefficients reported in table 19.

The arithmetic average of the correlation coefficient based on first differenced Solow residuals is 0.19; for first differences Solow residuals, the average cross-country correlation is 0.21.

Changing the technology parameter η does not have a strong effect on the estimated cross-country correlations: for $\eta=0.25$, the average cross-country correlations of first differences and of linearly detrended log Solow residuals for the countries in the sample are 0.17 and 0.13 respectively. For $\eta=0.45$, the corresponding average cross-country correlations are 0.21 and 0.29 respectively.

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[20]From the identity $K_{t+1}=(1-d)*K_t+I_t$, where 'd' is the depreciation rate, I is gross investment capital, capital stocks can be constructed for periods $t=1,\dots,T$ if one has an estimate of the capital stock in the 'initial' period t=0. In a steady state where GNP grows at a constant rate 'g' and where the capital to GNP and the investment to GNP ratios are constant, we have that $k=i/(g+d)$, where k and i are the steady state capital to GNP and the steady state investment to GNP ratios respectively. I estimate the 'initial' capital stock by $\hat{K}_0\equiv(\hat{i}/(\hat{g}+d))*\overline{Y}_0$, where $\hat{g}$ is the average GNP growth rate during the sample period and $\overline{Y}_0$ is the average GNP during the first four quarters of the sample, while $\hat{i}\equiv(1/T)*\sum I_t/Y_t$ is the sample average of the investment to GNP ratio.

Table 20 estimates trivariate AR(1) processes using detrended log Solow residuals for the US, Japan and France (the Solow residuals are constructed using η=0.35).

The autocorrelation (RHO) matrices which are estimated for first differences of log Solow residuals, are never significantly different from zero.

Almost all off-diagonal elements of the estimated autocorrelation matrices which are estimated using linearly detrended log Solow residuals fail to be significantly different from zero at conventional significance levels.

In the simulations presented below, I therefore assume that the off-diagonal elements of the 'RHO' matrix are zero.

5.2. *Simulation Results: Fixed Labor Supplies*

In the first set of simulations, reported in tables 22-27 I assume that labor supplies are fixed. I start the simulations of the model with this case because of the strong empirical evidence according to which labor supply elasticities are small (particularly for males).[21]

Intuitively we expect that the effects of limitations on international asset markets are stronger, the smaller the cross-country correlations of the exogenous shocks. The simulation in tables 22-27 show that when labor supplies are fixed and when shocks are not correlated across countries, the incompleteness of asset markets has a strong effect on the correlations of consumption and of life-time utilities across countries: in all the cases shown in tables 22-27, the cross-country correlation of consumption in the

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[21] See Pencavel (1986).

presence of incomplete asset markets is smaller than 0.25. Furthermore the correlation between the expected life-time utility levels of the two countries tends to be *negative*. In the presence of complete asset markets, consumption and expected lifetime utility levels would be perfectly correlated across countries (given the assumptions of identical preferences in the two countries and fixed labor supplies).

As noted above, when both countries have identical preferences and technologies and when labor supplies are fixed, the approximate equilibrium has the property that the behavior of physical capital is the same with complete and with incomplete capital markets. The following discussion of the differences between the effects of multiplicative and additive shocks on investment and output applies thus to both asset market structures. The simulations reveal striking differences between the effects of multiplicative and additive shocks. When the model is subjected to multiplicative technology shocks (tables 22-24), investment in a given country tends to be very volatile and in addition investment tends to be negatively correlated across countries, unless the serial correlation of the technology shocks is low; this can be explained by the fact that when a persistent (positively serially correlated) country-specific multiplicative technology shock occurs in one of the countries, the increase in the marginal product of capital which results from the shock induces a strong increase in that country's investment while the other country's investment falls. This increase in physical investment in the country affected by a serially correlated positive multiplicative shock increases the country's demand for imports and thus implies that the trade balance is countercyclical.

When the model is subjected to multiplicative shocks, output is negatively correlated across countries, unless the serial correlation of the technology shocks is low. Simulations (not reported in the appendix) of the fixed labor supply model show that in order to get positive cross-country correlations of investment in the presence of serially correlated technology shocks, one has to assume that the multiplicative technology shocks are strongly correlated across countries. The same thing is true for output but the cross-country correlations of the multiplicative shocks required to make output positively correlated across countries seem to be lower than those needed to get positive cross-country correlations of investment.

When the fixed labor supply model is subjected to additive shocks (tables 25-27), then investment (and hence capital) is perfectly correlated across countries: as the marginal product of capital in the two countries is unaffected by additive technology shocks, the equalization of the marginal products of capital across the two countries (see (19)) implies that - with fixed labor supplies - the capital stocks of the two countries are perfectly correlated. An additive shock affecting one of the countries induces changes in the capital stocks of all countries in the world which have the same sign (provided the shock affects the world interest rate). Even with additive shocks which are not correlated across countries, the cross-country correlation of output $(y_t^i=\theta_t^i * f(K_t^i,L_t^i)-g_t^i)$ is positive (although it is relatively small). Note however that if one interprets the *additive* shocks g_t^1 and g_t^2 as government consumption financed by lump-sum taxes (according to this interpretation $g_t^i>0$), then the output of country 'i' is given by $q_t^i \equiv \theta_t^i * f^i(K_t^i)$.

This implies that if government consumption (financed by lump-sum taxes) is the only source of fluctuations and if labor supplies are fixed, then both investment *and* output are perfectly correlated across countries. A drawback of a model based on additive shocks is that these shocks induce procyclical behavior of the trade balance.

5.3. *Simulation Results: Variable Labor Supplies*

Simulation results for the model with variable labor supplies and multiplicative shocks are presented in table 28.

The table presents simulation results for a grid of values of the intertemporal elasticity of substitution parameter σ and the labor supply elasticity els.[22]

As labor is immobile internationally, non-separabilities between consumption and work effort reduce the international covariation of consumption.[23] When the labor supply elasticity is sufficiently low and/or σ is sufficiently close to unity,[24] the simulation results are similar to those which obtain with fixed labor supplies; in particular, cross-country correlations of consumption are much lower with incomplete asset markets, than with complete markets.

An increase in σ strengthens the non-separability between consumption and

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[22] The following values of these parameters are considered: σ=1.1, 2, 3, 5, 7 and els=0.1, 0.5, 1.0, 2.0.

[23] See Devereux, Gregory & Smith (1991) for detailed discussions of this point.

[24] For σ=1, the period utility function specified in (21) becomes $u=\ln(c)-\psi(L)$, i.e. it becomes separable in consumption and hours worked.

hours (provided that $\sigma>1$) and holding constant the labor supply elasticity it therefore reduces the cross-country correlation of consumption. Likewise, increases in the labor-supply elasticty reduce the cross-country correlation of consumption (holding constant σ).

By selecting values of σ and els which are sufficiently large, the cross-country correlation of consumption can be made rather small, and that even when asset markets are complete; it appears however that this can only be achieved by letting consumption and hours be strongly correlated within the same country, which is inconsistent with the data.

We see that for $\sigma=1.1$ and els=0.1, the cross-country correlations of consumption with complete and with incomplete asset markets are 0.99 and 0.30 respectively, and that the correlation of consumption and hours within the same country is 0.18 with complete and 0.11 with incomplete asset markets. Increasing σ and els to $\sigma=2$ and els=2 reduces the cross-country correlation of consumption to 0.07 with complete and to 0.02 with incomplete asset markets. We see however that the within-country correlation of consumption and hours now becomes 0.77 with complete and 0.67 with incomplete asset markets, which is too large compared to the average within-country correlation observed in the data.

The simulations results therefore confirm the finding of Gregory et al. (1991) that non-separabilities between consumption and labor allow to strongly reduce the cross-country correlation of consumption, but they suggest that non-separabilities do not allow to reduce the cross-country correlation of consumption to the level seen in the data, while ensuring that within-country correlations of consumption and hours are small. A model with incomplete asset markets however is able to match both the low cross-country correlations of consumption and the low within-country correlations of consumption and hours, and that for 'realistic' cross-country correlations of technology shocks.

6. Summary

This chapter extends the Real Business Cycle (RBC) theory of international economic fluctuations by presenting an RBC model with incomplete international asset markets. In this model only debt contracts can be used for international capital flows.

The paper shows that with incomplete asset markets, the cross-country correlation of consumption can be markedly smaller than in the presence of complete asset markets. The simulations suggest that allowing for non-separabilities between consumption and labor in a one good international Real Business Cycle model with complete asset markets does not allow to reduce the cross-country correlation of consumption to the levels seen in the data while at the same time ensuring that the within-country correlations of consumption and hours are small. A model with incomplete asset markets however is able to match both the low cross-country correlations of consumption and the low within-country correlations of consumption and hours which are observed in the data, and that for 'realistic' cross-country correlations of technology shocks.

The essay shows also that a model which allows for *additive* technology shocks (which can be interpreted as shocks to government consumption) is better able to explain the observed positive correlations of investment and output across countries than standard business cycle theories in which *multiplicative* shocks to total (and marginal) factor productivities are the source of fluctuations.

APPENDIX

nd: non-durables; s: services; ctot: total private consumption;
g: government consumption.
nd, s, ctot, g, GNP series are in constant prices and in per capita terms.
JA: Japan, FR: France, IT: Italy, CA: Canada.

TABLE 1.-- AVERAGE ANNUAL GROWTH RATES (IN %) OF PRIVATE AND GOVERNMENT CONSUMPTION AND OF GNP

	nd	s	ctot	g	GNP
US	0.96	2.36	2.02	0.85	1.72
JA	1.68	3.73	2.98	2.77	3.45
FR	1.66	3.08	2.26	2.26	2.05
UK	0.80	2.83	2.38	1.50	2.11
IT	1.94	3.26	3.00	2.47	2.66
CA	0.92	3.24	2.73	1.58	2.87

NOTE: The sample period is 1971:II-1988:I.
GNP data for US and Japan from International Financial Statistics; for the remaining countries from OECD Quarterly National Accounts.

TABLE 2.-- AVERAGE ANNUAL RATES OF CHANGE OF BILATERAL REAL EXCHANGE RATES

	JA	FR	UK	IT	CA
US	-.05 (-13.29)	-.024 (-6.29)	-.022 (-6.71)	-.020 (-5.80)	-.009 (-8.21)
JA		.031 (11.57)	.034 (8.94)	.036 (12.72)	.047 (10.76)
FR			.030 (1.05)	.005 (3.11)	.015 (3.47)
UK				.002 (0.84)	.012 (3.28)
IT					.010 (2.71)

NOTE: The sample period is 1971:II-1988:I
Average rates of change (p.a.) of bilateral real exchange rates in terms of non-durables consumption. In parentheses: t-statistic for the test of the hypothesis that the unconditional expected value of first differences of quarterly log real exchange rates is zero. The t-statistics are corrected for serial correlation (using the Newey & West (1987) method with 10 lags).

TABLE 3.-- SHARES OF TOTAL PRIVATE CONSUMPTION EXPENDITURES (IN %) ACCOUNTED FOR BY NON-DURABLES AND BY SERVICES IN 1971 AND 1987

	non-durables		services	
	1971	1987	1971	1987
US	41.6	35.4	46.8	49.0
JA	36.3	29.5	45.5	50.7
FR	40.2	36.9	34.6	38.8
UK	39.2	30.7	38.6	40.8
IT	45.0	37.9	30.0	31.1
CA	36.6	26.6	41.6	45.1

TABLE 4.-- AUTOCORRELATIONS OF QUARTERLY CONSUMPTION GROWTH RATES

	n=1	n=2	n=3	n=4	n=5	n=6	n=7	n=8	n=9	n=10
US										
nd	.32**	.02	.34**	.09	-.32**	-.16	.01	-.21	-.21**	-.00
s	.01	-.04	.34**	.05	-.20	.27**	-.06	-.18	.18	-.02
JA										
nd	.64**	.54**	.38**	.34**	.17	.13	.07	.17	.07	.12
s	.67**	.56**	.51**	.31**	.11	.03	-.06	-.13	-.25	-.27
FR										
nd	-.37**	-.18	.35**	-.18	-.04	.05	-.09	.03	-.02	-.03
s	-.30**	-.11**	-.07	.08	.06	.10	-.01	-.12	.02	-.03
UK										
nd	-.02	-.06	.01	-.05	.05	.12	.05	-.20	-.11	-.00
s	.19	.20	.29**	.15	.29**	.05	.11	.13	.05	.11
IT										
nd	.67**	.31**	-.14	-.27	-.37	-.21	-.14	-.01	-.12	-.10
s	.39**	.10	-.16**	.17	-.13	-.11	-.23	.09	-.19**	-.10
CA										
nd	-.15	.00	.09	.04	.15	.11	-.07	.03	.09	.16**
s	.17	.17	.12	.27**	.08	.02	.01	-.07	.07**	.01

NOTE: The sample period is 1971:II-1988:I
n: order of autocorrelation.
**: significantly different from zero at 5% level.
Estimates of the standard deviation of autocorrelations were obtained using the Newey & West (1987) method, allowing for 10 lags.

TABLE 5.-- AUGMENTED DICKEY-FULLER TESTS FOR UNIT ROOTS IN QUARTERLY LOG CONSUMPTION

	k=0	k=1	k=2	k=3	k=4	k=5	k=6
US							
nd	-1.58	-2.29 §	-2.12	-3.65 **	-3.07 ‡	-2.55 §	-2.33
s	-1.68	-1.78	-1.80	-2.32 §	-2.27	-1.92	-2.35
c	-1.36	-1.98	-1.97	-3.32 *	-2.97 ‡	-2.23 §	-2.33
JA							
nd	-2.80 ‡	-3.07 ‡	-3.85 **	-3.80 **	-3.83 **	-3.11 ‡	-3.01 ‡
s	-2.84 ‡	-2.80 §	-3.94 **	-5.02 ***	-4.85 ***	-4.24 ***	-3.69 **
c	-3.26 *	-3.07 ‡	-4.03 **	-4.79 ***	-4.44 ***	-3.66 **	-3.13 ‡
FR							
nd	-4.62 ***	-3.60 **	-2.74 §	-3.13 ‡	-3.16 ‡	-3.23 *	-2.85 ‡
s	-2.98 ‡	-2.65 §	-1.35	-1.35	-1.54	-1.90	-2.07
c	-3.28 *	-2.86 ‡	-2.49	-2.38 §	-2.29	-2.47 §	-2.29
UK							
nd	-2.56	-2.53 §	-2.68 §	-2.64 §	-2.38	-2.51 §	-2.97 ‡
s	1.68	1.42	1.06	0.46	0.32	-0.22	-0.07
c	0.41	-0.08	-0.38	-0.54	-0.29	-1.12	-1.18
IT							
nd	-1.29	-3.32 *	-2.86 ‡	-1.80	-2.70 §	-1.77	-3.02 ‡
s	-1.80	-2.75 §	-2.36	-2.36 §	-3.51 **	-2.38 §	-2.88 ‡
c	-1.38	-3.33 *	-2.80 ‡	-2.11	-2.78 §	-2.08	-2.48
CA							
nd	-3.02 ‡	-2.61 §	-2.53	-2.28 §	-2.22	-2.24 §	-2.86 ‡
s	-1.62	-1.72	-1.99	-1.99	-2.19	-2.47 §	-2.40 §
c	-2.32	-2.01	-2.11	-2.00	-1.99	-2.34 §	-2.86 ‡

NOTE: The sample period is 1971:II-1988:I.
nd: per capita consumption of non-durables.
s: per capita consumption of services.
c: per capita consumption of non-durables plus services.
The test results are for logs of the variables indicated in the first column.
k: the number of lagged Δz terms included on right-hand side of (2.2); 1971:II-1988:I is the sample period used for k=0; for k>0, the beginning of the sample period is shifted forward, e.g., for k=1, the sample period is 1971:III-1988:I.

***, **, *, ‡ and § indicate rejection of null-hypothesis $\phi=0$ at 1%, 5%, 10%, 20% and 50% significance levels respectively.

For 50 observations, the critical values for the Augmented Dickey-Fuller test are -4.15, -3.50, -3.18, -2.81 and -2.60 at the 1%, 5%, 10%, 20% and 50% levels respectively (the 1%, 5% and 10% critical values are taken from table 8.5.2 of Fuller (1976); The 20% and 50% critical values were obtained by Monte Carlo simulations (5000 replications) of random walks with N(0,1) innovations).

TABLE 6.-- AUGMENTED DICKEY-FULLER TESTS FOR UNIT ROOTS IN BILATERAL REAL EXCHANGE RATES

	k=0	k=1	k=2	k=3	k=4	k=5	k=6
US-JA	-1.45	-2.36	-2.14	-2.02	-1.67	-1.62	-1.44
US-FR	-1.53	-2.10	-1.95	-2.03	-1.89	-2.17	-2.25
US-UK	-1.54	-1.94	-1.83	-1.98	-2.14	-1.61	-2.21
US-IT	-0.66	-1.82	-1.92	-2.04	-1.72	-1.98	-1.79
US-CA	-0.87	-1.36	-1.60	-2.02	-2.21	-2.32	-2.16
JA-FR	-2.12	-3.41 *	-3.26 *	-3.71 **	-2.69 §	-2.40	-2.39
JA-UK	-1.62	-2.53	-2.34	-2.36	-2.04	-1.82	-1.73
JA-IT	-2.15	-3.55 **	-3.04 ‡	-3.66 **	-2.61 §	-2.26	-2.16
JA-CA	-1.52	-2.35	-2.16	-2.14	-1.77	-1.94	-1.80
FR-UK	-2.13	-2.23	-2.33	-2.60 §	-2.46	-2.46	-2.51
FR-IT	-1.76	-1.89	-1.70	-1.94	-1.56	-1.85	-2.00
FR-CA	-1.49	-2.09	-1.93	-2.13	-1.90	-2.42	-2.48
UK-IT	-1.42	-1.70	-1.24	-1.41	-1.83	-1.08	-1.33
UK-CA	-1.43	-1.84	-1.82	-1.98	-2.12	-1.72	-2.21
IT-CA	-1.04	-2.16	-2.18	-2.42	-2.07	-2.49	-2.35

NOTE: The sample period is 1971:II-1988:I.
The bilateral real exchange rates are defined in terms of non-durables.

k: the number of lagged Δz terms included on right-hand side of (2.2); 1971:II-1988:I is the sample period used for k=0; for k>0, the beginning of the sample period is shifted forward, e.g., for k=1, the sample period is 71:III-88:1.

***, **, *, ‡ and § indicate rejection of null-hypothesis $\phi=0$ at 1%, 5%, 10%, 20% and 50% significance levels respectively.

For 50 observations, the critical values for the Augmented Dickey-Fuller test are -4.15, -3.50, -3.18, -2.81 and -2.60 at the 1%, 5%, 10%, 20% and 50% levels respectively (the 1%, 5% and 10% critical values are taken from table 8.5.2 of Fuller (1976); The 20% and 50% critical values were obtained by Monte Carlo simulations (5000 replications) of random walks with N(0,1) innovations).

TABLE 7.-- A MONTE CARLO STUDY ON THE PARK TEST STATISTIC

	n=0	n=5	n=10	n=20	n=30	n=50	n=60
ρ=0	.13	.07	.18	.44	.67	.86	.97
ρ=0.25	.38	.09	.10	.37	.62	.92	.96
ρ=0.50	.49	.13	.22	.32	.68	.97	.94
R $\rho\rho$=0795	.786	.250	.199	.449	.725	.999	.987
ρ=1	.97	.64	.47	.57	.83	.99	1.00

NOTE:
The table shows the proportions of draws in which, using the 10% critical values for the asymptotic distribution of the Park statistic, the hypothesis is rejected that x and y are cointegrated. 'n' is the lag length used to correct for serial correlation in constructing the Park test statistic (t, t^2 and t^3 are used as superfluous regressors).

The table is based on the following process:

$x_t=x_{t-1}+\varepsilon_t$; $y_t=x_t+h_t$, $h_t=\rho*h_{t-1}+\eta_1$ for t=1,...,70 (and $x_0=h_0=0$),
where ε and η are N(0,1) white noises. Hence {x} follows a random walk; {x} and {y} are cointegrated if $|\rho|<1$. 100 draws of the {x} and {y} processes were simulated for ρ=0, 0.25, 0.50, 0.75, 0.95 and 1.

TABLE 8.-- PARK (1990) COINTEGRATION TESTS

	(1) Single Good Model		(2) One Country-specific Consumption Good		(3) Two Country-specific Consumption Goods	
US-JA	.33		.69	(P)	.00	
	.02		.02	(P)	.00	
			.05	(P)	.21	(P)
					.08	
					.32	
US-FR	.41		.61		.36	(P)
	.08		.27	(P)	.10	
			.03	(P)	.70	(P)
					.38	(P)
					.26	(P)
US-UK	.38		.63		.34	(P)
	.04		.03		.02	
			.09		.12	(P)
					.00	
					.03	(P)
US-IT	.24	(P)	.35	(P)	.30	(P)
	.08	(P)	.21	(P)	.43	
			.03	(P)	.26	
					.55	(P)
					.22	(P)
US-CA	.82		.26	(P)	.21	(P)
	.04		.02	(P)	.07	(P)
			.13	(P)	.03	(P)
					.12	(P)
					.12	(P)
JA-FR	.04		.03		.17	(P)
	.17		.16		.12	(P)
			.16		.15	(P)
					.20	(P)
					.23	(P)
JA-UK	.21		.03		.00	(P)
	.03	(P)	.02	(P)	.67	(P)
			.11	(P)	.10	(P)
					.00	
					.19	
JA-IT	.02		.02	(P)	.42	(P)
	.10		.11		.06	(P)
			.27	(P)	.19	(P)
					.04	
					.68	(P)

TABLE 8.-- continued

	(1) Single Good Model		(2) One Country-specific Consumption Good		(3) Two Country-specific Consumption Goods	
JA-CA	.14		.04		.15	
	.09		.29		.03	
			.73		.07	
					.69	
					.46	
FR-UK	.15		.54	(P)	.11	
	.07		.08	(P)	.84	(P)
			.09	(P)	.40	
					.04	
					.09	
FR-IT	.21		.23	(P)	.57	(P)
	.66		.72		.05	(P)
			.09	(P)	.02	(P)
					.25	(P)
					.11	(P)
FR-CA	.81		.82	(P)	.55	(P)
	.12		.11		.16	
			.24	(P)	.55	(P)
					.61	(P)
					.43	(P)
UK-IT	.01	(P)	.04		.61	
	.28		.06		.99	(P)
			.13		.02	
					.14	
					.18	
UK-CA	.17		.05		.40	
	.07		.02		.30	(P)
			.11	(P)	.01	
					.86	(P)
					.94	(P)
IT-CA	.87		.69		.46	
	.15		.10		.10	(P)
			.03	(P)	.63	(P)
					.10	(P)
					.16	(P)

NOTE: Sample period: 1971:II-1988:I (quarterly data).
The table reports p-values for test of null hypothesis of cointegration.
(1) Tests of the single good model: test whether $\ln(c^i)$ and $\ln(c^j)$ are cointegrated (see (2.5)); c: non-durables plus services.

TABLE 8.-- continued

(2) Tests of the model with one country-specific consumption good: test whether $\ln(c^i)$, $\ln(c^j)$ and $\ln(R^{i,j})$ are cointegreted (see (2.7)); c: non-durables plus services, $R^{i,j}$: real exchange rate in terms of 'c'.

(3) Tests of the model with two country-specific consumption goods: test whether $\ln(nd^i)$, $\ln(s^i)$, $\ln(nd^j)$, $\ln(s^j)$ and $\ln(RND^{i,j})$ are cointegrated (see (2.10 a)); nd: non-durables, s: services, $RND^{i,j}$: real exchange rate in terms of non-durables.

As the outcome of the Park test can depend on which 'x' variable is used on the left-hand side in the cointegrating regression (2.3), test results are reported for all possible choices for the left-hand side variable. The first two columns in the table indicate the country pair to which the test statistics provided in the remaining columns pertain. Let i and j denote the first and the second country listed for a given country pair (e.g. for US-JA, i=US and j=Japan). For the single good model, the first line of test statistic provided for a given country pair uses $\ln(c^i)$ on the left-hand side of (2.3), while the second line uses $\ln(c^j)$ as the left-hand side variable. For the model with one country-specific good, $\ln(c^i)$ and $\ln(c^j)$ and $\ln(R^{i,j})$ are used as left-hand side variables for the 1st, 2nd and 3rd test statistics respectively. For the model with two country - specific goods, $\ln(nd^i)$, $\ln(nd^j)$, $\ln(s^i)$, $\ln(s^j)$ and $\ln(R^{i,j})$ are used as left-hand side variables for the 1st, 2nd, 3rd and 4th test statistic respectively.

For the single good model, '(P)' indicates that coefficients of canonical cointegrating regressions (3.4) imply that the restriction $(\sigma^i-1)/(\sigma^j-1)>0$ is rejected *and* that one fails to reject the hypothesis that $(\sigma^i-1)/(\sigma^j-1)=0$ at 5% level.

For model with one country-specific consumption good, '(P)' indicates that preference parameters recovered from coefficients of canonical cointegrating regressions violate the restriction $\sigma^k<1$ *and* that one fails to reject the hypothesis that $\sigma^k-1=0$ at the 5% significance level for k=i and/or k=j.

For the model with two country-specific consumption goods, '(P)' indicates that preference parameters recovered from coefficients of canonical cointegrating regressions violate one of the following restrictions: $\sigma^i+\mu^i<1$, $\sigma^i*\mu^i>0$, $\sigma^j+\mu^j<1$, $\sigma^j*\mu^j>0$ *and* that in addition this violation is 'significant', (when $\sigma+\mu<1$ is violated, then '(P)' indicates that the hypothesis $\sigma+\mu-1=0$ is rejected at the 5% level, when $\sigma*\mu>0$ is violated, then '(P)' indicates that the hypothesis $\sigma*\mu=0$ is rejected at the 5% level).

TABLE 9.-- CRITICAL VALUES FOR THE PHILLIPS & OULIARIS (1990) $\hat{Z}_{\alpha}$ AND $\hat{Z}_{t}$ TEST STATISTICS

(a) $\hat{Z}$ statistic

Size	0.01	0.05	0.10	0.20	0.50	0.80
n=1	-35.41	-27.08	-23.19	-18.69	-12.29	-7.36
n=2	-40.34	-32.22	-27.78	-23.36	-15.89	-9.69
n=4	-53.61	-42.45	-37.73	-32.58	-23.76	-16.69

(b) $\hat{Z}_{t}$ statistic

Size	0.01	0.05	0.10	0.20	0.50	0.80
n=1	-4.36	-3.80	-3.51	-3.15	-2.53	-1.90
n=2	-4.64	-4.15	-3.84	-3.50	-2.85	-2.17
n=4	-5.36	-4.74	-4.46	-4.13	-3.50	-2.90

NOTE:
n is the number of x-variables on the right-hand side of the cointegrating regression (3.3).

The null-hypothesis of *no* cointegration is rejected if the value of the test statistic falls *below* the chosen critical value.

Phillips & Ouliaris provide 1%, 5% and 10% critical values for the $\hat{Z}_{\alpha}$ and $\hat{Z}_{t}$ test statistics (see tables Ic and IIc of their paper). I obtained 20%, 50% and 80% critical values using Monte Carlo simulations.

TABLE 10.-- PHILLIPS & OULIARIS (1990) TESTS OF NO-COINTEGRATION HYPOTHESIS

	(1) Single Good Model		(2) One Country-specific Consumption Good		(3) Two Country-specific Consumption Goods	
	$\hat{Z}_\alpha$	$\hat{Z}_t$	$\hat{Z}_\alpha$	$\hat{Z}_t$	$\hat{Z}_\alpha$	$\hat{Z}_t$
US-JA	-11.25	-2.95 §	-11.16	-2.81	-13.10	-2.67
	-8.72	-2.04	-11.39	-2.42	-8.37	-1.88
			-9.70	-2.03	-19.12	-3.46
					-12.98	-2.66
					-13.01	-2.60
US-FR	-30.69 *	-4.37 *	-43.74 ***	-5.57 ***	-22.70	-3.80 §
	-13.90 §	-2.69 §	-13.45	-2.64	-56.94 ***	-6.97 ***
			-22.49 §	-3.51 ‡	-20.48	-3.75 §
					-31.85 §	-4.53 *
					-15.51	-2.92
US-UK	-6.11	-1.26	-5.08	-1.10	-15.05	-2.95
	-9.24	-2.14	-9.20	-2.13	-21.37	-4.02 §
			-7.26	-1.93	-23.98 §	-3.91 §
					-6.66	-1.23
					-13.66	-2.75
US-IT	-6.80	-1.87	-5.36	-1.54	-10.57	-2.61
	-8.69	-2.06	-9.54	-2.18	- 9.68	-2.17
			-8.00	-1.70	-13.92	-3.02
					-10.97	-2.46
					-8.61	-1.83
US-CA	-8.33	-2.50	-8.02	-2.30	-29.30 §	-4.53 *
	-9.47	-2.17	-8.82	-2.03	-60.22 *	-6.88 ***
			-9.21	-2.01	-23.92 §	-4.35 ‡
					-43.91 **	-5.52 ***
					-32.24 §	-4.12 §
JA-FR	-23.22 *	-3.64 *	-24.55 ‡	-3.73 ‡	-22.83	-3.52 §
	-11.82	-2.79 §	-13.61	-3.08 §	-54.02 ***	-6.75 ***
			-11.60	-2.58	-31.45 §	-4.65 *
					-42.94 **	-5.43 ***
					-15.97	-3.41
JA-UK	-5.06	-1.07	-6.00	-1.24	-12.65	-2.37
	-11.04	-2.93 §	-11.04	-2.92 §	-29.78 §	-4.27 ‡
			-8.36	-2.13	-12.73	-2.43
					-14.00	-2.21
					-11.57	-2.59
JA-IT	-11.62	-2.58 §	-11.26	-2.52	-10.58	-2.40
	-14.94 §	-3.44 ‡	-14.75	-3.26	-15.19	-3.19
			-8.06	-2.08	-20.66	-3.78 §
					-13.69	-2.68
					-9.08	-2.77

TABLE 10.-- continued

	(1) Single good model		(2) One country-specific consumption good		(3) Two country-specific consumption goods	
	$\hat{Z}_{\alpha}$	$\hat{Z}_{t}$	$\hat{Z}_{\alpha}$	$\hat{Z}_{t}$	$\hat{Z}_{\alpha}$	$\hat{Z}_{t}$
JA-CA	-7.25	-1.87	-11.23	-2.44	-9.65	-1.99
	-11.19	-2.50	-11.40	-2.53	-33.11 ‡	-4.60 *
			-11.48	-2.49	-14.21	-2.82
					-21.61	-3.68 §
					-11.99	-2.65
FR-UK	-3.85	-0.84	-5.30	-1.12	-43.56 **	-6.36 ***
	-18.85 ‡	-3.35 ‡	-25.23 ‡	-3.83 ‡	-28.80 §	-4.48 *
			-11.43	-2.49	-31.20 §	-4.32 ‡
					-11.52	-1.95
					-14.65	-2.83
FR-IT	-8.86	-2.13	-8.90	-2.15	-41.52 *	-6.09 ***
	-22.38 ‡	-3.72 *	-36.47 **	-4.75 ***	-10.57	-2.35
			-18.09 §	-3.18 §	-49.35 **	-6.06 ***
					-16.93	-3.09
					-30.80 §	-4.25 ‡
FR-CA	-20.72 ‡	-3.54 *	-12.23	-2.80	-39.38 *	-5.67 ***
	-32.35 **	-4.61 *	-36.79 **	-5.11 ***	-31.89 §	-4.59 *
			-16.40 §	-3.08 §	-39.91 *	-5.28 **
					-31.98 §	-4.75 **
					-17.01	-3.05
UK-IT	-6.67	-1.84	-7.25	-1.91	-20.86	-3.67 §
	-5.14	-1.09	-10.17	-1.85	-13.24	-2.74
			-12.61	-2.59	-6.08	-1.38
					-10.86	-2.41
					-14.85	-2.90
UK-CA	-7.33	-2.37	-8.27	-2.53	-18.75	-3.42
	-4.90	-1.06	-4.20	-0.93	-39.26 *	-5.48 ***
			-8.64	-2.13	-8.20	-1.51
					-33.95 ‡	-5.09 **
					-23.06	-3.76 §
IT-CA	-9.37	-2.57 §	-9.47	-2.65	-9.76	-2.19
	-8.47	-2.09	-8.19	-2.04	-37.31 ‡	-5.13 **
			-8.66	-2.00	-11.79	-2.46
					-32.55 §	-4.38 ‡
					-12.55	-2.41

TABLE 10.-- continued

NOTE:
Sample period is 1971:II-1988:I (quarterly data).

***, **, *, ÷, §, £: rejection of null hypothesis of no cointegration at 1%, 5%, 10%, 20%, 50% and 80% levels respectively.

Columns labeled '(1) Single Good Model': tests of hypothesis that $\ln(c^i)$ and $\ln(c^j)$ are not cointegrated (c: non-durables plus services)

Columns labeled '(2) Model With One Country-specific Consumption Good': tests of hypothesis that $\ln(c^i)$, $\ln(c^j)$ and $\ln(R^{i,j})$ are not cointegreted (c: non-durables plus services).

Columns labeled '(3) Model With Two Country-specific Consumption Goods': tests of hypothesis that $\ln(nd^i)$, $\ln(s^i)$, $\ln(nd^j)$, $\ln(s^j)$ and $\ln(RND^{i,j})$ are not cointegrated (nd: non-durables; s: services).

As the outcome of the Park test can depend on which 'x' variable is used on the left-hand side in the cointegrating regression (2.3), test results are reported for all possible choices for the left-hand side variable. The first two columns in the table indicate the country pair to which the test statistics provided in the remaining columns pertain. Let i and j denote the first and the second country listed for a given country pair (e.g. for US-JA, i=US and j=Japan). For the single good model, the first line of test statistic provided for a given country pair uses $\ln(c^i)$ on the left-hand side of (2.3), while the second line uses $\ln(c^j)$ as the left-hand side variable. For the model with one country-specific good, $\ln(c^i)$ and $\ln(c^j)$ and $\ln(R^{i,j})$ are used as left-hand side variables for the 1st, 2nd and 3rd test statistics respectively. For the model with two country - specific goods, $\ln(nd^i)$, $\ln(nd^j)$, $\ln(s^i)$, $\ln(s^j)$ and $\ln(R^{i,j})$ are used as left-hand side variables for the 1st, 2nd, 3rd and 4th test statistic respectively.

TABLE 11.-- REGRESSIONS OF CONSUMPTION GROWTH RATES ON LAGGED CONSUMPTION GROWTH RATES

	$\bar{R}^2$		$\hat{\rho}$			r		
US-JA	.07	.35	.03	(-.45	.51)	.16	(-.08	.38)
US-FR	.03	.06	.74	(-.13	.96)	.30	(.05	.51)
US-UK	.06	.02	.11	(-.71	.80)	.27	(.01	.50)
US-IT	.28	.27	.27	(-.22	.65)	-.04	(-.27	.18)
US-CA	.04	.13	.37	(-.25	.78)	.06	(-.17	.30)
JA-FR	.32	.03	.23	(-.31	.66)	.22	(.18	.42)
JA-UK	.29	.00	.57	(-.33	.93)	.24	(.41	.05)
JA-IT	.32	.39	.64	(.00	.90)	.27	(.01	.49)
JA-CA	.28	.10	.53	(-.03	.84)	.11	(-.09	.31)
FR-UK	.14	.00	.33	(-.31	.77)	.25	(.08	.40)
FR-IT	.07	.26	-.01	(-.39	.36)	.00	(-.17	.17)
FR-CA	.02	.01	.43	(-.57	.91)	.13	(-.12	.37)
UK-IT	.03	.34	.68	(.12	.91)	.19	(.04	.34)
UK-CA	.00	.07	.10	(-.56	.68)	.20	(-.07	.46)
IT-CA	.27	-.01	.27	(-.47	.79)	.30	(.07	.49)

NOTE:
Regressions use quarterly data for the period 1971:IV-19881:I.
For each country pair i,j, the table considers regressions of $\Delta\ln(c_t^i)$ and $\Delta\ln(c_t^j)$ on a constant, $\Delta\ln(nd^i)$, $\Delta\ln(s^i)$, $\Delta\ln(nd^j)$, $\Delta\ln(s^j)$ lagged two and three periods ($c^i = nd^i + s^i$; nd: non-durables; s: services).

$\bar{R}^2$: Adjusted R^2. The first (second) $\bar{R}^2$ statistic reported for a given pair of countries uses as the left-hand side variable the consumption growth rate of the first (second) country listed for that pair.

$\hat{\rho}$: Sample correlation coefficient between fitted values of $\Delta\ln(c_t^i)$ and $\Delta\ln(c_t^j)$ from regressions.
(): 95% confidence intervals (calculated using the methods of Cumby & Huizinga (1991) and Newey & West (1987), allowing for 10 autocorrelations).

r: cross-country correlation of actual consumption growth rates.
(): 95% confidence intervals (calculated using the method of Newey & West (1987), allowing for 10 autocorrelations).

TABLE 12.-- VELU ET AL. (1986) TESTS: INCOMPLETE ASSET MARKETS MODEL WITH A SINGLE CONSUMPTION GOOD

(a) p-values for Test of the Hypothesis That rank(Ψ)=n-1

	h=0	h=1	h=2	h=3	h=4
US-JA	.00	.03	.00	.00	.00
US-FR	.06	.09	.11	.19	.04
US-UK	.45	.44	.32	.46	.05
US-IT	.00	.01	.00	.00	.00
US-CA	.08	.24	.14	.12	.07
JA-FR	.00	.01	.00	.00	.00
JA-UK	.25	.49	.77	.84	.31
JA-IT	.00	.00	.00	.00	.00
JA-CA	.03	.24	.04	.12	.16
FR-UK	.10	.11	.19	.39	.18
FR-IT	.00	.00	.00	.00	.00
FR-CA	.03	.15	.02	.06	.09
UK-IT	.12	.34	.59	.69	.38
UK-CA	.42	.69	.19	.20	.06
IT-CA	.08	.37	.10	.27	.04

(b) Estimates of $(\sigma^i-1)/(\sigma^j-1)$

	h=0	h=1	h=2	h=3	h=4
US-JA	12.60 (29.72)	37.24 (251.20)	52.50 (600.99)	-21.00 (100.27)	-4.14 (5.63)
US-FR	-1.80 (1.00)	-2.77 (2.15)	-1.66 (0.55)	-1.59 (0.51)	-1.28 (0.45)
US-UK	-1.45 (0.80)	-1.59 (1.24)	-0.64 (0.52)	-0.54 (0.47)	0.00 (0.65)
US-IT	3.31 (3.50)	6.41 (12.53)	-2.27 (2.10)	-3.69 (5.65)	-6.63 (16.97)
US-CA	-0.97 (1.07)	-1.22 (1.57)	-0.81 (0.50)	-0.61 (0.51)	-0.31 (0.44)
JA-FR	0.14 (0.26)	0.23 (0.27)	0.11 (0.26)	-0.02 (0.25)	-0.19 (0.23)
JA-UK	-0.41 (0.27)	-0.45 (0.27)	-0.44 (0.27)	-0.49 (0.26)	-0.64 (0.29)
JA-IT	-0.66 (0.24)	-0.76 (0.22)	-0.78 (0.20)	-0.78 (0.20)	-0.81 (0.18)
JA-CA	-0.38 (0.21)	-0.37 (0.21)	-0.50 (0.23)	-0.48 (0.22)	-0.48 (0.22)
FR-UK	-0.48 (0.61)	0.35 (0.66)	0.08 (0.48)	0.01 (0.48)	0.33 (0.85)
FR-IT	2.31 (2.26)	3.82 (6.51)	-4.04 (7.79)	12.46 (64.57)	-25.29(137.5)
FR-CA	0.41 (0.70)	0.17 (0.42)	-0.03 (0.63)	0.01 (0.68)	-0.09 (0.85)
UK-IT	-1.46 (1.03)	-1.16 (0.60)	-0.96 (0.40)	-0.98 (0.40)	-0.87 (0.31)
UK-CA	-0.81 (0.58)	-0.72 (0.58)	-3.26 (6.64)	-25.29(507.61)	-9.32(219.75)
IT-CA	-0.47 (0.33)	-0.42 (0.30)	-0.37 (0.31)	-0.49 (0.30)	-0.68 (0.30)

NOTE: Standard Deviations in Parentheses.
Sample period used for h=0 is 1971:II-1988:I. For h>0, the beginning of the sample period is shifted forward, e.g., for h=1, the sample period is 1971:III-1988:I.
h: the Velu et al. (1986) test uses lagged values of the nx1 vector X_{t+1} as instruments: $X_{t+1}=\Psi' Z_t+\epsilon_{t+1}$, where $Z_t = (X_t', X_{t-1}', \ldots, X_{t-h}')'$ and

TABLE 12.-- continued

$X_{t+1}=(\Delta\ln(c^{i}_{t+1}),\ \Delta\ln(c^{j}_{t+1}))'$.

The Velu et al. method requires the vector X to have zero mean. As the X variables used in the empirical work do not have zero means, the analysis was conducted using deviations of all variables from their respective sample means. This does not affect the asymptotic properties of the Velu et al. test statistics.

'i' denotes the first country listed for a given pair of countries.

TABLE 13.-- VELU ET AL. (1986) TESTS: INCOMPLETE ASSET MARKETS MODEL WITH ONE COUNTRY-SPECIFIC CONSUMPTION GOOD

(a) p-values for Test of the Hypothesis That rank(Ψ)=n-1

	h=0	h=1	h=2	h=3	h=4
US-JA	.17	.40	.17	.17	.29
US-FR	.06	.08	.30	.07	.06
US-UK	.61	.55	.80	.89	.75
US-IT	.00	.00	.01	.02	.05
US-CA	.07	.32	.25	.25	.35
JA-FR	.02	.19	.33	.24	.00
JA-UK	.30	.72	.88	.92	.49
JA-IT	.00	.00	.01	.00	.00
JA-CA	.11	.53	.34	.47	.61
FR-UK	.62	.27	.41	.62	.40
FR-IT	.00	.01	.00	.00	.00
FR-CA	.06	.05	.03	.05	.09
UK-IT	.31	.45	.76	.69	.62
UK-CA	.92	.91	.85	.92	.94
IT-CA	.08	.51	.22	.34	.14

(b) Estimates of Preference Parameters for h=4.

	σ^i	σ^j
US-JA	-5.63 (2.39)	-0.34 (2.04)
US-FR	-14.30 (13.76)	-7.44 (9.94)
US-UK	2.84 (2.64)	4.99 (2.72)
US-IT	-1.01 (2.19)	2.57 (2.05)
US-CA	1.02 (2.10)	-2.13 (3.57)
JA-FR	-0.16 (2.34)	7.87 (6.25)
JA-UK	8.81 (9.25)	13.06 (12.63)
JA-IT	-7.32 (5.76)	-7.50 (7.35)
JA-CA	-7.13 (6.22)	-12.43 (10.35)
FR-UK	-3.31 (7.08)	8.51 (19.00)
FR-IT	-2.57 (2.68)	1.06 (1.71)
FR-CA	-2.47 (11.06)	-14.97 (17.75)
UK-IT	8.92 (7.99)	10.13 (9.80)
UK-CA	3.31 (2.41)	-3.38 (2.83)
IT-CA	-5.49 (11.76)	-8.04 (12.13)

NOTE: Standard Deviations in Parentheses. Sample period used for h=0 is 1971:II-1988:I. For h>0, the beginning of the sample period is shifted forward, e.g., for h=1, the sample period is 1971:III-1988:I.

TABLE 13.-- continued

h: the Velu et al. (1986) test uses lagged values of the nx1 vector X_{t+1} as instruments: $X_{t+1}=\Psi' Z_t+\epsilon_{t+1}$, where $Z_t = (X_t', X_{t-1}', \ldots, X_{t-h}')'$, where $X_{t+1}=(\Delta\ln(c^i_{t+1}), \Delta\ln(c^j_{t+1}), \Delta\ln(R^{i,j}_t))'$.

The Velu et al. method requires the vector X to have zero mean. As the X variables used in the empirical work do not have zero means, the analysis was conducted using deviations of all variables from their respective sample means. This does not affect the asymptotic properties of the Velu et al. test statistics.
'i' denotes the first country listed for a given pair of countries.

TABLE 14.--GMM TESTS OF THE INCOMPLETE ASSET MARKETS MODEL WITH A SINGLE CONSUMPTION GOOD

	$\hat{\bar{\gamma}}$		J	p-value	J'	p-value
US-JA	0.33	(0.16)	10.82	.69	18.28	.95
	0.71	(0.17)	13.57	.48		
US-FR	0.68	(0.17)	19.36	.15	19.46	.92
	0.39	(0.14)	9.53	.79		
US-UK	-0.02	(0.09)	12.56	.56	19.23	.93
	0.52	(0.21)	14.13	.44		
US-IT	-0.01	(0.23)	14.49	.41	18.93	.94
	0.06	(0.13)	15.42	.34		
US-CA	-0.14	(0.21)	13.00	.52	19.11	.93
	-0.28	(0.22)	11.26	.66		
JA-FR	0.63	(0.22)	9.02	.82	17.29	.96
	0.25	(0.12)	10.94	.69		
JA-UK	0.22	(0.11)	11.10	.67	18.48	.95
	1.05	(0.25)	7.60	.90		
JA-IT	1.66	(0.32)	11.91	.61	18.55	.94
	0.37	(0.07)	10.83	.69		
JA-CA	0.85	(0.19)	9.90	.76	16.32	.97
	0.62	(0.15)	9.82	.77		
FR-UK	-0.06	(0.10)	13.98	.45	18.50	.94
	-0.36	(0.25)	12.72	.54		
FR-IT	0.47	(0.25)	12.63	.55	17.39	.96
	0.23	(0.09)	13.28	.50		
FR-CA	0.22	(0.11)	9.77	.77	17.50	.96
	0.51	(0.19)	10.79	.70		
UK-IT	-0.29	(1.00)	7.12	.92	19.48	.92
	0.01	(0.08)	13.42	.49		
UK-CA	-0.44	(0.40)	7.91	.89	18.42	.95
	-0.14	(0.21)	9.90	.76		
IT-CA	0.37	(0.05)	11.37	.65	17.62	.96
	1.79	(0.33)	10.60	.71		

TABLE 14.-- continued

NOTE: The sample period is 71:IV-88:I (quarterly data).
The table reports tests of the model with incomplete asset markets and a single consumption good. The tests are use the Generalized Method of Moments, exploiting the orthogonality condition $E\gamma'*X_{t+1}*Z_t=0$. The table tests condition (7.5'):

$$\Delta\ln(c^k_{t+1})=\tilde{\mu}+\tilde{\gamma}*\Delta\ln(c^h_{t+1})+\tilde{\eta}_{t+1}, \tag{A.1}$$

where $\tilde{\gamma}\equiv(\sigma^k-1)/(\sigma^h-1)$ and $\tilde{\eta}_{t+1}$ has mean zero and is orthogonal to all instruments in the period t information set.

The following instruments are used: $\Delta\ln(nd^k)$, $\Delta\ln(nd^h)$, $\Delta\ln(s^k)$, $\Delta\ln(s^h)$, r^k_{nd}, r^h_{nd}, r^k_s, r^h_s, lagged two and three periods. Here r_{nd} and r_s are ex post real short term interest rates in terms of non-durables and services respectively.
The first column in the table indicates the country pair. Let i and j denote the first and the second country listed for a given country pair (e.g. for US-JA, i=US and j=Japan). For a given country pair, the first line with test statistics uses $\Delta\ln(c^i)$ on the left-hand side of (A.1), while the second line uses $\Delta\ln(c^j)$ on the left-hand side of (A.1).

$\hat{\tilde{\gamma}}$ is the GMM estimate of γ (standard deviation in parentheses).

J: Hansen's (1982) J-statistic for test of the hypothesis that the residual $\tilde{\eta}_{t+1}$ (see (A.1)) is orthogonal to the instruments. If the orthogonality conditions hold , the J-statistic is asymptotically χ^2 with 14 degrees of freedom. 'p-value' is the p-value of the J statistic.

J': This statistic tests the joint hypothesis that consumption growth rates for a given pair of countries are orthogonal to instruments.
Specifically, the statistic tests whether $\omega^k_{t+1}\equiv(\Delta\ln(c^k_{t+1})-\mu^k)$ and $\omega^h_{t+1}\equiv(\Delta\ln(c^h_{t+1})-\mu^h)$ are orthogonal to the instruments for some constants μ^k and μ^h. Under the hypothesis that ω^k_{t+1} and ω^h_{t+1} are orthogonal to the instruments, the J' statistic is asymptotically χ^2 with 30 degrees of freedom.

The real interest rates which are used as instruments are computed on the basis of short term interest rates from International Finacial Statistics (US: see line 60C in IFS, Japan: 60B, France: 60B, UK: 60C, Italy: 60B, Canada: 60C).

TABLE 15.-- CROSS-COUNTRY CORRELATIONS OF DETRENDED TOTAL PRIVATE CONSUMPTION

	JA	FR	UK	IT	CA
(a) Cross-country Correlations of Quarterly Growth Rates of Total Private Consumption.					
US	.22 (-.01 .44)	.27 (-.00 .50)	.24 (-.08 .51)	.02 (-.23 .28)	.26 (.07 .43)
JA		.25 (.04 .44)	.30 (.17 .43)	.55 (.34 .70)	.22 (-.03 .46)
FR			.21 (-.02 .43)	.17 (-.04 .37)	.23 (.02 .42)
UK				.10 (-.02 .21)	.17 (-.08 .40)
IT					.22 (-.03 .45)
(b) Cross-country Correlations of Linearly Detrended Logs of Quarterly Total Private Consumption.					
US	.12 (-.10 .33)	.19 (-.33 .62)	.52 (.05 .80)	-.18 (-.48 .17)	.39 (.02 .66)
JA		.59 (.21 .82)	.18 (-.10 .44)	.66 (.45 .80)	.58 (.17 .82)
FR			-.15 (-.41 .12)	.63 (.42 .77)	.64 (.23 .85)
UK				-.09 (-.37 .20)	.01 (-.37 .38)
IT					.38 (.04 .63)

NOTE: The figures in parentheses are 95% confidence intervals (based on estimates of the standard errors of the sample correlation coefficients which were obtained using the method of Newey & West (1987), allowing for 10 autocorrelations). Sample period is 1971:I-1988:I (1971:II-1988:I for statistics based on growth rates).

TABLE 16.-- CROSS-COUNTRY CORRELATIONS OF DETRENDED GNP

	JA	FR	UK	IT	CA
(a) Cross-country Correlations of Quarterly Growth Rates of GNP.					
US	.44 (.22 .62)	.29 (.01 .53)	.16 (-.02 .33)	.24 (.01 .45)	.47 (.21 .67)
JA		.37 (.21 .50)	.28 (-.04 .55)	.04 (-.18 .25)	.29 (.17 .40)
FR			.24 (.00 .45)	.56 (.33 .73)	.22 (-.04 .44)
UK				.03 (-.18 .24)	.18 (-.14 .47)
IT					.15 (-.10 .39)
(b) Cross-country Correlations of Linearly Detrended Logs of Quarterly GNP.					
US	.54 (.28 .72)	.25 (-.23 .64)	.71 (.58 .81)	.29 (.01 .53)	.60 (.22 .83)
JA		.14 (-.07 .33)	.36 (.07 .59)	.15 (-.14 .42)	.07 (-.22 .35)
FR			.21 (-.27 .61)	.83 (.73 .89)	.61 (.12 .86)
UK				.13 (-.20 .44)	.55 (.28 .74)
IT					.56 (.32 .73)

NOTE: The figures in parentheses are 95% confidence intervals (based on estimates of the standard errors of the sample correlation coefficients which were obtained using the method of Newey & West (1987), allowing for 10 autocorrelations). Sample period is 1971:I-1988:I (1971:II-1988:I for statistics based on growth rates).

TABLE 17.-- CROSS-COUNTRY CORRELATIONS OF DETRENDED GROSS FIXED CAPITAL FORMATION

(a) Cross-country Correlations of Quarterly Growth Rates of Gross Fixed Capital Formation.

	JA	FR	UK	IT	CA
US	.27 (.05 .47)	.16 (-.12 .43)	.12 (-.06 .31)	-.03 (-.28 .21)	.12 (-.25 .47)
JA		.46 (.29 .60)	.42 (.21 .60)	.49 (.31 .63)	.15 (-.03 .33)
FR			.39 (.05 .65)	.31 (.19 .42)	.35 (.18 .50)
UK				.09 (-.11 .28)	.11 (-.09 .32)
IT					.29 (.05 .49)

(b) Cross-country Correlations of Linearly Detrended Logs of Quarterly Gross Fixed Capital Formation.

	JA	FR	UK	IT	CA
US	.43 (.09 .68)	.26 (-.20 .63)	.41 (.00 .70)	.64 (.38 .80)	-.00 (-.49 .48)
JA		.57 (.13 .82)	.72 (.48 .86)	.45 (.03 .74)	.07 (-.31 .45)
FR			.74 (.46 .88)	.37 (-.08 .70)	.67 (.43 .82)
UK				.39 (-.09 .73)	.33 (.00 .60)
IT					.25 (-.32 .70)

NOTE: The figures in parentheses are 95% confidence intervals (based on estimates of the standard errors of the sample correlation coefficients which were obtained using the method of Newey & West (1987), allowing for 10 autocorrelations). Sample period is 1971:I-1988:I (1971:II-1988:I for statistics based on growth rates).

TABLE 18.-- WITHIN-COUNTRY CORRELATIONS OF DETRENDED TOTAL PRIVATE CONSUMPTION AND EMPLOYMENT

(a) Within-country Correlations Between Detrended Citibase/ILO Hours Measure and Total Private Consumption.

	growth rates	linearly detrended logs
US	.43 (.21 .60)	.69 (.47 .82)
JA	.36 (.04 .62)	-.15 (-.59 .36)
FR	.01 (-.19 .22)	-.42 (-.65 -.11)

(b) Within-country Correlations Between Employment Measure from International Financial Statistics (IFS) and Total Private Consumption.

	growth rates	linearly detrended logs
US	.38 (.19 .55)	.56 (.21 .78)
JA	.01 (-.23 .25)	-.45 (-.79 .08)
FR	.03 (-.20 .25)	.28 (-.07 .57)
IT	.12 (-.02 .27)	.70 (.42 .86)
CA	.20 (-.05 .43)	.70 (.54 .82)

NOTE: The figures in parentheses are 95% confidence intervals (based on estimates of the standard errors of the sample correlation coefficients which were obtained using the method of Newey & West (1987), allowing for 10 autocorrelations). Sample period is 1971:I-1988:I (1971:II-1988:I for statistics based on growth rates).

TABLE 19.-- CROSS-COUNTRY CORRELATIONS OF DETRENDED SOLOW RESIDUALS

(a) Cross-country Correlations of Growth Rates of Solow Residuals, Constructed Using Citibase/ILO Hours Data.

	JA	FR
US	.35 (.00)	.08 (.16)
JA		.19 (.00)

(b) Cross-country Correlations of Linearly Detrended Log Solow Residuals, Constructed Using Citibase/ILO Hours Data.

JA		.19 (.00)
	JA	FR
US	.54 (.00)	-.18 (.79)
JA		-.12 (.86)

(c) Cross-country Correlations of Growth Rates of Solow Residuals, Constructed Using IFS Employment Measure.

	JA	FR	IT	CA
US	.30 (.00)	.30 (.00)	.21 (.01)	.11 (.13)
JA		.32 (.00)	-.01 (.57)	-.02 (.60)
FR			.48 (.00)	.04 (.32)
IT				.26 (.05)

TABLE 19-- continued

(d) Cross-country Correlations of Linearly Detrended Log Solow Residuals, Constructed Using IFS Employment Measure.

	JA	FR	IT	CA
US	.16 (.24)	.26 (.02)	.45 (.00)	.60 (.00)
JA		.52 (.00)	.18 (.06)	-.20 (.88)
FR			.56 (.00)	-.13 (.88)
IT				.17 (.14)

NOTE: The figures in parentheses are p-value for one-sided test of null hypothesis that cross-country correlation is zero (the p-values are based on estimates of the standard errors of the sample correlation coefficients which were obtained using the method of Newey & West (1987), allowing for 10 autocorrelations). Sample period is 1971:I-1988:I (1971:II-1988:I for statistics based on growth rates).

Solow residual (S): $\ln(S)=\ln(GNP)-\eta*\ln(K)-(1-\eta)*\ln(L)$.
K: capital stock. L: hours. $\eta=0.35$ used in this table.

TABLE 20. TRIVARIATE AR(1) MODELS FITTED TO DETRENDED SOLOW RESIDUALS IN THE US, JAPAN AND FRANCE

(a) Trivariate AR(1) Models Fitted to Growth Rates of Solow Residuals.

$$RHO=\begin{bmatrix} 0.04 & 0.13 & -0.17 \\ (0.47) & (1.87) & (-1.35) \\ 0.08 & 0.26 & -0.14 \\ (0.60) & (1.44) & (-0.56) \\ -0.05 & 0.06 & -0.01 \\ (-0.47) & (0.75) & (-0.11) \end{bmatrix}$$

$$V=\begin{bmatrix} 5.07*10^{-5} & 2.13*10^{-5} & 0.34*10^{-5} \\ (6.56) & (1.83) & (0.74) \\ & 8.29*10^{-5} & 1.19*10^{-5} \\ & (2.40) & (3.04) \\ & & 4.40*10^{-5} \\ & & (9.67) \end{bmatrix}.$$

(b) Trivariate AR(1) Models Fitted to Linearly Detrended Logs of Solow Residuals.

$$RHO=\begin{bmatrix} 0.85 & 0.04 & -0.11 \\ (13.15) & (0.53) & (-1.98) \\ 0.03 & 0.83 & -0.16 \\ (0.58) & (9.06) & (-3.02) \\ 0.04 & 0.01 & 0.90 \\ (0.82) & (0.36) & (20.52) \end{bmatrix}$$

$$V=\begin{bmatrix} 4.74*10^{-5} & 2.03*10^{-5} & 0.30*10^{-5} \\ (6.59) & (1.67) & (0.71) \\ & 7.72*10^{-5} & 0.95*10^{-5} \\ & (2.05) & (2.13) \\ & & 4.00*10^{-5} \\ & & (11.06) \end{bmatrix}.$$

TABLE 20.-- continued

NOTE:
Sample period is 1971:II-1988:I (1971:III-1988:I for statistics based on growth rates).

The Solow residuals used for this table were constructed using the hours Citibase/ILO hours data, assuming an elasticity of output with respect to capital of η=0.35.

Let $X_t=(x_t^{US}, x_t^{JA}, x_t^{FR})'$ be the vector of detrended Solow residuals in the US, Japan and France.

The following model is estimated:

$$X_t=RHO*X_{t-1}+\epsilon_t, \tag{A.2}$$

where RHO is a 3x3 matrix and ϵ_t is a random variable of dimension 3x1 with covariance matrix $V=E\eta*\eta'$.

The tables report estimates of RHO (obtained using OLS for each equation in (A.2)) and of the covariance matrix of ϵ_t. t-statistics for these estimates are in parentheses.

As first differences of log Solow residuals have non-zero means in the data, (A.2) was estimated using deviations of the first differences from their respective sample means.

The adjusted R^2 coefficients for the AR(1) model fitted to growth rates of Solow residuals are 0.01, 0.05 and -0.02 in the first through third equations of the model and the Box-Pierce Q statistics (with 25 degrees of freedom) for these equations have the following p-values: 0.99, 0.98 and 0.84.
For the AR(1) model fitted to linearly detrended log Solow residuals, the corresponding R^2 statistics are 0.79, 0.73 and 0.82 respectively and the Box-Pierce Q statistics have the following p-values: 0.99, 0.94 and 0.79.

TABLE 21.-- NOTATION FOR TABLES WITH SIMULATION RESULTS

$q_t^i \equiv \theta_t^i * (K_t^i)^\eta * (L_t^i)^{1-\eta}$;

g_t^i: *additive* shock;

$y_t^i \equiv q_t^i - g_t^i$;

c_t^i: consumption of country 'i' with incomplete asset markets;

$\bar{c}_t^i$: consumption of country 'i' in presence of complete asset markets;

V_t^i: expected life-time utility of country i;

I_t^i: gross investment in country 'i';

TB_t^i: trade balance of country i (incomplete asset markets).

$\beta(u)$: β evaluated at steady state value of period utility function.

$(\partial\beta/\partial u)*(u/\beta)$: elasticity of β with respect to instantaneous utility, evaluated at the steady state;

d : depreciation rate of captal stock;

γ : g/y (g: steady state value of additive shock);

RHO_θ: autocorrelation matrix of multiplicative technology shocks;

RHO_g: autocorrelation matrix of additive shocks;

V_θ: covariance matrix of innovations to multiplicative technology shocks;

V_g: covariance matrix of innovations to additive shocks.

TABLE 22.-- SIMULATION RESULTS: FIXED LABOR SUPPLIES, MULTIPLICATIVE TECHNOLOGY SHOCKS. VARIATIONS IN THE INTERTEMPORAL ELASTICITY OF SUBSTITUTION

	σ=2	σ=3	σ=5	σ=7
std(c_t^i)	0.37	0.22	0.19	0.19
std($\bar{c}_t^i$)	0.29	0.17	0.14	0.13
std(I_t^i)	2.42	2.67	2.73	2.74
corr(c_t^i,c_t^j)	0.21	0.20	0.03	-0.07
corr(V_t^i,V_t^j)	-0.66	-0.72	-0.77	-0.80
corr(y_t^i,y_t^j)	0.00	0.00	0.01	0.21
corr(I_t^i,I_t^j)	1.00	1.00	1.00	1.00
corr(TB_t^i,y_t^i)	0.68	0.69	0.69	0.69

NOTE:

$RHO_\theta=\begin{bmatrix} 0 & 0 \\ 0 & 0 \end{bmatrix}$; $V_\theta=\begin{bmatrix} 1 & 0 \\ 0 & 1 \end{bmatrix}$; $(\partial\beta/\partial u)*(u/\beta)=-0.1$; $\beta(u^i)=0.99$; $\eta=0.35$; $\gamma=0.0$; d=0.025.

Standard deviations are relative to standard deviation of output (y).
Corr: correlation.

TABLE 23.-- SIMULATION RESULTS: FIXED LABOR SUPPLIES, MULTIPLICATIVE TECHNOLOGY SHOCKS. VARIATIONS IN THE STEADY STATE RATE OF TIME PREFERENCE

	β=0.93	β=0.95	β=0.97
std(c_t^i)	0.41	0.37	0.34
std($\bar{c}_t^i$)	0.29	0.27	0.25
std(I_t^i)	6.65	5.30	3.91
corr(c_t^i,c_t^j)	0.05	0.09	0.13
corr(V_t^i,V_t^j)	-0.69	-0.68	-0.68
corr(y_t^i,y_t^j)	0.01	0.00	0.00
corr(I_t^i,I_t^j)	1.00	1.00	1.00
corr(TB_t^i,y_t^i)	0.65	0.66	0.67

NOTE:

$RHO_\theta = \begin{bmatrix} 0 & 0 \\ 0 & 0 \end{bmatrix}$; $V_\theta = \begin{bmatrix} 1 & 0 \\ 0 & 1 \end{bmatrix}$; $(\partial\beta/\partial u)*(u/\beta)=-0.10$; $\sigma=2$; $\eta=0.35$; d=0.025. $\gamma=0.0$;

Standard deviations are relative to standard dev. of y.
Corr: correlation.

TABLE 24.-- SIMULATION RESULTS: FIXED LABOR SUPPLIES, MULTIPLICATIVE TECHNOLOGY SHOCKS. VARIATIONS IN SERIAL CORRELATION OF SHOCKS

	p_θ=0.25	p_θ=0.50	p_θ=0.90
$std(c^i_t)$	0.45	0.55	0.94
$std(\bar{c}^i_t)$	0.34	0.41	0.60
$std(I^i_t)$	13.03	19.72	13.40
$corr(c^i_t,c^j_t)$	0.16	0.11	-0.17
$corr(V^i_t,V^j_t)$	-0.55	-0.23	0.83
$corr(y^i_t,y^j_t)$	-0.02	-0.11	-0.24
$corr(I^i_t,I^j_t)$	-0.93	-0.97	-0.97
$corr(TB^i_t,y^i_t)$	-0.26	-0.14	-0.04

NOTE:

$RHO_\theta = \begin{bmatrix} p_\theta & 0 \\ 0 & p_\theta \end{bmatrix}$; $V_\theta = \begin{bmatrix} 1 & 0 \\ 0 & 1 \end{bmatrix}$; $(\partial\beta/\partial u)*(u/\beta)=-0.10$; $\beta(u)=.99$; $\sigma=2$; $\eta=0.35$. $\gamma=0.0$; $d=0.025$.

Standard deviations are relative to standard deviation of output (y).
Corr: correlation.

TABLE 25.-- SIMULATION RESULTS: FIXED LABOR SUPPLIES, ADDITIVE SHOCKS. VARIATIONS IN THE INTERTEMPORAL ELASTICITY OF SUBSTITUTION

	$\sigma=2$	$\sigma=3$	$\sigma=5$	$\sigma=7$
$\text{std}(c_t^i)$	0.30	0.22	0.19	0.19
$\text{std}(\bar{c}_t^i)$	0.29	0.17	0.14	0.13
$\text{std}(I_t^i)$	2.42	2.67	2.73	2.74
$\text{corr}(c_t^i,c_t^j)$	0.21	0.20	0.03	-0.07
$\text{corr}(V_t^i,V_t^j)$	-0.66	-0.72	-0.77	-0.80
$\text{corr}(y_t^i,y_t^j)$	0.51	0.60	0.01	0.02
$\text{corr}(q_t^i,q_t^j)$	1.00	1.00	1.00	1.00
$\text{corr}(I_t^i,I_t^j)$	1.00	1.00	1.00	1.00
$\text{corr}(TB_t^i,y_t^i)$	0.68	0.69	0.69	0.69

NOTE:

$RHO_g = \begin{bmatrix} 0 & 0 \\ 0 & 0 \end{bmatrix}$; $V_g = \begin{bmatrix} 1 & 0 \\ 0 & 1 \end{bmatrix}$; $(\partial\beta/\partial u)*(u/\beta) = -0.10$; $\beta(u) = 0.99$; $\gamma = 0.0$; $\eta = 0.35$; $d = 0.025$.

Standard deviations are relative to standard dev. of y.
Corr: Correlation.

TABLE 26.-- SIMULATION RESULTS: FIXED LABOR SUPPLIES, ADDITIVE SHOCKS. VARIATIONS IN THE STEADY STATE RATE OF TIME PREFERENCE

	β=0.93	β=0.95	β=0.97
std(c_t^i)	0.41	0.37	0.34
std($\bar{c}_t^i$)	0.29	0.27	0.28
std(I_t^i)	6.65	5.30	3.91
corr(c_t^i,c_t^j)	0.05	0.09	0.13
corr(V_t^i,V_t^j)	-0.69	-0.68	-0.68
corr(y_t^i,y_t^j)	0.01	0.00	0.00
corr(q_t^i,q_t^j)	1.00	1.00	1.00
corr(I_t^i,I_t^j)	1.00	1.00	1.00
corr(TB_t^i,y_t^i)	0.65	0.66	0.67

NOTE:

$RHO_g = \begin{bmatrix} 0 & 0 \\ 0 & 0 \end{bmatrix}$; $V_g = \begin{bmatrix} 1 & 0 \\ 0 & 1 \end{bmatrix}$; $(\partial\beta/\partial u)*(u/\beta)=-0.10$; $\sigma=2$; $\eta=0.35$; d=0.025. $\gamma=0.0$;

Standard deviations are relative to standard dev. of y.
Corr: Correlation.

TABLE 27.-- SIMULATION RESULTS: FIXED LABOR SUPPLIES, ADDITIVE SHOCKS. VARIATIONS IN SERIAL CORRELATION OF SHOCKS

	p_g=0.25	p_g=0.50	p_g=0.90
std(c_t^i)	0.47	0.62	1.32
std($\bar{c}_t^i$)	0.37	0.48	0.94
std(I_t^i)	2.19	1.86	0.00
corr(c_t^i,c_t^j)	0.20	0.19	0.00
corr(V_t^i,V_t^j)	-0.66	-0.64	-0.56
corr(y_t^i,y_t^j)	0.01	0.02	0.00
corr(q_t^i,q_t^j)	1.00	1.00	1.00
corr(I_t^i,I_t^j)	1.00	1.00	1.00
corr(TB_t^i,y_t^i)	0.66	0.63	0.45

NOTE:

$RHO_g = \begin{bmatrix} p_g & 0 \\ 0 & p_g \end{bmatrix}$; $V_g = \begin{bmatrix} 1 & 0 \\ 0 & 1 \end{bmatrix}$; $(\partial\beta/\partial u)*(u/\beta)=-0.10$; $\beta(u)=0.99$; $\sigma=2$; $\gamma=0.0$; $\eta=0.35$; d=0.025.

Standard deviations are relative to standard dev. of y.
Corr: Correlation.

TABLE 28.-- SIMULATION RESULTS: VARIABLE LABOR SUPPLIES

(i) Cross-country Correlation of Consumption

	els=0.1	els=0.5	els=1	els=2
σ=1.1	.30 [.99]	.36 [.97]	.40 [.92]	.43 [.82]
σ=2	.27 [.98]	.24 [.81]	.18 [.52]	.02 [.07]
σ=3	.23 [.98]	.19 [.73]	.10 [.35]	-.09 [-.13]
σ=5	.21 [.98]	.16 [.67]	.05 [.23]	-.17 [-.27]
σ=7	.20 [.97]	.14 [.64]	.02 [.18]	-.20 [-.33]

(ii) Correlation of Consumption and Hours Within the Same Country

	els=0.1	els=0.5	els=1	els=2
σ=1.1	.11 [.18]	.14 [.22]	.17 [.26]	.23 [.32]
σ=2	.36 [.26]	.46 [.47]	.55 [.62]	.67 [.77]
σ=3	.42 [.29]	.53 [.54]	.63 [.71]	.75 [.85]
σ=5	.46 [.31]	.58 [.58]	.69 [.76]	.80 [.89]
σ=7	.47 [.32]	.60 [.60]	.71 [.78]	.82 [.90]

NOTE: $RHO_\theta = \begin{bmatrix} 0.9 & 0 \\ 0 & 0.9 \end{bmatrix}$; $V_\theta = \begin{bmatrix} 1 & 0.2 \\ 0.2 & 1 \end{bmatrix}$;

$\beta(u)=0.99$; $(\partial\beta/\partial u)*(u/\beta)=-0.1$; $\gamma=0.0$; $\eta=0.35$; $d=0.025$.

$els=\sigma/(\sigma*\nu_2+(\sigma-1)*\nu_1)$ (elasticity of labor supply).

Numbers not in brackets: correlations in economy with incomplete asset markets.

Numbers in brackets: correlations in economy with complete asset markets.

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