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Current account determination in the intertemporal framework: an empirical analysis

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Current account determination in the intertemporal framework:

An empirical analysis

by

Ferdous Hossain

**A Dissertation Submitted to the
Graduate Faculty in Partial Fulfillment of the
Requirements for the Degree of
DOCTOR OF PHILOSOPHY**

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For the Graduate College

**Iowa State University
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CHAPTER 1. INTRODUCTION

Over the years an extensive empirical literature has grown that attempts to model the behavior of the current account balance of a country. In recent years such modeling efforts have shifted the focus away from the static Keynesian saving and investment functions, towards analyses based on intertemporal optimization. The current account of a country is, by national income accounting relationship, the difference between saving and investment of that country. Therefore it follows that the factors that are relevant for saving and investment decisions have direct bearing on the current account balance of the country. However, saving and investment decisions are inherently forward-looking and dynamic in the sense that they involve intertemporal choice: present as well as future considerations enter the decision making process. Static analysis in general, and Keynesian static models in particular, are not suitable to accommodate the intertemporal nature of the decision-making process that determines saving and investment, and hence current account balance. Recently many authors have tried to capture the intertemporal aspect of the issue by casting the problem of current account determination in an intertemporal framework. In this framework of analysis, response of an economic agent to exogenous disturbances in his/her economic environment depends on the entire time path of the disturbance sequence.

Real income and relative price (terms of trade) changes have always occupied the central position in the analysis of current account. Unlike the traditional static models, the

intertemporal analysis distinguishes between the effects of permanent and transitory changes in income and relative price. For example, Sachs (1981) demonstrates how a permanent real income change may have a different effect on current account compared with that of a transitory change in income. Incorporating a version of the Life Cycle Permanent Income Theory of Consumption, he shows that a permanent change in real income leaves the current account balance unaffected as both consumption and income change by similar magnitude. On the other hand, a transitory increase in income improves the current account balance as consumption increases by less than income. However, if the income increase is induced solely by demand expansion then such an income change tends to deteriorate the current account balance.

The effects of changes in relative price on the current account balance of a country (assumed to be a small open economy) have been the subject of some controversy since the 1950s when Harberger (1950) and Laursen and Metzler (1950) developed what has come to be known in the literature as the Laursen-Metzler effect. In essence they argue that a deterioration of the terms of trade (relative price across countries) of a country lowers its real income, and therefore, lowers savings. Given investment, the fall in savings leads to a deterioration of its current account balance. The effect of terms of trade changes on current account has been re-evaluated within the intertemporal framework of analysis. The re-appraisal was initiated by Obstfeld (1982) and has been further pursued within a variety of related frameworks by other authors. All these analyses carry one essential message: transitory and permanent, as well as anticipated and unanticipated,

disturbances in income and terms of trade have different effects on the current account balance of a country.

Another new approach to modeling the current account balance is to follow the methodology widely used in the real business cycle literature. This approach involves building an artificial economy by specifying tastes and technology (i.e., utility and production functions), and making assumptions about the stochastic shocks affecting the artificial economy. The equations describing the artificial economy are then solved and the stochastic properties of the solution functions are examined to see if the solutions do match real world behavior of the relevant variables. The authors of these models show that it is important to isolate the sources of the disturbance to properly understand the observed movements in macroeconomic data. They emphasize the source over timing and duration of the disturbances in their efforts to explain the observed movements in US current account balance. In empirical implementation of these models, calibration method is used instead of econometric techniques. So far, these models have produced results that match observed data only partially.

Our approach in this study is in the spirit of the intertemporal models. Despite their elegance and theoretical appeal, empirical application of the intertemporal models has been limited by our inability to separate permanent and transitory components from observed movements in income and terms of trade. Empirical models have either relied on anecdotal evidence or used traditional models. In traditional models, an increase in income tends to reduce current account surplus while depreciation of real exchange

rate - a macroeconomic analog of terms of trade deterioration - tends to improve the current account surplus. However, no distinction is made between transitory and permanent changes in income and real exchange rate. Recent developments in time series techniques provide us with ways to identify transitory and permanent components in a nonstationary variable, and thus present us the opportunity to empirically test the propositions of the intertemporal models using the decomposed data series. Specifically, two different methods have been used to obtain permanent and transitory components in real income and real exchange rate. One is the univariate method suggested by Beveridge and Nelson (1981), and the other is the bivariate method proposed and implemented (in the context of decomposing US GNP into permanent and transitory components) by Blanchard and Quah (1989). The transitory and permanent components of the data series obtained via these two methods are used in the empirical analysis.

The objective of this study is to examine whether the testable propositions generated by the intertemporal models are supported by observed data. More specifically, we shall examine whether the behavior of the current account balances of the United States and Japan are consistent with the intertemporal models, given actual data on income, real exchange rate, and government consumption. The study is organized as follows: chapter 2 presents a brief review of the literature concerning the determination of the current account balance of a country. In chapter 3, a simple conceptual framework is presented that highlights the potentially different effects of transitory and permanent movements in income and real exchange rate on the current account balance of a country.

This conceptual model provides the motivation for the empirical study that follows. Chapter 4 discusses the significance of the unit roots issue in macroeconomic data. There we discuss the methodology to test for the presence of unit roots in the data series. Later in this chapter the empirical results of the tests for unit roots in the data are presented. In chapter 5 we discuss the methodology used to obtain the transitory and permanent components of income and real exchange rate. Chapter 6 discusses the concept of cointegration, the methodology to test for the presence of potential cointegrating relations, and presents empirical results of such tests. The results of cointegration test and estimation essentially relate to long-run relationships among the variables in the model. Thus, the cointegration analysis will allow us to examine if there is any long-run relationship among current account balance, real income, and real exchange rate. However, any particular long-run relationship among a set of variables may be consistent with many different short-run dynamic behavior of the variables of the model. The cointegration analysis is incapable of capturing the short-run dynamic interaction among the variables. The short-run dynamics among the variables are the subject of discussion in chapter 7. Finally, we summarize our findings in the concluding chapter (chapter 8).

CHAPTER 2. SURVEY OF LITERATURE

Theoretical Analysis

During the 1970s, world economy as a whole saw major shifts in the behavior of international capital flows, current account balances, and exchange rate (both nominal and real) movements. In particular, several developed and developing countries experienced what are by historical standards large disturbances in their terms of trade and current account balances. These changes may be attributed to several factors such as sudden and large changes in the relative prices of some intermediate inputs, specially oil, large changes in the relative prices of final manufactured goods and final primary commodities. There was no consensus among economists regarding the causes of such current account imbalances. Different analyst had different explanation for the current account imbalances of countries¹. However, the current account difficulties of the 1970s generated a renewed interest, during the 1980s, among economists about the impacts of various disturbances to income and terms of trade on macroeconomic variables such as spending, saving, investment, and in particular the current account balance.

Although the revival of interest during the 1980s in the current account balance of a country was due to the recent experiences of many countries, theoretical discussions on the subject were prompted by the general dissatisfaction among economists with the traditional approaches to the issue. Specifically, it was felt that the traditional static framework was not appropriate for analyzing the current account determination.

Income and relative price have always occupied a central position in any discussion of current account balance of a country. Traditionally, theoretical as well as empirical analyses of current account were performed within the framework of the Keynesian model of income determination in an open economy. In these models, trade or current account balance of a country is assumed to be a function of domestic income, foreign income, and relative price. Both domestic and foreign general price levels are assumed to be fixed. Relative price moves in response to devaluation of nominal exchange rate. Income is assumed to be demand determined. In such models, an increase in domestic income brought about by an increase in domestic aggregate demand (for example, an increase in government consumption) leads to a trade/current account deficit as part of the increased expenditure is spent on imports. Even if the effects of increased demand for imports (in the home country) on foreign income is taken into account, the qualitative results remain unchanged. Until the beginning of the 1980s, this has been the characteristic of the discussions of the effects of income changes on trade/current account balance.

Since the publication of Harberger (1950), and Laursen and Metzler (1950), the effects of relative price changes on the current account balance of a country have been discussed quite extensively. The general wisdom has been that if the Marshall-Lerner condition is satisfied then, starting from an initial position of balanced trade, a devaluation (and the associated increase in the relative price of imports) leads to an improvement in the trade balance. If unilateral transfers are ignored, then that implies an

improvement in the current account balance in response to a terms of trade deterioration. However, both Harberger (1950), and Laursen and Metzler (1950) focuses on the response of saving to real income decline resulting from a deterioration in the terms of trade. In accord with the Keynesian theory of consumption function, Laursen and Metzler (1950) observes, on the basis of statistical evidence, that (in the *short run* of a cycle) a rising proportion of income is saved. Therefore, they argue, a deterioration of the terms of trade leads to a reduction in real income which causes an increase in expenditure relative to income. Given investment, that implies a current account deficit. Harberger (1950) also focuses on the effects of terms of trade deterioration on domestic saving, measured in domestic exportable good. He assumes that saving, measured in exportable good, is a functions of real income only. Changes in prices do not induce any substitution between saving and consumption. Further, he assumes that the marginal propensity to save for changes in real income due to changes in terms of trade is the same as that for changes in output at constant prices, and that the effects of a change in the terms of trade on real income may be approximated by the change in the cost of the initial value of imports. With these assumptions, Harberger (1950) argues that a deterioration in the terms of trade brings about a decline in domestic saving as a result of a decline in the real income. Given investment, this then leads to a current account deficit. This apparent counter intuitive conclusion is what has come down in the literature as the Laursen-Metzler effect². Harberger (1950) is criticized by Day (1954) on the ground that savings and

imports may be substitutes. Laursen and Metzler (1950) is criticized by White (1954) who argues that time lags (in the adjustment process) make the long-run behavior of saving as ratio of income more relevant, and that the long-run saving ratio is constant³.

Now, in order to get some insight into the determination of current account, let us start with the national income identity:

$$Y \equiv C + I + G + (X - M),$$

where Y is income, C is aggregate consumption, I is investment, G is government consumption, and $(X - M)$ is net exports. Subtracting net taxes T (taxes less domestic transfers) from both sides, and adding net international transfer receipts R to both sides of the above identity, we get:

$$Y + R - T \equiv C + I + (G - T) + (X + R - M).$$

Rearranging, we can write:

$$(X + R - M) \equiv (Y + R - C) - I + (T - G).$$

The term $(X + R - M)$ is the current account balance of the country, $(Y + R - C)$ is the domestic saving, and $(T - G)$ is the government budget surplus. Hence we can express the current account balance of a country as:

$$CA \equiv S - I + (T - G),$$

which says that the current account balance is equal to the difference between national saving and investment. Therefore, saving and investment are integral parts of any analysis of current account determination.

During the 1970s, the traditional Keynesian theories of consumption and investment, among others, came under serious scrutiny and re-examination. In particular, it was felt that a static Keynesian framework is inappropriate for analyzing consumption saving decision of individual economic agents. Economists generally agreed that consumption-saving, and investment are inherently forward looking and dynamic variables that should to be analyzed in a dynamic optimizing setting. Since current account balance of country is, by national income identity, the difference between saving and investment of a nation, it was felt that current account also should be analyzed in a dynamic optimizing framework.

As has been mentioned, the revival of interest in the current account determination was due at least in part to the 1973 oil price shock. The conventional wisdom at that time was that, given limited substitution possibilities in production, an increase in the price of an intermediate input would lead to current account deficit (Findlay and Rodríguez 1977; Buiters 1978; and Bruno and Sachs 1979). Obstfeld (1980), and Sachs (1981) challenge such conventional wisdom. Their analytical models generate results that are contrary to those of the traditional models. Their models predict that a permanent oil price increase (a permanent terms of trade deterioration for an oil importing country) may actually improve the current account of an oil importing country. Obstfeld's (1980) prediction is based on the assumption of low price elasticity of demand for oil. In his model he demonstrates that it is only when there is a significant degree of substitution between imported intermediate goods and domestically available factors of production,

and between finished imports and exports is it possible to have a current account deficit in response to an increase in the imported raw material price (oil price). Sachs (1981) emphasizes the depressing effect of a permanent oil price increase on investment to predict a current account improvement in response to such a price shock. In the context of the observed large current account deficits of the oil importing countries, he argues that higher oil price alone can not explain the cross-country differences in current account deficit among the oil importing countries. Rather, he goes on to argue, permanently higher oil price lowered world interest rate, and induced investment boom (in the US) which led to the current account deficit (of the US) as national savings fell short of booming investment. Marion (1984), however, shows that under a more realistic model structure that includes nontradable goods (e.g., services) the strong conclusions of Obstfeld (1980) and Sachs (1981) do not necessarily hold. With the introduction of nontraded goods, both temporary and permanent oil price increases may have ambiguous effects on the current account balance. These conflicting results about the effects of oil price shocks on current account provided new impetus for a theoretical re-examination of the issue of current account determination.

The renewed discussions of current account determination during the 1980s is characterized by a shift in the focus away from static Keynesian savings and investment functions, toward analyses based on intertemporal optimization. These discussions typically distinguish between temporary and permanent, as well as anticipated and unanticipated, changes in income and terms of trade in analyzing the current account

balance. Sachs (1981, 1982) is perhaps one of the first to discuss how the current account is affected by temporary and permanent disturbances in income. Assuming a periodic utility function that is additively separable, a constant rate of time preference, and a given world interest rate, Sachs (1982) shows that a temporary increase in income leads to current account improvement unless the rate of time preference is much higher than the given world interest rate. In that case, consumption smoothing response to the income shock dominates the substitution effect. If the rate of time preference greatly exceeds the world rate of interest, the wealth effect may dominate the consumption smoothing effect such that a temporary increase in income may lead to current account deterioration. A permanent income increase that raises wealth improves current account if and only if the rate of interest exceeds the constant rate of time preference. An anticipated income increase that raises (lifetime) wealth typically worsens current account, and necessarily does so if the interest rate is lower than the rate of time preference. In the benchmark case where the rate of time preference is equal to the world interest rate, a temporary income increase improves current account while a permanent income increase leaves the current account largely unaffected. Obstfeld (1983) confirms the conclusion arrived by Sachs (1982), and adds that such responses of current account to output disturbances are independent of the risk aversion by the representative consumer.

Although disturbances to both income and terms of trade are important to current account movements, most of the recent discussions on the issue are on the effects of terms of trade shocks on current account balance of a country. Since the publication of

Laursen and Metzler (1950), and Harberger (1950), the issue of the effects of a terms of trade shock on the current account balance of a small open economy has been revisited many times by different authors. Recent reexamination of the issue was initiated by Obstfeld (1982), and has been pursued within a variety of frameworks by a number of authors. Obstfeld (1982) challenges the prediction of Laursen-Metzler analysis by arguing that a permanent deterioration of the terms of trade may result in increased saving and an improvement in the current account balance of the country experiencing the terms of trade deterioration. Obstfeld (1982) uses a model of a small open economy consisting of an infinitely lived representative consumer with an Uzawa (1968)-type utility function where the rate of time preference is assumed to be an increasing function of utility. This consumer can lend or borrow freely at a given world rate of interest. Such an economy has a target level of real wealth, at the point where the rate of time preference is equal to the (given) world rate of interest. When such an economy faces a terms of trade deterioration, its real wealth level is lowered⁴. In order to reach the target level of wealth, the country must accumulate foreign assets, and hence save more. Therefore, in contrast to Laursen-Metzler effect, the permanent deterioration of the terms of trade leads to current account surplus.

It should be mentioned here that Obstfeld (1982) considers a pure endowment economy where there is no production and investment. Further, his assumption about the rate of time preference being a positive function of utility is critical in obtaining the results⁵. Also, he considers only unanticipated and permanent terms of trade shocks.

No consideration is given to the potential effects of temporary or (anticipated) future shocks to terms of trade on current account. However, in another discussion, Obstfeld (1983) considers the effects of temporary, (anticipated) future, and permanent movements in the terms of trade and their effects on current account. There he shows that a temporary deterioration of the terms of trade leads to a current account deficit for the duration of the adverse movement in the terms of trade. When the terms of trade returns to its pre-shock level, current account moves into surplus and eventually returns to its long-run equilibrium path.

Obstfeld (1982) is followed by a number of analyses that examine the issue under a variety of related frameworks. Svensson and Razin (1983) examine the issue raised by Obstfeld (1982) under a more general setting. First, they use a two-period intertemporal optimizing model to examine the effects of different types of terms of trade shocks on domestic spending and current account of a small open economy. They show that the effects of a terms of trade deterioration on spending and current account can be separated into three effects:

1. a direct effect, consisting of a revaluation of the consumption and net exports.
2. a wealth effect on spending or, more precisely, a welfare effect, because the change in terms of trade changes welfare level of the representative agent, which corresponds to an equivalent change in wealth at constant prices.
3. substitution effects, due to changes in spending pattern because of changes in

relative prices within and between periods, at constant welfare level.

Svensson and Razin (1983) find that without some additional structures are imposed on preferences, the direct and wealth effects can in some cases be signed unambiguously whereas the substitution effect is ambiguous. In particular, if the terms of trade deterioration is temporary then the sum of the direct and wealth effects on current account is negative. For a future terms of trade deterioration, in Svensson and Razin (1983) model, the direct effect is zero and the wealth effect is positive. So, if the substitution effect is ignored, current account deteriorates in response to a temporary terms of trade deterioration in the first period, and improves for an (anticipated) future terms of trade deterioration. If the terms of trade deterioration is permanent, the direct and wealth effects on current account balance are in general ambiguous.

Imposing the additional assumption that the utility function of the representative consumer is weakly separable in time, where the subutility functions in each period are linearly homogeneous, they find the following results⁶:

- (i) Holding constant the real discount factor, a temporary deterioration in the terms of trade has the unambiguous effect of deteriorating the current account. This effect is nothing but the sum of direct and wealth effects.
- (ii) A temporary terms of trade deterioration also affects the real discount factor (used to convert the second period values to present value in the first period). If such deterioration leads to an increase in the (real) discount factor, then the substitution effect of a temporary terms of trade deterioration works counter to the

direct plus wealth effects, and the overall effect of such a terms of trade deterioration on current account is ambiguous.

(iii) An (anticipated) future terms of trade deterioration, holding real discount factor constant, leads to an unambiguous improvement in current account balance. However, such a change in terms of trade will also affect the real discount factor and therefore will have substitution effect. So, the overall effect will depend on whether the substitution effect reinforce direct plus wealth effects, or work counter to each other.

(iv) Making an additional assumption that preferences are not only homothetically weakly separable, but also identically so, Svensson and Razin (1983) derive more clear results. Under these assumptions, the real current account balance deteriorates, or improves, in response to a permanent terms of trade deterioration depending on whether the marginal propensity to spend in the second period exceeds or falls short of the marginal propensity to spend in the first period.

Under the assumptions made in Svensson and Razin (1983), the marginal propensity to spend and rate of time preference are intimately related. In particular, the rate of time preference for a given level of welfare (utility) is equal to the (absolute value of) marginal rate of substitution between real spending in the second period and real spending in the first period minus one, evaluated at the point where real spending in the two periods are equal. Put differently, the rate of time preference is the subjective interest rate along the 45-degree ray from the origin. A necessary condition for the marginal

propensity to spend in the second period to exceed that in the first period is that the rate of time preference decreases with the level of welfare.

Svensson and Razin (1983) also extends their discussion from a two-period model to an infinite horizon model. In the infinite horizon case, they assume that the real discount factor between any two consecutive periods is constant. The necessary and sufficient condition for stability of this infinite horizon model is that the rate of time preference increases with the level of welfare. This is precisely the assumption made by Obstfeld (1982). If stability is assumed (i.e., the behavior of rate of time preference is identical in Obstfeld (1982) and Svensson and Razin (1983)), then the results of Obstfeld (1982) are confirmed by Svensson and Razin (1983).

Thus, Svensson and Razin (1983) indeed provide a more generalized treatment of the issue discussed by Obstfeld (1982), Sachs (1981, 1982). However, the two-period model of Svensson and Razin (1983) can only accommodate limited dynamics of the variables. On the other hand, the infinite horizon models of Obstfeld (1982, 1983) and Svensson and Razin (1983) require, for stability, the condition that the rate of time preference is increasing in utility. Although such a property of the rate of time preference is not inconsistent with basic axioms of neoclassical utility maximization models⁷, such a behavior of the rate of time preference seems counter-intuitive. Also, as Persson and Svensson (1985) argues, the assumption of infinitely lived agents gives rise to a very high degree of consumption smoothing and intertemporal substitution in the models of Sachs (1982), Obstfeld (1982, 1983), and Svensson and Razin (1983). Persson and Svensson

(1985), therefore, suggest that a finite planning horizon may be a more realistic approach to modeling the problem under consideration.

Persson and Svensson (1985) develop a finite horizon overlapping generations model (without private intergenerational gifts). Using a simple production technology and preferences that are similar to those in Svensson and Razin (1983), they show that it is important to distinguish not only between permanent and temporary shocks to terms of trade but also between anticipated and unanticipated shocks. They argue that any attempt to derive unqualified statements about the dynamic adjustment of current account to terms of trade changes is, what they call, *a futile exercise*. Another phenomenon of the dynamic adjustment of current account to terms of trade shocks in their model is that, unlike Obstfeld (1982, 1983) and Svensson and Razin (1983), current account shows oscillating behavior rather than smooth and steady convergence to stationary state value. The forces working behind such behavior of the current account balance are the indirect effects of terms of trade shocks on real interest rate and investment.

Ostry (1988) and Edwards (1989) independently extend Svensson and Razin (1983) by including a nontraded good in the model. Ostry (1988) provides a more comprehensive discussion on the relationship between terms of trade and current account than Edwards (1989). Therefore here only the analysis of Ostry (1988) is presented briefly. The framework used by Ostry (1988) is very similar to Svensson and Razin (1983) except that a nontraded good is included. Using a model of a small open economy, he demonstrates that the incorporation of a nontraded good in the model has important

effects on the model predictions. Like other analyses mentioned earlier, Ostry (1988) examines the effects of temporary and permanent shocks to terms of trade on current account balance. He finds that temporary terms of trade disturbances will in general have different effect on current account balance depending on whether or not the model includes a nontradable good. The effect of a temporary shock to the terms of trade on current account critically depends on three factors: (i) initial debt position of the country; (ii) the relative magnitudes of temporal and intertemporal elasticities of substitution; and (iii) the relative magnitudes of the intertemporal elasticity of substitution and the ratio of imports to consumption of importables. The response of current account to a permanent terms of trade shock will be qualitatively similar in models with and without nontradable goods if the elasticity of substitution between domestic nontradable and import goods exceeds the ratio of imports to consumption, assuming that initially the current account is balanced.

Thus far the discussions about the effects of temporary and permanent shocks to terms of trade on current account do not include investment⁸, and assume fixed labor employment in the economy. Sen and Turnovsky (1989) does fill this gap by including investment and by making employment endogenous through labor-leisure choice⁹. They use an infinite horizon utility maximizing model with two goods: one domestically produced and the other imported. They allow capital accumulation and derive a q-theoretic investment function¹⁰. Effects of three types of shocks to terms of trade on current account are considered, namely an anticipated permanent, an unanticipated

temporary, and a future anticipated permanent shock. The essential message from their discussion is that the response of current account to all three types of shocks to terms of trade depend critically upon the long-run effects of terms of trade shocks on the stock of capital. This in turn consists of a negative substitution effect and a positive income effect. The final outcome depends on which effect dominates. Although neither effect can be ruled out *a priori*, Sen and Turnovsky (1989) suggest that the more plausible case is where the substitution effect dominates. In that case, irrespective of whether the terms of trade shock is temporary or permanent, anticipated or unanticipated, current account improves in response to terms of trade deterioration when the news of the shock arrives. However, the long-run effect on current account of a (unanticipated) permanent terms of trade shock is greater than that of a temporary or an anticipated future terms of trade shock. This is contrary to Laursen-Metzler effect and this happens through a fall in investment. In the other case, i.e., when the income effect dominates, the news of terms of trade deterioration generates an investment boom and drives the current account into deficit. In particular, a permanent terms of trade deterioration (either anticipated or unanticipated) leads to a steady accumulation of capital accompanied by current account deficit. A temporary shock leads to a permanent increase in capital stock and a permanent reduction of the stock of foreign bond holdings. While this vindicates Harberger (1950) and Laursen-Metzler (1950) predictions, the source of such a movement in current account is the adjustment in investment rather than savings as argued in Harberger (1950)

and Laursen-Metzler (1950) analyses.

It should be mentioned that the results of Sen and Turnovsky (1989) depend on certain assumptions made in their analysis. First, both import and export goods, and leisure are assumed to be complementary, as are the two goods themselves. Second, labor supply is endogenous. Finally, the country initially has some positive holding of foreign asset. The results of their analysis would be significantly altered without these assumptions.

The above mentioned analyses are the major contributions in the area of current account determination in the intertemporal framework that distinguish among various types of income and terms of trade shocks. Other important contributions are Bean (1986), Matsuyama (1987), Sen (1990), Brock (1988), Ostry and Reinhart (1992). The essential message of all these discussions is that the current account of a country ought to be analyzed in a dynamic framework, and that distinctions should be made among different types of shocks in real income and relative prices.

Recently some attempts have been made to model the relationship among trade balance, income, and real exchange rate within a dynamic general-equilibrium framework. These efforts extend the stochastic general-equilibrium models used by the real business cycle school of thought by incorporating external trade. The main contributors in this class of models are Backus, Kehoe, and Kydland (1992, 1994), Mendoza (1991), Stockman and Tesar (1991), and Tesar (1992). The theoretical structures used in these analyses are extensions of Stockman and Svensson (1987), and

Hodrick (1989), who develop simple general-equilibrium models in which both terms of trade and trade balance are endogenously determined. All these models attempt to construct artificial economies and calculate the general-equilibrium solutions of the endogenous variables by assigning *reasonable values* to the parameters of the model constructed. Then they examine the responses of important macroeconomic variables of their model to various shocks, and compare the behavior of the response functions with real world data. The ultimate objective of such modeling effort is to generate the general equilibrium solutions for the endogenous variables such that the behavior of the variables in the artificial economy resemble as closely as possible their empirical counterparts in real world.

Backus, Kehoe, and Kydland (1992, 1994) use a two-country, two-good framework. Each country specializes in the production of one good. Both goods are consumed in each country, and these two goods are imperfect substitutes. Both goods are used for consumption as well as for capital formation. Technology of production incorporates the *time-to-build* feature of Kydland and Prescott (1982). Each country has one infinitely lived representative consumer who derives utility from consumption of goods as well as leisure. In these models fluctuations arise from persistent shocks to aggregate productivity and government consumption of goods. The behavior of income, trade balance, and terms of trade are examined by solving the model with *reasonable values* of the parameters. They find that, in their model, the dynamic responses of the endogenous variables to aggregate productivity shocks are markedly different from the

responses of the endogenous variables to government consumption shocks. Specifically, a favorable domestic productivity shock leads to an increase in domestic output, a decrease in the relative price of the domestic good (a terms of trade deterioration), and a deterioration in the trade balance. The counter-cyclical behavior of the trade balance, and the negative correlation between terms of trade and trade balance are due to the dynamic responses of consumption and investment to a (positive) domestic productivity shock. Initial positive productivity shock generates higher income and higher consumption, and creates a temporary investment boom as capital moves to the more productive sector. Increases in consumption and investment together exceed the gain in output, and the domestic trade balance deteriorates. However, when the experiment is repeated under the assumption of no capital and investment, the (positive) productivity shock (in the home country) leads to an improvement in trade balance, and thus generates a pro-cyclical movement in trade balance. Hence, investment dynamics play a central role in explaining the observed co-movements in income, terms of trade, and trade balance. On the other hand, persistent government consumption shock generates positive instantaneous correlation between terms of trade and trade balance: trade balance improves as terms of trade deteriorates. Hence, Backus, Kehoe, and Kydland (1992, 1994) argue that it is important to identify the source of the disturbance in understanding the observed movements in trade balance, income, and terms of trade. In other words, they suggest that there is no structural relationship among income, terms of trade, and trade balance. One

can not characterize the relationship among these variables without specifying the source of the fluctuation.

Stockman and Tesar (1991) use a very similar model with the exception that they incorporate a nontraded good in their model. They allow productivity as well as preference shocks to examine the co-movements in income, terms of trade, and trade balance. Their results are qualitatively similar to those of Backus, Kehoe, and Kydland (1992, 1994).

It should be mentioned that all these models are calibrated rather than econometrically estimated from real world data. Results of this type of models are often very sensitive to the choice of the parameter values used to solve for the endogenous variables. Specifically, the solutions for the endogenous variables in both Backus, Kehoe, and Kydland (1992, 1994), and Stockman and Tesar (1991) are sensitive to the value assigned to the elasticity of substitution between domestic good and the import good. Both Backus, Kehoe, and Kydland (1992, 1994), and Stockman and Tesar (1991) find that the co-movements of income, terms of trade, and trade balance from their artificial economies are consistent with observed data when there is an aggregate productivity shock. However, whether the countries studied in these papers actually experienced such productivity shock over the period of study is an empirical issue. In other words, the consistency of the model predictions and observed data is of any value only if actual economies experienced the same shock that generates the data series in the models of the artificial economies. Finally, as the authors of these studies concede, there are many

important conflicts between the predictions of these models and the data from the countries that these models are supposed to mimic. The behavior of some data series generated by the models and their co-movements are not always consistent with real world data. Therefore, this new approach to the study of the relationship among income, terms of trade, and trade balance is an important avenue to explore, but additional work needs to be done before one can make a more definitive judgment about the usefulness of this line of research.

Empirical Analysis

Empirical works on the issue of external balance more or less exclusively use trade balance as the dependent variable. Essentially almost all of the studies approach the issue from the standpoint of trade theory rather than from the perspective of open economy macroeconomics. In addition, almost all of the studies use the more traditional models in estimating the effects of changes in income and real exchange rate on the balance of trade of a country. The usual approach is to specify an import demand function where demand for the imported good is a positive function of domestic real income, a negative function of foreign income, and a positive function of (relative) price of the import good. In addition, an export supply function is specified where the quantity of export is assumed to be a positive function of the relative price of the foreign good. Domestic import and export goods are assumed to be imperfect substitutes for each other. Trade balance (or net exports) is then defined as the difference between domestic exports and imports. A classic example of such model specification can be found in Goldstein

and Khan (1985). Goldstein and Khan (1985) reviews the empirical studies that model trade balance in this framework¹¹. Their imperfect substitutes model is specified as follows:

$$I_i^d = f(Y_i, PI_i, P_i), \quad f_1, f_3 > 0, f_2 < 0,$$

$$X_i^d = g(Y^*e, PX_i, P^*e), \quad g_1, g_3 > 0, g_2 < 0,$$

$$I_i^s = h[PI^*(1 + s^*), P^*], \quad h_1 > 0, h_2 < 0,$$

$$X_i^s = j[PX_i(1 + s_i), P_i], \quad j_1 > 0, j_2 < 0,$$

$$PI_i = PX^*(1 + T_i)e,$$

$$PI^* = \frac{PX_i(1 + T^*)}{e},$$

$$I_i^d = I_i^s e,$$

$$X_i^d = X_i^s,$$

where:

I_i^d = quantity of imports demanded in country i,

X_i^d = quantity of country i's exports demanded by rest of the world,

I_i^s = quantity of imports supplied to country i from rest of the world,

X_i^s = quantity of exports supplied from country i to rest of the world,

PI_i = domestic currency price paid by importer of country i,

PI^* = domestic currency price paid by importers of rest of the world,

PX_i = domestic currency price received by exporters of country i,

PX^* = domestic currency price received by exporters of rest of the world,

Y = level of nominal national income of country i,

Y^* = level of nominal national income of the of the world,

P = general price level of country I,

P^* = general price level in rest of the world,

T = proportional tariff on imports in country i,

T^* = proportional tariff on imports in rest of the world,

S = subsidy rate applied to imports in country i,

S^* = subsidy rate applied to imports by rest of the world,

e = nominal exchange rate (domestic currency price of foreign currency).

Similar specification can be found in Helkie and Hooper (1988), and Marques and Ericsson (1993). The econometric model estimated by Marques and Ericsson (1993), which is representative of the vast majority of the empirical studies on this issue, is the following:

$$\ln X_{it} = \alpha_{1i} + \alpha_{2i} \ln Y_t^* + \alpha_{3i} \ln \left(\frac{P_{xit}}{e_t P_t^*} \right) + u_{xit},$$

$$\ln M_{it} = \beta_{1i} + \beta_{2i} \ln Y_t + \beta_{3i} \ln \left(\frac{P_{mit}}{P_t} \right) + u_{mit},$$

$$\ln P_{xit} = \gamma_{1i} + \gamma_{2i} \ln C_t + \gamma_{3i} \ln(e_t P_t^*) + \dot{v}_{xit},$$

$$\ln P_{mit} = \delta_{1i} + \delta_{2i} \ln(e_t P_t^*) + v_{mit},$$

where

C = production cost,

e = nominal exchange rate (domestic currency price of foreign exchange),

P_m = import price,

P_x = export price,

X = real export,

M = real import,

P = GDP deflator,

Y = GDP, and

u_{jit} and v_{jit} are random disturbances. Asterisk denotes foreign variable.

The above specification of a model that intends to explain the trade / current account balance suffers from two fundamental problems. First, conceptually such a model is unable to capture the dynamic saving-investment behavior involved in the determination of trade or current account balance. It may be used with micro-level data to analyze the net exports of a single good, but is not suitable for capturing the underlying dynamic decision-making process that determines trade/current account balance. Also, this type of model does not distinguish between temporary and permanent movements in income and terms of trade. Second, from pure econometric perspective, we observe that some of the variables of the model are non-stationary (i.e., they contain a unit root in their autoregressive component) as has been demonstrated by numerous studies¹². It is well known that the presence of nonstationary variables in an econometric model may have serious consequences on the estimation method and on the statistical properties of the commonly used estimators (such as the OLS). Specifically, many of the standard estimators lose their usual properties if the data contain nonstationary variables. Further, simple differencing of the variables (to obtain stationarity of the nonstationary variables) may not be appropriate. Potential cointegrating relations among the nonstationary

variables should be estimated, and then incorporated in further analysis. These methodological weaknesses imply that the results of these studies may have very low reliability. However, some of the recent empirical analyses do recognize the issue of unit roots in the variables, and existence of potential cointegration among them. Rose and Yellen (1989), Rose (1991), Burada and Gerlach (1992), and Peruga (1992) are a few of such studies. All of these works focus on the relation between trade balance and exchange rate (real and / or nominal), but do not distinguish between temporary and permanent changes in income and real exchange rate. In contrast to all of these works, we focus on the current account from the perspective of the intertemporal optimizing models, where we attempt to isolate the transitory and permanent movements in income and real exchange rate, and use the decomposed series to empirically test whether observed data support the hypotheses of the intertemporal models. Methodologically, we provide a more rigorous treatment of the issues of unit roots and cointegration than some of the recent analyses. After dealing with the issue of unit roots in the variables, cointegration analysis is used to identify long-run relations among the nonstationary variables. Finally, we use the Vector Autoregression (VAR) technique along with impulse response functions and innovation accounting to investigate the short-run dynamics in the variables included in the model.

Endnotes

1. Paul Volcker (1977) held large imports of oil responsible for US current account deficits, while McKinnon (1980) emphasized low saving rate and / or high investment for the deficit.
2. In fact Harberger (1950) was published before Laursen and Metzler (1950), and it is perhaps more appropriate to identify this result as the Harberger-Laursen-Metzler effect.
3. See Harry Johnson (1967) for more discussion on this issue.
4. It is assumed that the economy has no initial debt, positive or negative. An economy that starts with net foreign asset (positive or negative) may generate results that are different from those obtained by Obstfeld (1982).
5. Obstfeld (1982) recognizes that the assumption about the rate of time preference and exclusion of all but single asset (claims on foreigners) make his analysis narrow. Different assumptions about the structure of preferences and the range of assets may yield different results, and as such the findings of this analysis should be viewed only as an example of how the Harberger-Laursen-Metzler relation may be reversed.
6. All the results are predicated upon the implicit assumption that the country has no initial net debt.
7. See Koopmans, Diamond, and Williamson (1964) on this issue.
8. Matsuyama (1988), in a different context, allows investment in discussing the effects of terms of trade shocks on current account in a Heckscher-Ohlin framework. However, that discussion focuses primarily on factor reallocation rather than current account balance.
9. Bean (1986) is an exception, but the discussion is not comprehensive.
10. For discussions on the q-theoretic investment function used in Sen and Turnovsky (1989), see Hayashi (1982), and Abel and Blanchard (1983).
11. See Goldstein and Khan (1985) for detailed discussion and additional references.
12. Starting with Nelson and Plosser (1981), numerous studies have shown that GDP, aggregate price indices are non-stationary variables. Examples are Cochrane (1988, 1994), Campbell and Mankiw (1989). Similarly, Meese and Singleton (1992), Huizinga

(1987), Corbae and Ouliaris (1988) Enders (1988), Meese and Rogoff (1988), and Mark (1990) all show that nominal and real exchange rates are also nonstationary variables. Rose (1991) shows that trade balance of the US is nonstationary. As we shall see later, the current account balance of the United States, and Japan are also nonstationary.

CHAPTER 3. MODEL AND DATA

Conceptual Model

This chapter outlines a conceptual framework that underlies the empirical analysis presented in the study. In order to keep the discussion simple, while at the same time highlighting the theoretical underpinning that motivates the empirical analysis, the effects of income and relative price on the current account balance of a country are discussed separately. The discussion is in the intertemporal dynamic optimizing framework of analysis.

Income

In order to illustrate how transitory and permanent shocks to real income affect the current account balance of a country, let us abstract initially from the effects of relative price (terms of trade) changes and assume that there is one aggregate consumption good. The infinitely lived representative consumer derives his/her utility from the consumption of this aggregate commodity. The periodic utility function representing the preference structure of the representative agent is:

$$u_t = u(C_t), \quad u' > 0, \quad u'' < 0,$$

where C_t is the consumption in period t ($t = 0, 1, 2, \dots, \infty$). The representative consumer chooses $\{C_t\}_{t=0}^{\infty}$ to maximize:

$$U = E_0 \sum_{t=0}^{\infty} \beta^t u(C_t), \quad 0 < \beta < 1, \quad (3.1)$$

where E_t is the mathematical expectation operator conditional on information available to the agent at time t and β is the subjective discount factor. The consumer faces the lifetime budget constraint given by

$$A_{t+1} = R_t[A_t + Y_t - C_t], \quad t = 0, 1, 2, \dots, \infty,$$

where $R_t = (1 + r_t)$ is the gross rate of return on saving between period t and $t+1$, A_t is the asset (or debt, if negative) holding in period t , Y_t is the real income (other than income from investment of previous savings) in period t . It is assumed that Y_t follows a stochastic process that is beyond the control of the agent. However, the agent is assumed to know the law of motion of that describes the evolution of Y_t . At time t , the consumer's information set contains at least $\{Y_{t-s}, s \geq 0; A_t\}$. Further, it is assumed that $\beta R^2 > 0$, $\lim_{j \rightarrow \infty} E_t \beta^j Y_{t+j} = 0$ ($\forall t$), and that $E_0(A_t) \geq k > -\infty$. The last restriction rules out indefinitely large and growing borrowing.

For our present exposition, we assume that $u(C_t)$ is quadratic in its arguments.

Specifically, we assume that:

$$u_t = a_0 + a_1 C_t - \frac{a_2}{2} C_t^2, \quad a_0, a_1, a_2 > 0. \quad (3.2)$$

The above utility function satisfies $u' > 0$, $u'' < 0$ only for C_t in the interval $0 \leq C_t \leq \frac{a_1}{a_2}$.

It is assumed that the satiation level of consumption, (a_1/a_2) , is large relative to the typical value of Y_t . More precisely, we assume for all $t \geq 0$, with probability arbitrarily close to unity, that we have the following condition:

$$E_t \sum_{j=0}^{\infty} R^{-j} Y_{t+j} < \frac{a_1}{a_2}. \quad (3.3)$$

This assumption is made simply to prevent consuming at the satiation level (a_1/a_2) for $t \geq 0$. For algebraic simplicity, we further assume that $R_t = R =$ (positive) constant. Given the above assumptions and restrictions, we may derive, via utility maximization, a consumption function of the following form¹:

$$C_t = \left(\frac{-\alpha}{R-1} \right) + \left(1 - \frac{1}{\beta R^2} \right) \left[\sum_{j=0}^{\infty} \left(\frac{1}{R} \right)^j E_t Y_{t+j} + A_t \right], \quad (3.4)$$

where

$$\alpha = \frac{a_1 (1 - (\beta R^2)^{-1})}{a_2}, \text{ and } \beta R^2 \text{ is assumed to be greater than unity.}$$

Further, if we assume that $\beta R = 1$, as is often done in the literature, then the consumption function may be written as:

$$C_t = \left(1 - \frac{1}{R} \right) \left[\sum_{j=0}^{\infty} \left(\frac{1}{R} \right)^j E_t Y_{t+j} + A_t \right].$$

Assuming no initial asset holding ($A_0 = 0$), the consumption function may be written as:

$$C_t = \left(1 - \frac{1}{R} \right) \left[\sum_{j=0}^{\infty} \left(\frac{1}{R} \right)^j E_t Y_{t+j} \right].$$

Denoting $\frac{1}{R}$ by δ , we may rewrite the consumption function as²:

$$C_t = (1 - \delta) \sum_{j=0}^{\infty} \delta^j E_t Y_{t+j}. \quad (3.5)$$

The stochastic process describing the evolution of real income is given by the following two equations:

$$Y_t = \tau_t + \varepsilon_{Tt},$$

$$\tau_t = \tau_{t-1} + \varepsilon_{pt},$$

where ε_{Tt} and ε_{pt} are stationary stochastic processes³. The above description of the real income process has the characteristic that Y_t is affected by two distinct types of disturbances: first is the transitory shock represented by the ε_{Tt} disturbance; second is the permanent shock represented by the ε_{pt} disturbance. Such a specification of the stochastic process for Y_t is in the accord with the general notion that real income of a country is better described as a unit root process plus a stationary component. Here τ_t represents the unit root process and shocks to τ_t are permanent in nature. The stationary component is represented by ε_{Tt} . The effects of shocks to the transitory component decay with time. We assume that there is no investment. Hence, current account of the country is the difference between real income and consumption in each time period.

In this framework, assuming no debt at the beginning of period t , the current account balance of a country in period t can be described with the following equations:

$$C_t = (1 - \delta) \sum_{j=0}^{\infty} \delta^j E_t Y_{t+j}, \quad (3.6)$$

$$Y_t = \tau_t + \varepsilon_{Tt}, \quad (3.7)$$

$$\tau_t = \tau_{t-1} + \varepsilon_{pt}, \quad (3.8)$$

$$b_t = Y_t - C_t, \quad (3.9)$$

where b_t is the current account balance in period t . In this framework, a permanent shock to income in period t will leave the current account unaffected since both consumption and income will move by the same amount. On the other hand, a transitory shock to income in period t will lead to a current account surplus in period t since consumption in period t will rise by less than the increase in income in that period⁴.

Relative Price

In this section, we shift our attention to the effects of transitory and permanent movements in the real exchange rate on the current account balance of a country. For simplicity, a discrete time two-period model is used to demonstrate how transitory and permanent movements in real exchange rate may have differential effects on the current account balance of a country. First the assumptions about the technology of production is described⁵. Then the preference structure of the consumers is outlined. Throughout the discussion competitive market structure is assumed. Using the described economic environment (technology, preferences, and market structure), we derive the functions describing the current account balance of a country. These functions allow us to examine the effects of transitory and permanent real exchange rate movements on the current account balance of a country.

Technology

Let us consider the following simple two-good production structure. The home country produces a nontradable good (N) and an exportable good (X). The foreign country produces its nontradable good (N*) and its export good (M*), the import good of

the home country. The production structure described here is based on constant returns to scale in production and competitive markets in both countries. For notational simplicity, the production process is described without any time subscripts attached to variables.

Home Country

The two sectors of production in the home country, X and N, use each other's output as input in own production processes. Production and gross output in each sector is assumed to be separable into value added (returns to labor and capital) in each sector and input from the other domestic sector. Value added in sectors X and N, V_x and V_n , are functions of labor and capital in respective sector. Assuming a Cobb-Douglas functional form and constant returns to scale with respect to labor and capital, value added in sectors X and N in the home country may be expressed as:

$$V_x = L_x^{(1-c_x)} K_x^{c_x},$$

$$V_n = L_n^{(1-c_n)} K_n^{c_n},$$

where L and K denote labor and capital used in respective sectors. Expressing in natural logarithm, the value added functions may be rewritten as:

$$v_x = (1 - c_x)l_x + c_x k_x, \quad (3.10)$$

$$v_n = (1 - c_n)l_n + c_n k_n. \quad (3.11)$$

Now, the lowercase letters l and k denote natural logarithms of labor (L) and capital (K) used in respective sectors.

Gross output in each sector is assumed to be a function of value added in that sector and input from the other domestic sector plus a sector specific technology factor. Again assuming a Cobb-Douglas functional form and constant returns to scale, gross output in sectors X and N, Q_x and Q_n , are given by⁶:

$$Q_x = A_x V_x^{c_{x1}} Q_{nx}^{c_{x2}}, \quad A_x = e^{h_x t}, \quad (3.12)$$

$$Q_n = A_n V_n^{c_{n1}} Q_{xn}^{c_{n2}}, \quad A_n = e^{h_n t}, \quad (3.13)$$

where Q_{nx} is the quantity of good N used in the production of Q_x , Q_{xn} is the quantity of good X used in the production of Q_n , and $c_{x1} + c_{x2} = c_{n1} + c_{n2} = 1$.

Prices

Given above technology, optimization may be performed sequentially in two stages. In the first stage V_x and V_n are chosen through optimal choice of labor and capital (L and K). Having chosen V_x and V_n in the earlier stage, in the second stage Q_x and Q_n are chosen through the choice of Q_{nx} and Q_{xn} . For simplicity of exposition, it is assumed that there is no depreciation of capital and no investment. Hence, capital is assumed to be fixed. Let P_{vx} and P_{vn} denote the value added deflators (prices of a unit of value added) in sectors X and N, respectively. Then P_{vx} and P_{vn} are related to the parameters of the production functions as follows: cost functions in sectors X and N are

$$C_x = V_x \left(\frac{W}{1 - c_x} \right)^{(1-c_x)} \left(\frac{r}{c_x} \right)^{c_x},$$

$$C_n = V_n \left(\frac{W}{1 - c_n} \right)^{(1-c_n)} \left(\frac{r}{c_n} \right)^{c_n}.$$

The marginal cost functions are:

$$MC_x = \left(\frac{W}{1-c_x} \right)^{(1-c_x)} \left(\frac{r}{c_x} \right)^{c_x}, \quad (3.14)$$

$$MC_n = \left(\frac{W}{1-c_n} \right)^{(1-c_n)} \left(\frac{r}{c_n} \right)^{c_n}. \quad (3.15)$$

Expressing in natural logarithm, the marginal cost functions are:

$$\text{Ln}(MC_x) = (1-c_x)\text{Ln}\left(\frac{W}{1-c_x}\right) + c_x\text{Ln}\left(\frac{r}{c_x}\right),$$

$$\text{Ln}(MC_n) = (1-c_n)\text{Ln}\left(\frac{W}{1-c_n}\right) + c_n\text{Ln}\left(\frac{r}{c_n}\right).$$

Equating price with marginal cost, value added deflators for sectors X and N may be written as:

$$\text{Ln}(P_{vx}) = (1-c_x)\text{Ln}\left(\frac{W}{1-c_x}\right) + c_x\text{Ln}\left(\frac{r}{c_x}\right),$$

$$\text{Ln}(P_{vn}) = (1-c_n)\text{Ln}\left(\frac{W}{1-c_n}\right) + c_n\text{Ln}\left(\frac{r}{c_n}\right).$$

Prices of goods X and N may be derived in the same manner. Given production functions for sectors X and N, the cost functions are:

$$C_x^g = \frac{Q_x}{A_x} \left(\frac{P_{vx}}{c_{x1}} \right)^{c_{x1}} \left(\frac{P_n}{c_{x2}} \right)^{c_{x2}},$$

$$C_n^g = \frac{Q_n}{A_n} \left(\frac{P_{vn}}{c_{n1}} \right)^{c_{n1}} \left(\frac{P_x}{c_{n2}} \right)^{c_{n2}},$$

where C denotes total cost, P_x and P_n are prices of gross output of goods X and N respectively, the superscript g on C implies that the costs are those of gross output of sectors X and N . Again equating price with marginal cost, the price equations (in natural logarithm) may be written as:

$$\ln(P_x) = c_{x1} \ln\left(\frac{P_{vx}}{c_{x1}}\right) + c_{x2} \ln\left(\frac{P_n}{c_{x2}}\right) - h_x t, \quad (3.16)$$

$$\ln(P_n) = c_{n1} \ln\left(\frac{P_{nx}}{c_{n1}}\right) + c_{n2} \ln\left(\frac{P_x}{c_{n2}}\right) - h_n t. \quad (3.17)$$

Foreign Country

Production and cost in the foreign country have similar structure as those in the home country. The only difference is that the foreign country uses an imported input (for example oil) in its production process in both sectors. Assumptions of linear homogeneity of the production functions and competitive market environment are maintained. Further, as in the home country, it is assumed that production functions in sectors M^* and N^* in the foreign country are separable in value added in respective sectors, inputs of other domestic sector, and the imported good (I)⁷. Value added in sectors M^* and N^* are assumed to have the following structure:

$$V_m^* = (L_m^*)^{1-c_m^*} (K_m^*)^{c_m^*},$$

$$V_n^* = (L_n^*)^{1-c_n^*} (K_n^*)^{c_n^*}.$$

Gross output in sectors M^* and N^* are described by the following production functions:

$$Q_m^* = A_m^* (V_m^*)^{c_{m1}^*} (Q_{mm}^*)^{c_{m2}^*} (I_m^*)^{c_{m3}^*}, \quad A_m^* = e^{h_m^* t},$$

$$Q_n^* = A_n^* (V_n^*)^{c_{n1}^*} (Q_{nn}^*)^{c_{n2}^*} (I_n^*)^{c_{n3}^*}, \quad A_n^* = e^{h_n^* t},$$

where Q_m^* and Q_n^* are quantities of goods M^* and N^* produced, Q_{mm}^* is the quantity of good N^* used in the production of Q_m^* , Q_{nn}^* is the quantity of good M^* used in the production of Q_n^* , I_m^* and I_n^* are quantities of the imported input used in producing Q_m^* and Q_n^* , respectively. The assumption of constant returns to scale in both sectors imply $c_{m1}^* + c_{m2}^* + c_{m3}^* = 1$ and $c_{n1}^* + c_{n2}^* + c_{n3}^* = 1$.

Price determination in the foreign country is similar to that in the home country. Following the two-stage optimization process used in case of the home country, value added deflators in sectors M^* and N^* , P_{vm}^* and P_{vn}^* , are given by (in natural logarithm):

$$\text{Ln}(P_{vm}^*) = (1 - c_m^*) \text{Ln}\left(\frac{W^*}{1 - c_m^*}\right) + c_m^* \text{Ln}\left(\frac{r^*}{c_m^*}\right), \quad (3.18)$$

$$\text{Ln}(P_{vn}^*) = (1 - c_n^*) \text{Ln}\left(\frac{W^*}{1 - c_n^*}\right) + c_n^* \text{Ln}\left(\frac{r^*}{c_n^*}\right). \quad (3.19)$$

Similarly, setting prices equal to marginal costs, prices of gross output of sectors M^* and N^* , P_m^* and P_n^* , may be expressed as:

$$\text{Ln}(P_m^*) = c_{m1}^* \text{Ln}\left(\frac{P_{vm}^*}{c_{m1}^*}\right) + c_{m2}^* \text{Ln}\left(\frac{P_n^*}{c_{m2}^*}\right) + c_{m3}^* \text{Ln}\left(\frac{P_1}{c_{m3}^*}\right) - h_m^* t, \quad (3.20)$$

$$\text{Ln}(P_n^*) = c_{n1}^* \text{Ln}\left(\frac{P_{vn}^*}{c_{n1}^*}\right) + c_{n2}^* \text{Ln}\left(\frac{P_m^*}{c_{n2}^*}\right) + c_{n3}^* \text{Ln}\left(\frac{P_1}{c_{n3}^*}\right) - h_n^* t, \quad (3.21)$$

where P_1 is the exogenously given price of the imported raw material, and as assumed in the case of production functions, $c_{m1}^* + c_{m2}^* + c_{m3}^* = 1$ and $c_{n1}^* + c_{n2}^* + c_{n3}^* = \cdot$.

Production structures in the home and in the foreign country as described above determine price structures in the two countries. The assumptions regarding production technologies in the two countries are made keeping in mind the objectives of the dissertation while maintaining simplicity of the analysis as much as possible. The above assumptions about technologies in the two countries allow us to see how shocks to production affect prices in the two countries, the real exchange rate between the two countries, and the current account balance of the home country.

Preferences

In this section we describe the preference structure of the agents in the two countries. Demand functions are derived via intertemporal utility maximization. These demand functions, along with the production technology described in the previous section, will be used to highlight the differential effects of transitory and permanent shocks⁸. It is assumed that there is one representative agent in each country. For simplicity, it is assumed that there are only two periods, current period (period 0) and future period (period 1). Agents from both countries can borrow and lend freely in the world capital market at world interest rate. It assumed that the world interest rate at which agents can lend and borrow is exogenously determined. The representative agent in each country consumes its own nontradable good, its export good, as well as its import good

(export good of the foreign country). Good X is treated as the numeraire so that $P_x \equiv 1$.

Home Country

The representative agent in the home country maximizes lifetime utility subject to lifetime budget constraint. The periodic budget constraints facing the agent are:

$$\text{period 0: } c_{x0} + P_{m0}c_{m0} + P_{n0}c_{n0} = Q_{x0} + P_{n0}Q_{n0} - (1 + r_{x,-1})B_{-1} + B_0,$$

$$\text{period 1: } c_{x1} + P_{m1}c_{m1} + P_{n1}c_{n1} = Q_{x1} + P_{n1}Q_{n1} - (1 + r_{x,0})B_0$$

where Q_{i0} = Output of good i in period 0 ($i = x, n$),

Q_{i1} = Output of good i in period 1 ($i = x, n$),

P_{i0} = Price of good i in period 0 ($i = x, m, n$),

P_{i1} = Price of good i in period 1 ($i = x, m, n$),

c_{i0} = Consumption of good i in period 0 ($i = x, m, n$),

c_{i1} = Consumption of good i in period 1 ($i = x, m, n$),

B_k = Level of accumulated debt at the beginning of period k ,

r_{xt} = Interest rate, in terms of X, on one period debt accumulated in period t .

The representative agent's utility function is defined over the consumption set $\{c_{x0}, c_{x1}, c_{m0}, c_{m1}, c_{n0}, c_{n1}\}$. For simplicity, it is assumed that the utility function is separable in terms of aggregate consumption in each period. In other words, the individual's utility function may be expressed as a function of two components:

$C_0 = C_0(c_{x0}, c_{m0}, c_{n0})$, and $C_1 = C_1(c_{x1}, c_{m1}, c_{n1})$. The functions $C_0 = C_0(c_{x0}, c_{m0}, c_{n0})$, and $C_1 = C_1(c_{x1}, c_{m1}, c_{n1})$ may be viewed as subutility functions expressing utility derived by

the agent in each period as functions of consumption of all goods in each period. It is further assumed that both $C_0 = C_0(c_{x0}, c_{m0}, c_{n0})$, and $C_1 = C_1(c_{x1}, c_{m1}, c_{n1})$ are linearly homogeneous in their respective arguments.

The representative consumer's problem is to choose $\{c_{x0}, c_{x1}, c_{m0}, c_{m1}, c_{n0}, c_{n1}\}$ to maximize lifetime utility given by $U = U[C_0(c_{x0}, c_{m0}, c_{n0}), C_1(c_{x1}, c_{m1}, c_{n1})]$ subject to the lifetime budget constraint. Assuming no initial net indebtedness ($B_{-1} = 0$), the consolidated budget constraint is given by:

$$c_{x0} + P_{m0}c_{m0} + P_{n0}c_{n0} + \alpha_{x1}(c_{x1} + P_{m1}c_{m1} + P_{n1}c_{n1}) = Q_{x0} + P_{n0}Q_{n0} + \alpha_{x1}(Q_{x1} + P_{n1}Q_{n1})$$

where $\alpha_{x1} = \frac{1}{(1+r_{x0})}$ is the discount factor to convert period 1 variables to respective present value in period 0.

We define:

$$W_0 = Q_{x0} + P_{n0}Q_{n0} + \alpha_{x0}(Q_{x1} + P_{n1}Q_{n1}) = \text{present value in period 0 of lifetime wealth.}$$

$$Z_t = c_{xt} + P_{mt}c_{mt} + P_{nt}c_{nt} = \text{aggregate consumption expenditure in period } t \text{ (} t = 0, 1 \text{).}$$

Given the preference structure described above, the entire utility maximization problem of the representative consumer may be divided into two parts:

- (i) the temporal allocation of Z_t among c_{xt} , c_{mt} , and c_{nt} ($t = 0, 1$);
- (ii) intertemporal allocation of W_0 between Z_0 and Z_1 subject to $Z_0 + \alpha_{x1}Z_1 = W_0$.

The first stage of the problem may be viewed as one of minimizing the cost (spending) Z_t of obtaining subutility C_t ($t = 0, 1$). The assumption that C_t is linearly homogeneous with

respect to its arguments implies that the cost or the spending function may be written as:

$$Z_t = P_t(1, P_{mt}, P_{nt})C_t,$$

where $P_t(1, P_{mt}, P_{nt})$ may be viewed as the marginal cost of obtaining subutility C_t . The assumption of linear homogeneity of C_t provides the basis to express the marginal cost of C_t as a function of relative prices only. Defining as an index number, P_t may be expressed as:

$$P_t = (1)^{\beta_{xt}} (P_{mt})^{\beta_{mt}} (P_{nt})^{\beta_{nt}},$$

where $\beta_{xt} + \beta_{mt} + \beta_{nt} = 1$. Such an index number can be derived through expenditure minimization with the linear homogeneous subutility function of the following form::

$$C_t = (c_{xt})^{\beta_{xt}} (c_{mt})^{\beta_{mt}} (c_{nt})^{\beta_{nt}}, \quad \beta_{xt} + \beta_{mt} + \beta_{nt} = 1,$$

where the variable c_{it} ($i = x, m, n$) represents the consumption of good i in period t . The exponent β_{it} ($i = x, m, n$) has the property that it represents the share of expenditure on good i in period t total consumption expenditure. In other words, β_{it} ($i = x, m, n$) represents the budget share of the i^{th} good in period t consumption expenditure. The price index defined above, P_t , displays the usual properties of a price index.

In the second stage of the consumer's optimization process, the problem is to choose C_0 and C_1 to maximize lifetime utility, $U = U[C_0, C_1]$, subject to lifetime budget constraint given by:

$$P_0 C_0 + \alpha_{x1} P_1 C_1 = W_0.$$

The budget constraint may be rewritten as:

$$C_0 + \alpha_{c1}C_1 = W_{c0},$$

where

$$\alpha_{c1} = \alpha_{x1} \left(\frac{P_1}{P_0} \right) \text{ and } W_{c0} = \frac{W_0}{P_0}.$$

Here α_{c1} may be interpreted as the real discount factor applicable to total consumption expenditure in period 1. Solving the allocation problem of the second stage, the periodic demand function of the following form may be obtained:

$$C_t = C_t(\alpha_{c1}, W_{c0}).$$

In the periodic demand function above, α_{c1} represents the intertemporal price ratio of consumption in period 1 in terms of that in period 0.

Foreign Country

The representative agent in the foreign country is assumed to have preferences that are identical to that of the representative agent of the home country. The utility function of this agent is defined over the consumption set $\{c_{x0}^*, c_{x1}^*, c_{m0}^*, c_{m1}^*, c_{n0}^*, c_{n1}^*\}$, where c_{ij}^* ($i = X, M, \text{ and } N^*, j = 0, 1$) is the consumption of good i in period j by the representative agent of the foreign country. Since identical preference structure is assumed, equilibrium budget shares are exactly the same as those for the representative agent of the home country. In other words, in equilibrium, $\beta_{xt}^* = \beta_{xt}$, $\beta_{mt}^* = \beta_{mt}$, and $\beta_{nt}^* = \beta_{nt}$, ($t = 0, 1$). Since perfectly integrated capital market is assumed, the agent can borrow and lend freely in the world capital market at interest rate, r_{x0} . As has been

done for the home country, a price index for the foreign country can be defined as:

$$P_t^* = (1)^{\beta_m} (P_{mt})^{\beta_m} (P_{nt}^*)^{\beta_n}, \quad (t = 0, 1),$$

where P_{nt}^* is the price of foreign nontradable good (N^*). It is evident from the definition of price indices for the home and the foreign country that underlying these definitions we have assumed perfect commodity price arbitrage, and zero transportation cost. Despite the assumptions of perfect commodity price arbitrage, and zero transportation cost, P_t and P_t^* may diverge due to differential movements in P_{nt} and P_{nt}^* .

Current Account

Assuming no initial debt, the current account balance of the home country in period t may be defined as the difference between the real income and real consumption in that period. That is, the real current account balance in period 0 can be written as:

$$CA_0 = Y_0 - C_0(\alpha_{c1}, W_{c0}), \quad (3.22)$$

where

$$Y_0 = \frac{Q_{x0} + P_{n0}Q_{n0}}{P_0(P_{m0}, P_{n0})}. \quad (3.23)$$

For algebraic simplicity, henceforth we assume that $U = U(C_0, C_1)$ is a homothetic utility function. The effects of transitory and permanent shocks to terms of trade on the current account balance of period 0 can be obtained by differentiating (3.22) with respect to the variable subject to the shock. If there is a transitory terms of trade increase in period 0 ($dp_{m0} > 0$, $dp_{m1} = 0$), then the effect on the current account balance of period 0 is

given by:

$$\frac{d(CA_0)}{d \log p_{m0}} = \frac{\partial(CA_0)}{\partial \log p_{m0}} + \sum_{t=0}^1 \frac{\partial(CA_0)}{\partial \log p_{nt}} \cdot \frac{d \log p_{nt}}{d \log p_{m0}}. \quad (3.24)$$

Similarly, the effect of an anticipated shock to terms of trade in period 1 ($dp_{m0} = 0$, $dp_{m1} > 0$) on the current account balance of period 0 is given by:

$$\frac{d(CA_0)}{d \log p_{m1}} = \frac{\partial(CA_0)}{\partial \log p_{m1}} + \sum_{t=0}^1 \frac{\partial(CA_0)}{\partial \log p_{nt}} \cdot \frac{d \log p_{nt}}{d \log p_{m1}}. \quad (3.25)$$

The effect of a permanent shock to terms of trade in period 0 on the current account balance of period 0 is obtained by setting $dp_{m0} = dp_{m1} = dp_m$, and obtaining the total derivative of CA_0 with respect to p_m . That effect is given by:

$$\frac{d(CA_0)}{d \log p_m} = \frac{\partial(CA_0)}{\partial \log p_m} + \sum_{t=0}^1 \frac{\partial(CA_0)}{\partial \log p_{nt}} \cdot \frac{d \log p_{nt}}{d \log p_m}. \quad (3.26)$$

Now differentiating (3.22) and (3.23) with respect to p_{n0} , p_{n1} , and p_m , and simplifying, we can obtain:

$$\frac{\partial CA_0}{\partial \log p_{n0}} = \beta_{n0} [(1 - \lambda_0) + \gamma \sigma] C_0, \quad (3.27)$$

$$\frac{\partial CA_0}{\partial \log p_{n1}} = -\beta_{n1} (\gamma \sigma) C_0, \quad (3.28)$$

$$\frac{\partial CA_0}{\partial \log p_n} = \beta_n (1 - \lambda_0) C_0, \quad (3.29)$$

where λ_0 is the ratio of real domestic income (Y_0) to real spending in period 0.

Similarly, we can obtain:

$$\frac{\partial CA_0}{\partial \log p_{m0}} = \beta_{m0}[(1 - \lambda_0) + \gamma(\sigma - 1)]C_0, \quad (3.30)$$

$$\frac{\partial CA_0}{\partial \log p_{m1}} = -\beta_{m1}\gamma(\sigma - 1)C_0, \quad (3.31)$$

$$\frac{\partial CA_0}{\partial \log p_m} = \beta_m(1 - \lambda_0)C_0. \quad (3.32)$$

Using equations (3.27) through (3.32), and the expressions for $\frac{d \log p_{nt}}{d \log p_{mt}}$ and $\frac{d \log p_{nt}}{d \log p_m}$

(these expressions are derived in Appendix C), we can see how the current account of the home country in period 0 is affected by temporary and permanent changes in terms of trade. If, for example, there is a permanent deterioration of the terms of trade in period 0 ($dp_{m0} = dp_{m1} = dp_m > 0$), then the effect of such a change can be expressed, after simplification, as⁹:

$$\frac{d(CA_0)}{d \log p_m} = \beta_m(1 - \lambda_0)C_0 + \beta_n(1 - \lambda_0)C_0(\sigma_{nm} - 1)\Lambda_1, \quad (3.33)$$

where

$$\Lambda_1 = \frac{\beta_m}{(\beta_m \sigma_{nm} + \beta_x \sigma_{nx})} > 0.$$

It is evident from equation (3.33) that the effect of a permanent change in the terms of trade on the current account balance depends critically on the initial net indebtedness (λ_0). If the initial current account is balanced ($\lambda_0 = 1$), then the permanent deterioration of the terms of trade lowers domestic real income and spending by the same amount, and the current account balance is unaffected. Similarly, assuming that initially the current

account is balanced, the effect of a temporary deterioration in the terms of trade in period 0 ($dp_{m0} > 0$, $dp_{m1} = 0$) on the current account of period 0 is given by:

$$\frac{d(CA_0)}{d \log p_{m0}} = \beta_m \gamma (\sigma - 1) C_0 + \beta_n \gamma \sigma (\sigma_{nm} - \sigma) C_0 \Lambda_2, \quad (3.34)$$

where

$$\Lambda_2 = \frac{\beta_m}{\beta_m \sigma_{nm} + \beta_x \sigma_{nx} + \beta_n \sigma} > 0.$$

Also, it can be shown that $\frac{d(CA_0)}{d \log p_{m1}} = -\frac{d(CA_0)}{d \log p_{m0}}$.

Thus, starting from an initial current account balance (that is, $\lambda_0 = 1$ in equation (3.33)) a permanent terms of trade change will leave the current account unaffected, while the effect of a temporary terms of trade is not necessarily zero. The magnitude and direction of the effects of a temporary shock to terms of trade on current account balance of period 0 depends on the values of the parameters of the model. In fact, the first term of equation (3.34) may be interpreted as the direct effect while the second term may be viewed as the indirect effect of a temporary terms of trade increase on current account of period 0. Looking at the first term of equation (3.34), we observe that the direct effect is positive (that is, current account improves) if the intertemporal elasticity of substitution (σ) is greater than unity. Otherwise, the direct effect of a temporary terms of trade increase in period 0 has a negative effect on period 0 current account. Further, it is evident that the indirect effect (given by the second term of equation (3.34)) of a temporary terms of trade deterioration in period 0 depends on the relative values of the

intratemporal elasticity of substitution between domestic nontradable good and the import good (σ_{nm}), and the intertemporal elasticity of substitution (σ). If the temporal elasticity of substitution (σ_{nm}) dominates the intertemporal elasticity of substitution, then a temporary deterioration in the terms of trade in period 0 improves the current account in period 0. Otherwise, such a terms of trade deterioration leads to a current account deficit. Thus, *a priori* we can not unambiguously predict the effects of a temporary terms of trade shock on the current account balance in that period.

The conceptual model described above highlights the roles of income and relative prices in determining the current account balance of a country. However, several authors have emphasized the importance of government consumption as a determinant of current account balance. For instance Ahmed (1986, 1987) emphasizes government spending as one of the determinants of the trade balance of the United Kingdom. Lee and Enders (1991) focus on the relation between budget deficit and current account deficit of the United States¹⁰. In our study government consumption is included as an exogenous variable affecting the current account balance. However, no effort has been made to distinguish between transitory and permanent changes in government consumption as a determinant of current account balance.

Description of Data and Variables

In the conceptual model outlined above, we have discussed the effects of temporary and permanent changes in income and terms of trade on the current account balance of the home country. In empirical tests of the hypotheses that will be discussed in

later chapters, real exchange rate has been used as the macroeconomic analog of terms of trade. Essentially, terms of trade attempts to measure the relative prices of imports in terms of domestic exportable good. In fact terms of trade have been defined in the literature in more than one way. One definition of terms of trade is the ratio of import price to export price. Following the usual practice in empirical literature in international economics, we use the aggregate foreign price level relative to domestic price level as the empirical counterpart of the concept of terms of trade used in the last section.

Accordingly the real exchange rate (of the home country) is defined here as the relative aggregate price of the foreign country in terms of that of the home country. Denoting the real exchange rate in period t by R_t , the definition above may be written as:

$$R_t = \frac{P_t^*}{P_t}$$

From the construction of P_t and P_t^* it follows that R_t may change in response to differential movements in P_{nt} and P_{nt}^* . Given the structure of the model outlined in the last section, even though we allow complete price arbitrage, no transportation costs, and no impediments to trade, P_{nt} and P_{nt}^* may diverge due to exogenous shocks to oil prices, or differential rate of technological progress across sectors and between the two countries.

This study uses data on current account balance of the United States and Japan. The analysis is performed in terms of the US vs. the rest of the world, and Japan vs. rest of the world. In the case of US current account balance, the following country data are aggregated and treated as the rest of the world: Australia, Austria, Canada, France,

Germany, Italy, Japan, South Korea, Sweden, and the United Kingdom. Similarly, in analyzing the current account balance of Japan, the rest of the world comprises of the following countries: Australia, Austria, Canada, France, Germany, Italy, South Korea, Sweden, the United States, and the United Kingdom. The variables used in the analysis are real GNP/GDP, real government consumption of goods and services of each of the countries, and multilateral effective real exchange rate of the United States and Japan, as reported in the International Financial Statistics. Current account balance is defined as the balance of goods and services. All income and government consumption are defined in 1985 constant price and then converted to US\$. In the analysis of US current account, income and government consumption of the 10 countries (Australia, Austria, Canada, France, Germany, Italy, Japan, South Korea, Sweden, and the United Kingdom) are aggregated to obtain the rest of the world income and government consumption. Similarly, for Japan, income and government consumption of Australia, Austria, Canada, France, Germany, Italy, South Korea, Sweden, the United States, and the United Kingdom are aggregated to obtain the rest of the world income and government consumption. Whenever available, GNP is used as measure of national income. Otherwise, GDP is used. Specifically the following definitions of the variables are used:

Current account = $\text{Ln}(\text{export of goods \& services}) - \text{Ln}(\text{import of goods \& services})$.

Relative income = $\text{Ln}(\text{domestic income}) - \text{Ln}(\text{world income})$.

Real exchange rate = multilateral effective exchange rate reported by IMF.

Relative government consumption = $\text{Ln}(\text{domestic government consumption}) -$
 $\text{Ln}(\text{government consumption in the rest of the world}).$

In these definitions, Ln denotes natural logarithm. All income, government consumption, exports, and imports are defined in 1985 constant price and then converted into US \$.

All data used in the study are obtained from International Monetary Fund CD-ROM. The study uses quarterly data and the period covered is 1973:1 to 1993:2, the floating exchange rate period.

Endnotes

1. Detailed derivation of the consumption function is presented in appendix A1.
2. It should be noted that although the assumptions about the specific utility function, the restrictions on the behavior of Y_t , and constant R allow us to derive a consumption function equation that is very convenient to work with, it has some undesirable properties. Specifically, the equilibrium law of motion for asset holding satisfies the following condition:

$$A_{t+1} = \left(\frac{1}{R}\right)A_t + \frac{\alpha R}{(R-1)} + RY_t - \left(R - \frac{1}{\beta R}\right) \sum_{j=0}^{\infty} R^{-j} E_t Y_{t+j} \text{ (see Sargent 1987, pp. 365 - 366, for details).}$$

When $\beta R > 1$, the above condition implies that $E_t A_{t+j}$ converges as $j \rightarrow \infty$.

Further, it can be shown that when $\beta R > 1$, $\lim_{j \rightarrow \infty} E_t C_{t+j} = \frac{a_1}{a_2}$. This implies that if $\beta R > 1$, the system can be expected to converge to a level of asset holdings such that consumption varies stochastically around the bliss level of consumption. This is an unsatisfactory result because it implies that eventually the marginal utility of consumption fluctuates around zero. Despite such undesirable limiting property, the model generates sensible results at low consumption levels. This is one reason that we impose restriction specified by equation (3.3).

3. This specification for the stochastic process describing the evolution of the real income (Y_t) has the implication that Y_t may grow indefinitely large. That may lead to levels of consumption so high that the restriction specified by equation (3.3) may be violated. However, we maintain the assumption that the economy operates at levels where such a situation does not arise.
4. The effects of permanent and transitory changes in income in period t on the current account balance in period t are illustrated in appendix A2.
5. The production technology used in this exposition is very similar to the one used by Marston and Turnovsky (1985). Similar specification of production technology has been used by Bruno and Sachs (1979), and Bruno (1981) in the context of analyzing the effects of oil price increases. All these studies use a Cobb-Douglas specification for the value-added functions, and a CES specification for the gross output functions which are assumed to be function of value added and a raw material. We assume that both the value-added functions, as well as the gross output functions are of Cobb-Douglas form with constant returns to scale. In this exposition, instead of using a raw material, it is

assumed that both sectors use each others output as input. A second modification is that we incorporate technical progress parameter. The purpose of this modification is to bring out the point that differential technical progress in the two sectors, as well as in the two countries are sources of changes in relative prices which will ultimately affect the real exchange rate (defined later) between the two countries. The choice of Cobb-Douglas functional form is purely for the sake of simplicity of analysis.

6. As Marston and Turnovsky (1985) mentions (see their reference to Arrow (1974)), gross output must be separable in value added and other inputs for the value added function to be well defined.

7. The assumption that the home country does not import any input (oil) from the third country while the foreign country does so is made to highlight the fact that different countries have different degrees of dependence on imported raw materials. Therefore, major changes in the prices of raw materials may affect the prices and competitiveness of different countries in very different ways. Ultimately these differential impacts of raw material price changes will be reflected on the real exchange rate between any two countries that are affected differently.

8. The exposition in this section draws heavily from Frankel and Razin (1987) and Ostry (1988).

9. These results are derived under the condition that the expenditure shares remain unchanged between periods 0 and 1.

10. For theoretical discussion on this issue, see Frankel and Razin (1987) and the references there.

CHAPTER 4. UNIVARIATE PROPERTIES OF THE DATA: UNIT ROOTS

Unit Roots in Time Series

Traditional macroeconomic literature was built on the presumption that real macroeconomic variables could be described by stationary stochastic processes. In other words, such literature implicitly assumed that the data generating processes could be modeled as stable stochastic processes (i.e., the characteristic roots of the polynomial describing the autoregressive components of the data series were less than unity in absolute value). Although theoretical implications of the presence of unit roots (in the autoregressive components of data series) greater than or equal to unity (in absolute value) were long known, econometric studies tended to ignore such a possibility on the basis of the commonly held belief that real macroeconomic variables were indeed stationary.

Until recently, traditional macroeconomic models used to decompose a real economic variable into a secular trend, a cyclical component, and an irregular component. The secular trend was thought to be determined by long-run factors such as population growth, technological progress, human and non-human capital accumulation. These determinants of the trend were believed to change only slowly and over time, where the rates of change were believed to be constant. Consequently, the trend component in any macroeconomic series (if any) was modeled as a deterministic trend and approximated by

a polynomial in time. Thus, although it was recognized that the presence of the trend component rendered the data series non-stationary in mean, the deterministic nature of the trend allowed an econometrician to remove the trend (detrending of variables) by using a linear (or polynomial) time trend in the regression model. The cyclical and the irregular components were believed to be stationary stochastic processes. Hence, the issue of non-stationarity of real economic variables was largely ignored.

In their seminal work, Nelson and Plosser (1981) demonstrates that important macroeconomic variables are characterized not by deterministic trends, but by stochastic trends. In other words, their general conclusion is that the stochastic processes describing most macroeconomic variables contains at least one unit root in their autoregressive components. That is, most macroeconomic variables are difference stationary rather than trend stationary, as was presumed in traditional econometric literature.

The Nelson-Plosser (1981) study stimulated tremendous interest in the macroeconometric research, and over the last decade an enormous literature has grown that deals with the issue of potential unit roots in macroeconomic data series and their implications for econometric analysis. Despite some arguments to the contrary by Perron (1989, 1992), Rudebusch (1992, 1993), and others, the overwhelming opinion of the profession on this issue is that indeed most macroeconomic variables are characterized by the presence of unit roots.

Notwithstanding the contrary opinions of Perron (1989, 1992) and others, there is general agreement that in any econometric study that uses macroeconomic data series, it

is of utmost importance to examine first whether any variable used in the study contains one or more unit roots in its autoregressive component. This is because the presence of unit roots has profound implications for model building, estimation strategy, and statistical inference. For instance, if a data series actually contains a unit root then the effects of exogenous shocks on the variable in question are very different from the case where the series does not contain any unit root. More explicitly, if the data series contains a unit root then the exogenous disturbances have permanent effects on the variable in question in the sense that the effects of the shocks do not disappear over time. That is, shocks to such a variable has permanent effects. On the other hand, if there is no unit root in the data series then exogenous disturbances only have transient effects in the sense that the effects of such disturbances die down over time and, in the long-run, the variable reverts back to its long-run path. From pure statistical standpoint, presence of unit roots in the data series renders many standard methods of estimation and inference invalid. This is because the sampling distributions of most estimators are based on the assumption of covariance stationary behavior of the variables and error terms. If unit roots are present in the data, the standard sampling distribution results of most estimators do not remain valid and inferences made on the basis of the standard sampling distributions of the estimators no longer remain statistically reliable.

Further, presence of unit roots in the variables has serious implications for standard regression models. The classical regression model is based on the assumptions that both the dependent and the explanatory variables are stationary, and that the error

term has zero mean and finite variance. Presence of non-stationary variables may lead to what Granger and Newbold (1974) calls *spurious regression* results in the sense that estimation of a regression equation with non-stationary variables may generate high R^2 (coefficient of determination) and t-statistics that may appear to be significant when there exists no meaningful relationship between the dependent variable and the explanatory variables. It is therefore important that the variables used in the empirical analysis be examined carefully for the presence of potential unit roots.

Testing for Unit Roots

Testing for stationarity of a variable is equivalent to testing for the presence of unit roots in the autoregressive component of the stochastic variable. The methodology used to test for the presence of such unit roots is based on variations of the Augmented Dickey-Fuller (ADF) test. This section provides a brief discussion of the ADF testing procedure.

The Augmented Dickey-Fuller Test

To illustrate, let us consider the case where the data generating process for the stochastic variable Y_t can be well represented by the p^{th} order autoregressive process:

$$Y_t = a_0 + \sum_{i=1}^p a_i Y_{t-i} + \varepsilon_t, \quad (4.1)$$

where the disturbance term ε_t follows a covariance stationary process.

Equation (4.1) may be rewritten as:

$$\Delta Y_t = a_0 + \gamma Y_{t-1} + \sum_{i=2}^p \beta_i \Delta Y_{t+i-i} + \varepsilon_t, \quad (4.2)$$

where

$$\Delta Y_t = Y_t - Y_{t-1} = \text{first difference of } Y_t, \gamma = -\left(1 - \sum_{i=1}^p a_i\right), \text{ and } \beta_i = -\sum_{j=i}^p a_j.$$

If the coefficient γ in equation (4.2) is zero ($\gamma = 0$) then the equation is entirely in first difference, and therefore, equation (4.1) has a unit root. Hence, testing for the presence of a unit root focuses on testing for the null hypothesis that $\gamma = 0$ in equation (4.2). The Augmented Dickey-Fuller (ADF) test is essentially a test of the above null hypothesis. However, the standard t-statistic is not useful for testing the above null hypothesis. Rather, the ADF test is the appropriate tool for testing the above null hypothesis. However, the ADF test depends critically on the deterministic components included in the equation used to estimate the test statistic.

Dickey and Fuller developed the test statistic and calculated the critical values for three different representations of the data generating process:

Case I: The autoregressive representation of Y_t contains both an intercept term and a linear time trend. That is, Y_t may be represented as:

$$Y_t = a_0 + b_1 t + \sum_{i=1}^p a_i Y_{t-i} + \varepsilon_t, \text{ or equivalently,}$$

$$\Delta Y_t = a_0 + b_1 t + \gamma Y_{t-1} + \sum_{i=2}^p \beta_i \Delta Y_{t+1-i} + \varepsilon_t. \quad (4.3)$$

Testing for unit root now involves estimation of the above equation and compare the t-ratio of estimated γ ($\hat{\gamma}$) with the critical values of the τ_γ -statistic developed by Dickey and Fuller (1979).

Case II. The autoregressive representation of Y_t contains an intercept term, but no time trend. That is, Y_t may be represented as:

$$Y_t = a_0 + \sum_{i=1}^p a_i Y_{t-i} + \varepsilon_t, \text{ or equivalently,}$$

$$\Delta Y_t = a_0 + \gamma Y_{t-1} + \sum_{i=2}^p \beta_i \Delta Y_{t+i-i} + \varepsilon_t. \quad (4.4)$$

Testing for unit root (testing for the null hypothesis that $\gamma = 0$) follows exactly the same procedure as in case I. The only difference is that now the t-ratio of the estimated γ ($\hat{\gamma}$) is compared with the critical values of the τ_μ -statistic developed by Dickey and Fuller (1979).

Case III. In this case, Y_t contains no deterministic component and is well represented by:

$$Y_t = \sum_{i=1}^p a_i Y_{t-i} + \varepsilon_t, \text{ or equivalently,}$$

$$\Delta Y_t = \gamma Y_{t-1} + \sum_{i=2}^p \beta_i Y_{t+i-i} + \varepsilon_t. \quad (4.5)$$

The Augmented Dickey-Fuller test involves comparing the estimated t-ratio of the $\hat{\gamma}$ in equation (4.5) with the critical values of the τ -statistic in Dickey and Fuller (1979).

Practical Problems in Testing for Unit Roots

However, there are some important issues involved in implementing the ADF test.

The different forms of the ADF test are based on the assumption that the error term in

each representation of Y_t is identically and independently distributed (iid) with zero mean and constant variance. In reality, the true data generating process may contain both autoregressive and moving average components. In such a case estimation of the autoregressive equation (4.2) leaves the moving average component in the error term, and therefore, the iid assumption about the error process is violated. Another problem in implementing the test is that of choosing an appropriate lag-length for equations (4.3) through (4.5). Finally, as Campbell and Perron (1991) point out, the proper handling of the deterministic trend is a vital prerequisite for dealing with unit roots. Since it is not known *a priori* whether the deterministic components (intercept and time trend) should be included in estimating equations (4.3) through (4.5), the testing procedure should be careful about the inclusion of the intercept and the time trend in the equation.

The issue of potential moving average component in the error terms of equations (4.3) through (4.5) has been addressed by Said and Dickey (1984). They show that an unknown ARIMA(p,1,q) process can be well approximated by an ARIMA(n,1,0) autoregressive process of order $T^{1/2}$. Hence, a potentially mixed error process can be estimated using a finite autoregressive process. The ADF statistic may be applied to test may then be used to test for unit roots.

On the question of choosing appropriate lag-length for equations (4.3) through (4.5), Campbell and Perron (1991) suggest the following procedure. One should start with a relatively long lag-length n^* . If the t-statistic on lag n^* is statistically insignificant at

some specified critical value, the equation should be re-estimated with $(n^* - 1)$ lags. The procedure should be repeated until the t-statistic on the last lagged term is statistically significant. In this study the above procedure has been followed in selecting the lag-length to estimate the ADF statistics. Initially, equations are estimated with 12 lags, and standard t-statistic and F-statistic are used to reduce the number of lags in the equations. Once a tentative lag-length is selected, additional diagnostic checking is done to check if residual terms reveal any evidence of structural breaks or serial correlation. Specifically, plots of the residuals as well as their correlograms are examined for potential structural breaks and serial correlation. Formally, Ljung-Box (1978) Q-statistic is used to test for serial correlation in the residuals at lags 4 and above up to 24 lags.

Perhaps the most important issue in testing for unit roots using the ADF statistic is that of the treatment of deterministic components. The problem arises due to the fact that the distribution of the test statistics (τ_τ and τ_μ) are themselves dependent on the presence of particular deterministic variables included as regressors. To complicate the matter further, tests to determine the appropriateness of deterministic regressors are conditional on the presence of unit roots. On this issue, Campbell and Perron (1991) report the following results:

1. When the estimated regression includes at least all the deterministic elements in the actual data generating process, the distribution of $\hat{\gamma}$ is non-normal under

the null of a unit root. The distribution itself varies with the set of parameters included in the estimated equation.

2. If the estimated regression includes deterministic regressors that are not in the actual data generating process, the power of the unit root test against a stationary alternative decreases as additional deterministic regressors are added.
3. If the estimated regression omits an important deterministic trending variable present in the true data generating process, the power of the t-statistic goes to zero as sample size increases. If the estimated equation omits a non-trending variable, the t-statistic is consistent but its finite sample power decreases as the magnitude of the coefficient of the omitted variable increases.

The τ_τ , τ_μ , ϕ_1 , ϕ_2 , and ϕ_3 statistics (the ϕ_i statistics are test statistics to determine the appropriateness of including the intercept term and the linear trend in the autoregressive representation of Y_t , and are described later) have asymptotic distributions tabulated by Dickey and Fuller (1979, 1981). The critical values of the various statistics depend on sample size. However, the sample variance of $\{Y_t\}$ will be dominated by the presence of trend or drift, and as sample size increases, the τ_τ and τ_μ statistics converge to standard normal distribution. It is only when the time trend is zero in the estimated equation and in the true data generating process that the Dickey-Fuller distribution dominates.

The discussion above implies that in testing for stationarity of a data series, one may reach a wrong conclusion because of a misspecification of the deterministic

components in the regression equation. When the actual data generating process is unknown, Dolado, Jenkins, and Sosvilla-Rivero (1990) suggest the following sequential procedure in implementing the unit root tests:

1. One should start with the least restrictive of the plausible models (which will, in general, include an intercept term and a time trend in the regression equation) and use the τ_τ -statistic to test the null hypothesis of a unit root. Unit root tests generally have low power to reject the null hypothesis. Therefore, if the null of a unit root is *rejected*, one may draw the conclusion that the data series does not contain a unit root.
2. If the null of a unit root is *not rejected*, one should then test for the significance of the time trend under the null of a unit root by using the ϕ_3 -statistic described in Dickey and Fuller (1981) (the test is described later). If the time trend is not significant, then one should proceed to step 3. On the other hand, if the time trend is statistically significant, one should test for a unit root using the standard normal test. If such test rejects the null of a unit root, that implies no unit root in the data generating process. Otherwise, one may reach the conclusion that there is a unit root in the variable in question.
3. In this step, one should re-estimate the regression equation without a time trend term (i.e., estimate equation (4.4)) and use the τ_μ -statistic to test the null of a unit root in the variable. If the null of a unit root is *not rejected*, one should test for the significance of the intercept term in the regression equation

using the ϕ_1 -statistic described in Dickey and Fuller (1981). If the intercept term is not significant then one should go to step 4. If the intercept term is significant, one may use the standard normal distribution in testing the null hypothesis that $\gamma = 0$ in equation (4.4) and accept the conclusion of this test.

4. In this step, regression equation should be re-estimated without the intercept term and any time trend (i.e., estimate equation (4.5)). Now the τ -statistic should be used to test for the presence of a unit root and conclusion should be drawn on the basis of this test result.

In this study, the sequential procedure outline above has been followed in the empirical implementation of the ADF test.

Deterministic Components in the Data and Unit Roots Tests

It is evident from previous discussion that proper specification of the deterministic components in the regression equation is a prerequisite for the ADF test. In order to determine the appropriate deterministic terms to be included in the unit root test, the Dickey-Fuller (1981) procedure is followed. Their procedure is as follows:

Step 1. The regression equation is estimated with an intercept term and a linear time trend. Under the null of a unit root, test for the significance of the time trend can be performed by obtaining the F-statistic for the joint null hypothesis

$$H_0: \quad \gamma = \beta_1 = 0$$

$$H_A: \quad \text{Not } H_0 .$$

The estimated value of the F-statistic is then compared with the tabulated value of the ϕ_3 -statistic in Dickey and Fuller (1981).

Step 2. If the null hypothesis is not rejected in step 1 then equation (4.4) is estimated. Under the null of a unit root, test for the significance of the intercept term is equivalent to the test of the joint null and alternative hypotheses:

$$H_0: \quad \gamma = \beta_0 = 0$$

$$H_A: \quad \text{Not } H_0 .$$

The test involves obtaining the estimated F-statistic and comparing it with the tabulated values of the ϕ_1 -statistic reported in Dickey and Fuller (1981).

The test statistics for various null and alternative hypotheses are summarized below:

- | | | |
|---------|--|---|
| Case 1. | Estimated regression equation: | $\Delta y_t = \beta_0 + \beta_1 t + \gamma y_{t-1} + \sum_{i=2}^p \beta_i \Delta y_{t+1-i} + \varepsilon_t$ |
| | Test for unit root: | $H_0: \gamma = 0; H_A: \gamma < 0$
Test statistic: τ_τ -statistic. |
| | Test of significance of time trend in the presence of a unit root: | $H_0: \beta_1 = \gamma = 0; H_A: \text{Not } H_0.$
Test statistic: ϕ_3 -statistic. |
| Case 2. | Estimated regression equation: | $\Delta y_t = \beta_0 + \gamma y_{t-1} + \sum_{i=2}^p \beta_i \Delta y_{t+1-i} + \varepsilon_t$ |
| | Test for unit root: | $H_0: \gamma = 0; H_A: \gamma < 0$
Test statistic: τ_μ -statistic. |
| | Test of significance of time trend in the presence of a unit root. | $H_0: \beta_0 = \gamma = 0; H_A: \text{Not } H_0.$
Test statistic: ϕ_1 -statistic. |
| Case 3. | Estimated regression equation: | $\Delta y_t = \gamma y_{t-1} + \sum_{i=2}^p \beta_i \Delta y_{t+1-i} + \varepsilon_t$ |
| | Test for unit root: | $H_0: \gamma = 0; H_A: \gamma < 0$
Test statistic: τ -statistic. |

Empirical Results

Campbell and Perron (1991) report that the powers of unit root tests are higher over longer data span compared to those with shorter time span even if the latter has more observations due to more frequent sampling. Accordingly, empirical tests of unit roots are performed with data covering as long a period as available. Tests of unit roots are performed for US real income, World (vis-à-vis US, and Japan) real income, and Japanese real income over the period 1965Q1 to 1993Q2. For current account balance, relative government consumption, and multilateral real exchange rates of the US and Japan, tests are performed over the period 1970Q1 to 1993Q2. For each data series, starting with 12 lags, we reduce the number of lags in the equation on the basis of the significance of the last lag term and the joint significance of the last four lag terms. Finally, after choosing the lag-length the Ljung-Box test is used to verify the assumption of iid for the error term.

Results of Unit Roots Tests: The United States

The Augmented Dickey-Fuller test (in its different forms) has been performed to examine the potential for unit roots in different US data series, namely US and rest of the world GNP, multilateral effective real exchange rate, US real government consumption relative to rest of the world, and the current account balance. The test statistics used are the three forms of the τ -test, and the ϕ_i -test. The estimated values of the test statistics and their 95% critical values for the closest sample sizes are summarized in Table 4.1.

Table 4.1. Results of Unit Roots Tests: US Data

<u>The Augmented Dickey-Fuller test and the ϕ_i test</u>					
Series	No of Obs.	Statistic	$\tau_i(0.95)$	ϕ_i -Statistic	$\phi_i(0.95)$
US GNP	105	$\tau_\tau = -2.77$ $\tau_\mu = -0.88$	-3.45 -2.89	$\phi_3 = 4.47$	7.44
World Income	105	$\tau_\tau = -2.24$ $\tau_\mu = -2.40$	-3.45 -2.89	$\phi_3 = 4.76$	7.44
US Real Exchange Rate	89	$\tau_\mu = -1.68$ $\tau = -0.90$	-2.89 -1.95	$\phi_1 = 1.65$	4.71
Current Account Balance	89	$\tau_\mu = -1.81$ $\tau = -1.66$	-2.89 -1.95	$\phi_1 = 4.00$	4.71
Relative Govt. Consumption	89	$\tau_\tau = 0.73$ $\tau_\mu = -1.83$	-3.45 -2.89	$\phi_3 = 2.15$	4.71

US Real Income

The time series of US real income is plotted in Figure 4.1. It appears from the plot of the series that US real income may contain a linear time trend. Accordingly, first the lag-length is determined using the procedure outlined earlier¹. Once the lag-length is determined, the τ_τ -statistic is calculated by estimating equation (4.3). The sample value of the τ_τ -statistic is found to be -2.77. The 95% critical value of the τ_τ -statistic for a sample size of 100 is -3.45. The sample value the test statistic is calculated on the basis of 105 observations. Since the sample estimate of the test statistic is less (in absolute value) than

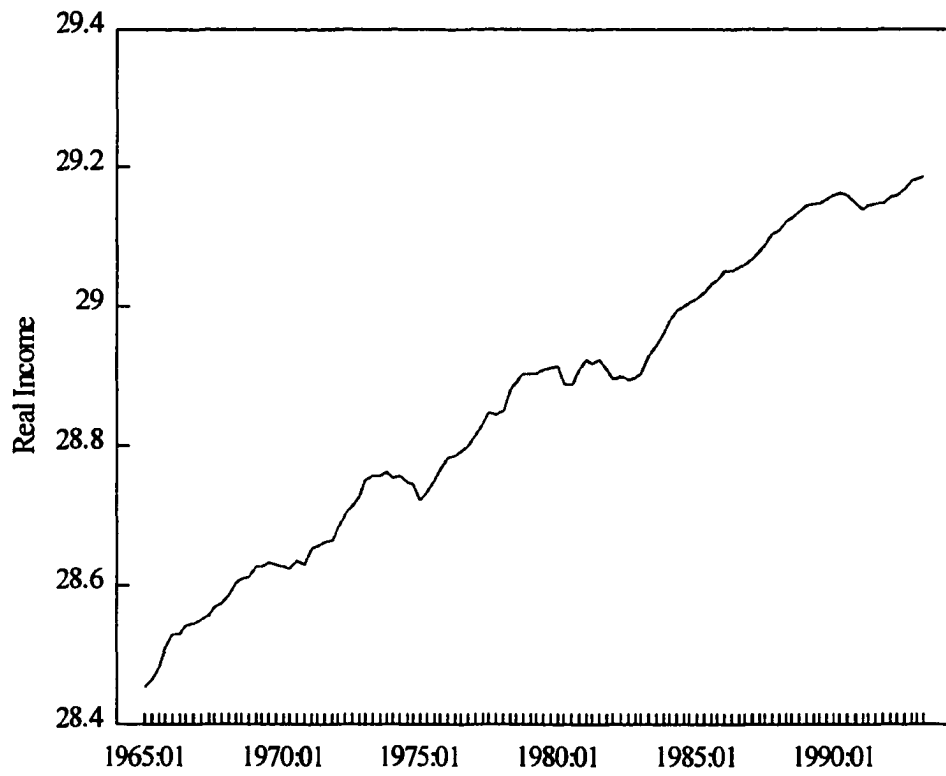


Figure 4.1 Real Income: The United States

the 95% critical value of the statistic, the null hypothesis of a unit root can not be rejected. In other words, the result of the τ_τ -test suggests that US real income series contains a unit root. Following the sequential testing strategy outlined earlier, the next logical step is to test for the appropriateness of including the linear time trend in equation (4.3). For that we use the ϕ_3 -statistic described in Dickey and Fuller (1981).

Having failed to reject the presence of a unit root using the τ_τ -statistic, test is performed for the following null and alternative hypotheses for equation (4.3):

$$H_0: \quad \beta_1 = \gamma = 0$$

$$H_A: \quad \text{Not } H_0 .$$

Sample estimate of the ϕ_3 -statistic is 4.47 which is well below the tabulated value of the test statistic at 95% level (which is 7.44 for 100 observations and 7.25 for 250 observations). Hence, the joint null of a unit root and no linear time trend can not be rejected.

On the basis of these results, equation (4.4) is estimated. Now the appropriate test for the presence of a unit root is the τ_μ -test. The estimated value of the test statistic is -0.88 which is well below the tabulated value of the test statistic at 95% significance level (which is -2.89 for 100 observations). Given the data series for real income of the US, it is reasonable to assume that the data generating process includes a drift term. On that assumption the τ_μ -statistic is considered as the appropriate test statistic. Given the result of the τ_μ -test above, it may be concluded that the real income series of the US contains a unit root. Test of stationarity of ΔY_t convincingly rejects the presence of any additional

unit root in Y_t . Hence, it concluded that there is one unit root in US real income.

World Income vis-à-vis the US

Plot of world (vis-à-vis the US) real income is presented in Figure 4.2. The plot suggests that there may be a linear time trend in world income. So, after determining the appropriate lag-length the τ_t -statistic is calculated. The sample estimate of the τ_t -statistic is -2.24 which is less (in absolute value) than the 95% critical value of the test statistic (which is -3.45 for 100 observations, and -3.43 for 250 observations). Hence, the null of a unit root can not be rejected. The test of the joint null of a unit root and no linear time trend (ϕ_3 -test) yields a sample estimate of 4.76 which is less than the 95% critical value of the ϕ_3 -statistic (which is 7.44 for 100 observations and 7.25 for 250 observations). Therefore, the joint null of a unit root and no linear time trend in world income (vis-à-vis the US) can not be rejected.

Following our sequential testing procedure, given above results, regression equation is re-estimated without any time trend, and the τ_μ -statistic is calculated. The estimated value of the test statistic is -2.40 which is below (in absolute terms) the 95% critical value (which is -2.89 for 100 observations and -2.88 for 250 observations). As in the case of US real income, it is reasonable to assume that the true data generating process contains a drift term. In that case the τ_μ -statistic is the appropriate tool to test for the presence of a unit root. Hence, it is concluded that world real income (vis-à-vis the US) contains a unit root. Formal test convincingly rejects any unit root in the first difference of the data series.

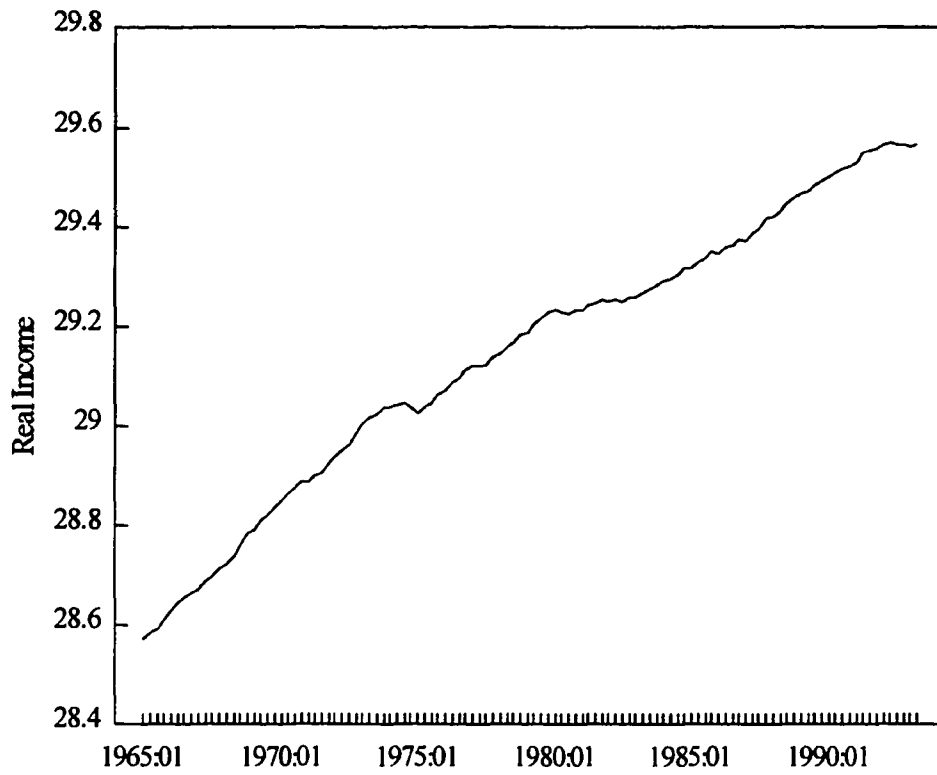


Figure 4.2 Real Income: Rest of the World vis-à-vis the U.S.

Multilateral Real Exchange Rate

Plot of US multilateral effective real exchange rate is shown in Figure 4.3. The graph does not suggest any linear time trend in the data. So, the regression equation is estimated with an intercept term and no linear time trend. Now the appropriate ADF test statistic is the τ_μ -statistic. The sample estimate of the test statistic is -1.68 which is well below (in absolute terms) the 95% critical value of the test statistic. Hence, the null hypothesis of a unit root in US multilateral real exchange rate can not be rejected.

Test of the significance of the intercept term in the estimated equation is performed with the ϕ_1 -test described in Dickey and Fuller (1981). The estimated value of the ϕ_1 -statistic is 1.65 which is lower than the 95% critical value (which is 4.71 for 100 observations and the sample size for US real exchange rate is 89). The ϕ_1 -test, therefore, suggests that in the presence of a unit root the intercept term is not significant.

On the basis of these results, equation (4.5) is estimated and the τ -statistic is calculated. The sample value of the τ -statistic is found to be -0.90 which is well below (in absolute terms) the 95% critical value (-1.95 for 100 observations). Hence, it is concluded that US multilateral real exchange rate contains a unit root. The potential for additional unit root is tested by performing the ADF test on the first difference of the real exchange rate. Results of that test decisively rejects the presence of any unit root in the first difference of the data series. Hence, our conclusion is that there is only one unit root in US multilateral effective real exchange rate data.

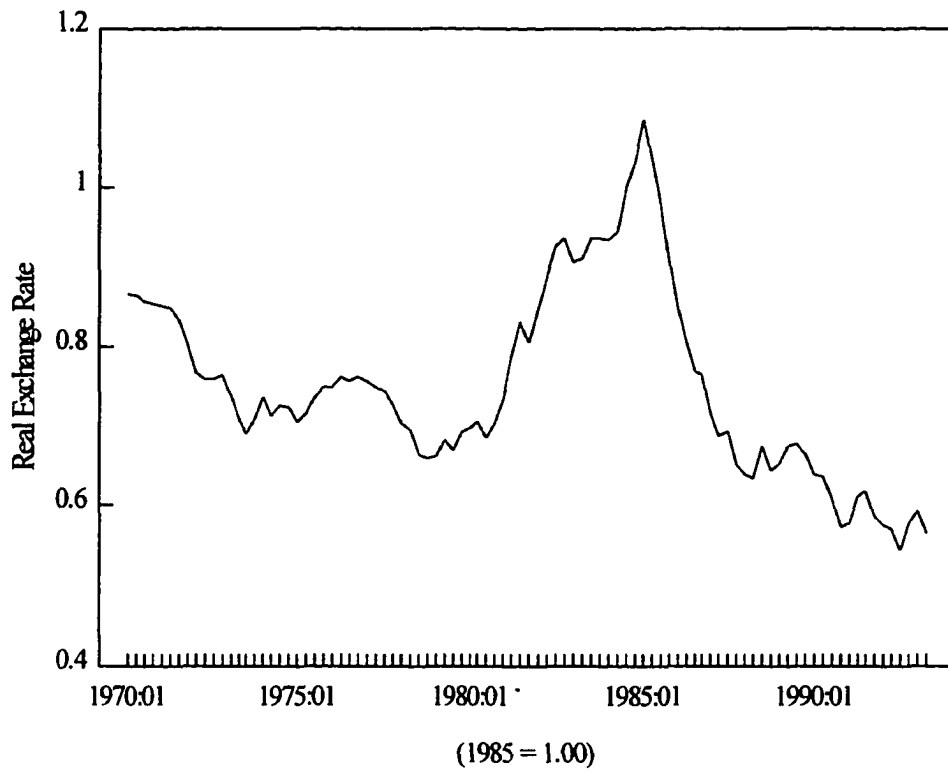


Figure 4.3 Multilateral Effective Real Exchange Rate: The United States

Current Account Balance

Graph of US current account balance is presented in Figure 4.4. In the graph, we do not see any evidence suggesting that US current account series contains a linear time trend. In this case, therefore, the appropriate test statistic is the τ_μ -test. Accordingly, first the regression equation (4.4) is estimated and τ_μ -statistic is calculated. The sample estimate of the τ_μ -statistic is -1.81. The 95% critical value of this test statistic for 100 observations is -2.89 and the sample size for US current account balance is 89. This suggests that there is a unit root in US current account balance.

Next we test for the significance of the intercept term in the presence of a unit root in the data. Test of significance of the intercept term in equation (4.4) in the presence of a unit root (the ϕ_1 -statistic) yields a sample estimate of 4.00 which is below its 95% critical value. So, the joint null of a unit root and no intercept term in equation (4.4) can not be rejected for US current account balance.

On the basis of the above test results, equation (4.5) is estimated and τ -statistic is calculated. The sample estimate of the τ -statistic is -1.66 which is below (in absolute terms) the 95% critical value of the test statistic. This result leads to the conclusion that there is a unit root in US current account balance series. Test for the presence of a unit root is performed on the first difference of the series. ADF test on the first difference of the series reveals no evidence of unit roots. Hence, we conclude that US current account series contains a unit root.

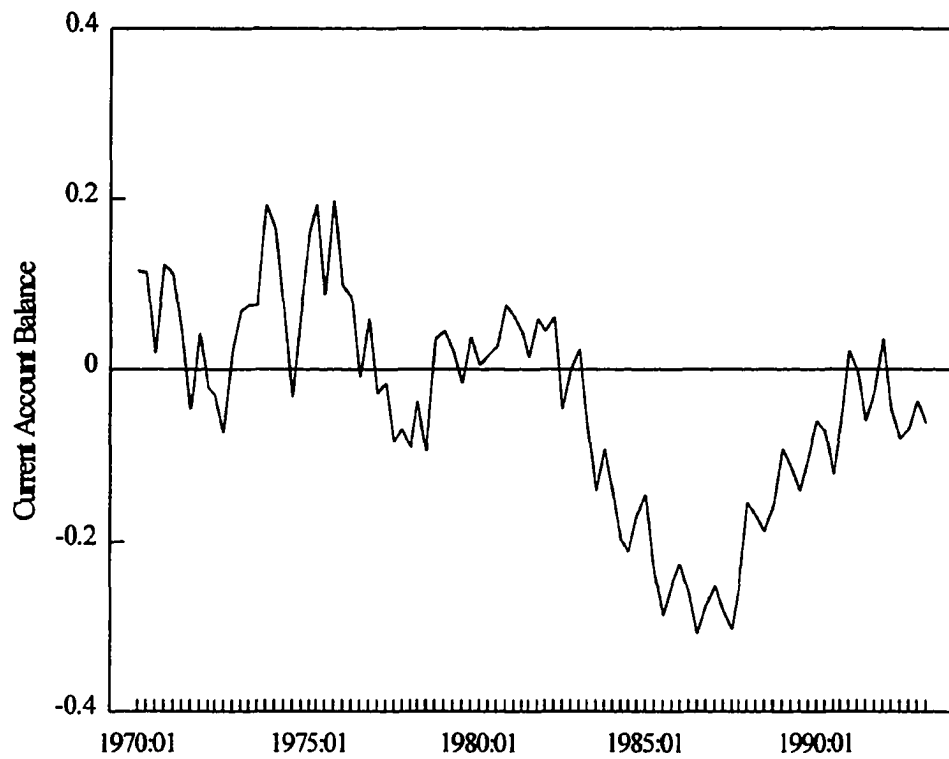


Figure 4.4 Current Account Balance: The United States

US Relative Government Consumption

Figure 4.5 represents the U.S. government consumption relative to rest of the world . The plot shows a general upward trend. This suggests that we should start with the τ_τ -test . The sample estimate of τ_τ -statistic is 0.73 which is below its 95% critical value (in absolute terms). This implies that the null of a unit root can not be rejected at conventional significance level. The significance of a linear time trend is tested with the ϕ_3 -statistic. When the ϕ_3 -statistic is calculated, the sample estimate is found to be 2.15 which is less than its 95% critical value. This implies that there is no (deterministic) linear time trend in the series. Given this result, our next logical step is to perform the τ_μ -test. When the τ_μ -statistic is estimated the sample value of the statistic is found to be - 1.83 which is less (in absolute value) than its 95% critical value. Hence, the null of a unit root can not be rejected at conventional significance level. Inspection of the graph suggests that the drift term should be included in performing the ADF test. Therefore, the result of the τ_μ -test is accepted and we conclude that US relative government consumption contains a unit root. As with other variable, test is performed on the first difference of the data series to investigate the potential for additional unit roots in the series. ADF test on the first difference of the series strongly rejects presence of additional unit root. All these results lead to the conclusion that there is only one unit root in US government consumption relative to rest of the world (vis-à-vis the US).

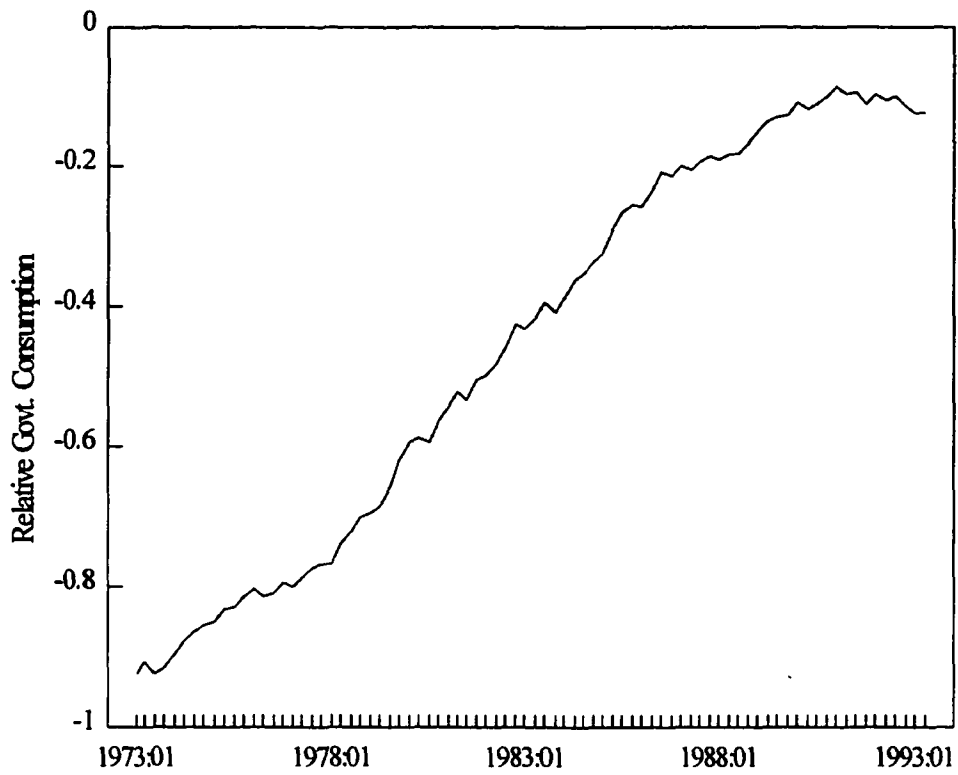


Figure 4.5 Relative Government Consumption: The United States

Results of Unit Roots Tests: Japan

Now we discuss the results of ADF test for Japanese data series. The procedure adopted is exactly the same as the one used for of U.S. data. The results of the Augmented Dickey-Fuller test for each of the variables are summarized in Table 4.2.

Real Income

Real income of Japan is plotted in Figure 4.6. The series clearly shows an upward trend suggesting the potential for a linear time trend. Following the same strategy as in the case of US data series, first the τ_τ -statistic is calculated for Japanese real income series.

Table 4.2. Results of Unit Roots Tests: Japan

<u>Augmented Dickey-Fuller and ϕ_i-test</u>					
Series	No of Obs	Statistic	$\tau_i(0.95)$	ϕ_i -Statistic	$\phi_i(0.95)$
Japanese GNP	105	$\tau_\tau = -1.85$ $\tau_\mu = -2.50$	-3.45 -2.89	$\phi_3 = 4.01$	7.44
World Income	105	$\tau_\tau = -1.84$ $\tau_\mu = -1.45$	-3.45 -2.89	$\phi_3 = 2.50$	7.44
Real Exchange Rate	89	$\tau_\mu = -1.58$ $\tau = 0.67$	-2.89 1.66	$\phi_1 = 1.67$	4.71
Current Account Balance	89	$\tau_\mu = -1.93$ $\tau = -1.11$	-2.89 -1.66	$\phi_1 = 1.89$	4.71
Relative Govt. Consumption	89	$\tau_\mu = -0.81$ $\tau = -0.50$	-2.89 1.66	$\phi_1 = 0.44$	4.71

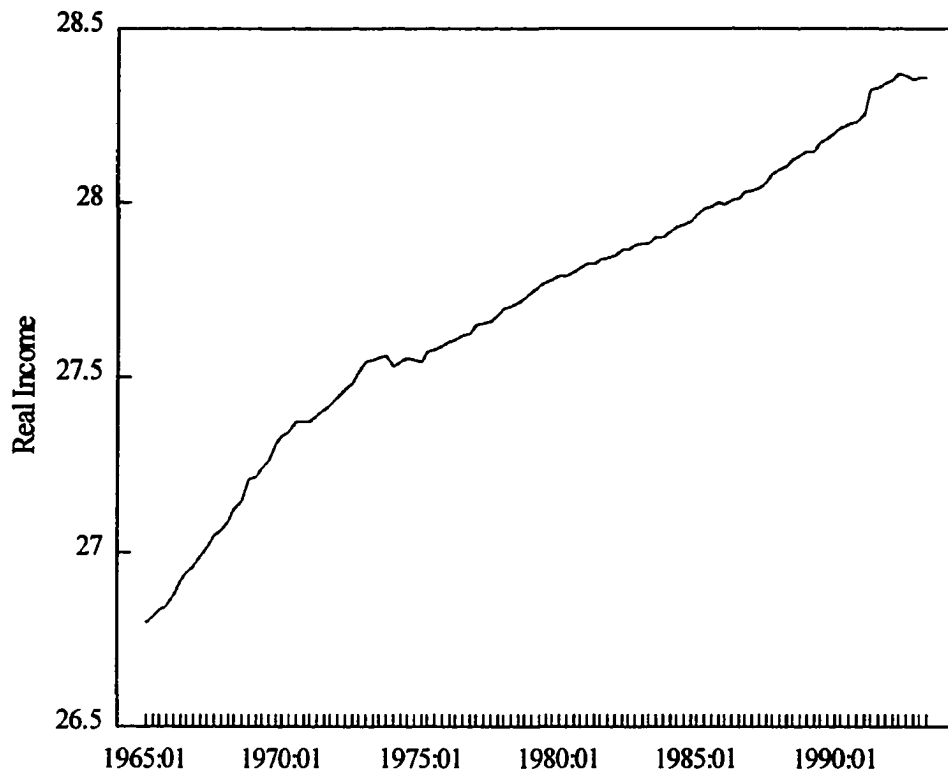


Figure 4.6 Real Income: Japan

From Table 4.2, we can see that the estimated value of the τ_τ -statistic is -1.85 which is well below (in absolute value) the 95% critical value of the test statistic. Therefore, on the basis of the τ_τ -test, the null hypothesis of a unit root can not be rejected at 5% significance level.

Having failed to reject the null hypothesis of a unit root on the basis of the τ_τ -test , we now test for the joint null of a unit root and no linear time trend in Japanese real income series. Again from Table 4.2, we see that the test for the joint null of a unit root and no linear time trend yields a sample value of the ϕ_3 -statistic equal to 4.01 which is less than the 95% critical value of the test statistic. Therefore, the above joint null can not be rejected.

On the basis of the above finding, equation (4.4) is estimated and the τ_μ -statistic is calculated. The estimated value of the τ_μ -statistic is -2.50 which is lower (in absolute value) than the 95% critical value of the test statistic. Again, it is reasonable to assume that real income data series contains a drift and the τ_μ -statistic is the appropriate test for the presence of a unit root. On the basis of the result of the τ_μ -test it is concluded that the real income series of Japan contains a unit root. ADF test on the first difference of Japanese GNP does not suggest any unit root in the first difference of the series. So, our conclusion is that the GNP series of Japan contains one unit root.

World Real Income vis-à-vis Japan

World real income (vis-à-vis Japan) is plotted in Figure 4.7. Again, the series shows an upward trend suggesting a potential linear time trend in the data. Therefore, first

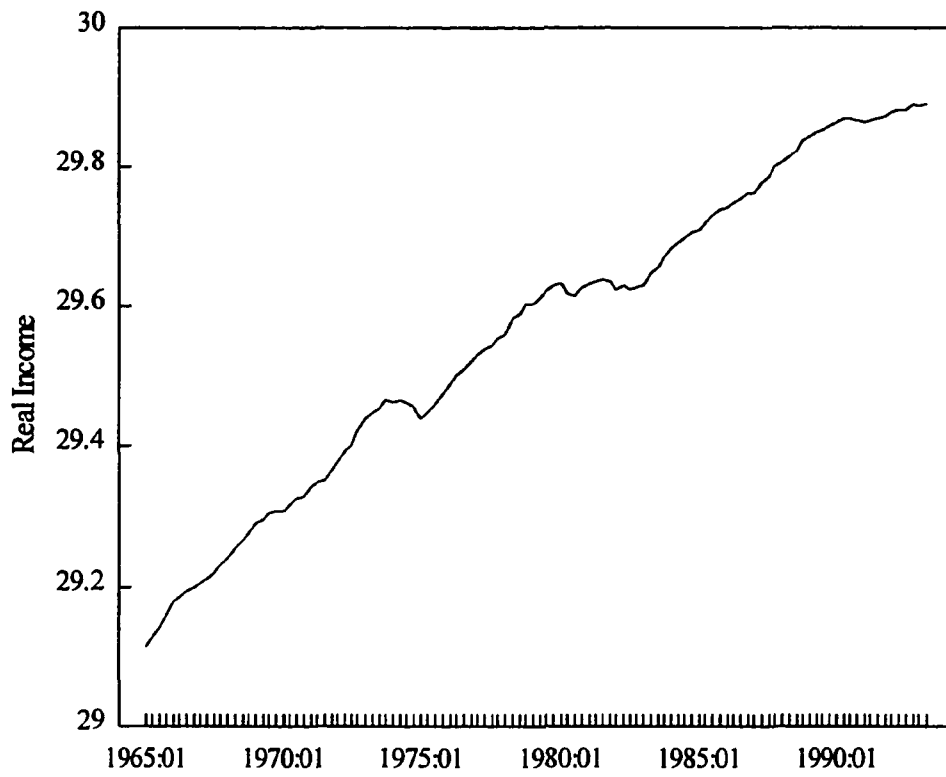


Figure 4.7 Real Income: Rest of the World vis-à-vis Japan

the τ_t -statistic is estimated. From Table 4.2 we can see that the sample estimate of the τ_t -test is found to be -1.84 which is below the 95% critical value of the test statistic. This suggests presence of a unit root in world income (vis-à-vis Japan). We then test the joint null of a unit root and no linear trend in world income (vis-à-vis Japan). Test of the joint null of a unit root and no linear time trend using the ϕ_3 -test yields an estimated value of the test statistic equal to 2.50 which is well below the 95% critical value of the test statistic (which is 7.44 for sample size 100). Hence, the joint null of a unit root and no linear time trend can not be rejected for world income (vis-à-vis Japan).

When equation (4.4) is estimated and the τ_μ -statistic is calculated, the sample estimate of the test statistic is found to be -1.45. Since this value is smaller (in absolute terms) than the 95% tabulated value of the test statistic, the null hypothesis of a unit root in world income can not be rejected. It is reasonable to assume that true data generating process contains a drift term, and therefore the τ_μ -statistic is the appropriate test statistic. On the basis of the outcome of the τ_μ -test, it may be concluded that world income (vis-à-vis Japan) contains a unit root. ADF test on the first difference of the series does not suggest any unit root in the differenced data.

Multilateral Real Exchange Rate

The plot of Japanese multilateral effective real exchange rate is presented in Figure 4.8. The plot does not suggest any time trend in the data series. So, equation (4.4) is estimated and the τ_μ -statistic is calculated. The estimated value of the test statistic is -1.58 and this is lower (in absolute value) than the 95% critical value of the test statistic.

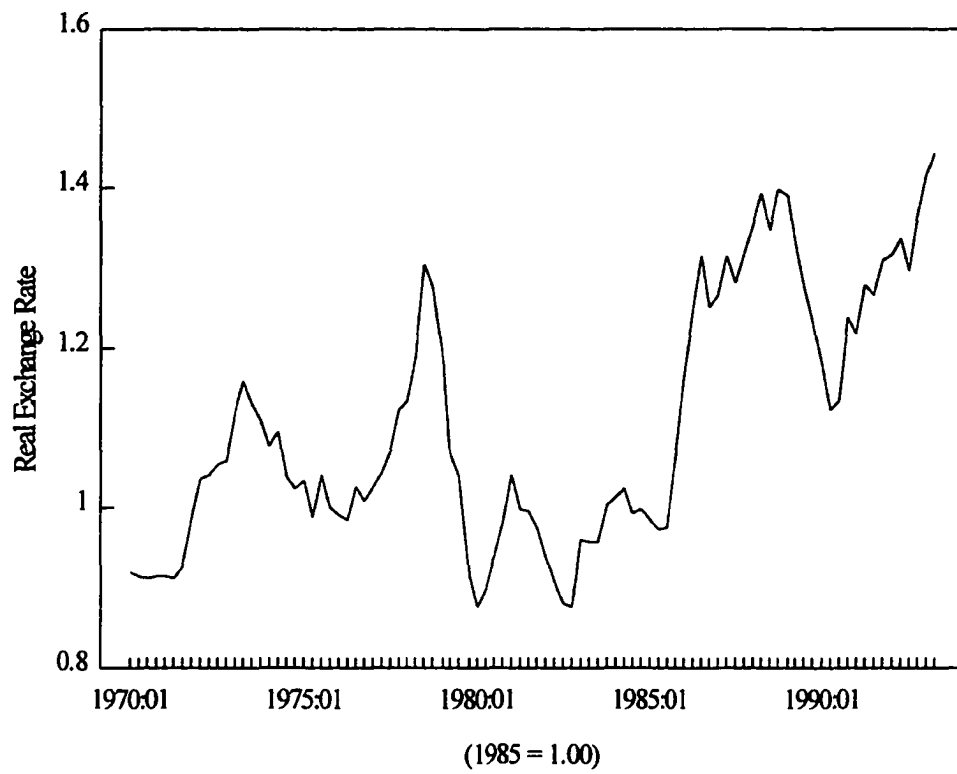


Figure 4.8 Multilateral Effective Real Exchange Rate: Japan

So the τ_μ -test fails to reject the null hypothesis of a unit root in Japanese multilateral real exchange rate. Having failed to reject the null of a unit root on the basis of τ_μ -test, we then test for the joint null of a unit root and no intercept term in equation (4.4). Test of this joint null hypothesis is the ϕ_1 -test. Sample estimate of the ϕ_1 -statistic is 1.67 which is lower than the 95% critical value of the test statistic (which is 4.71 for 100 observations). Therefore, the joint null of a unit root and no intercept term in equation (4.4) can not be rejected at 5% significance level.

Estimation of equation (4.5) and the τ -statistic yields a sample estimate of the statistic equal to 0.67. This is well below (in absolute terms) the 95% critical value of the τ -statistic (-1.95 for 100 observations). Therefore, the null hypothesis of a unit root can not be rejected at conventional significance level. Hence, it is concluded that the Japanese multilateral real exchange rate series contains a unit root. The ADF test on the first difference of the series clearly rejects the null of a unit root. Hence, the above test results leads us to the conclusion that there is only one unit root in the multilateral effective real exchange rate series of Japan.

Current Account Balance

Figure 4.9 shows the plot of the current account balance of Japan. The graph suggests no linear time trend in the data. This implies that the appropriate test statistic is the τ_μ -statistic. Accordingly, equation (4.4) is estimated and the τ_μ -statistic is calculated. The sample estimate of the τ_μ -statistic is -1.93 which is well below the 95% tabulated

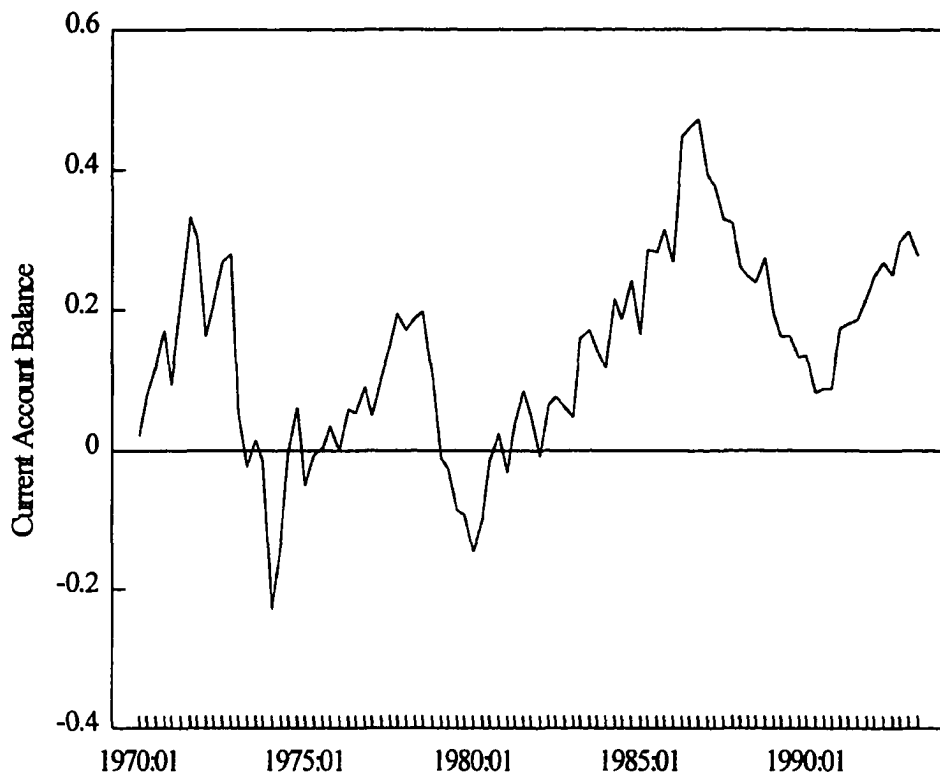


Figure 4.9 Current Account Balance: Japan

value of the test statistic. Therefore, on the basis of the τ_μ -test result, the null hypothesis of a unit root can not be rejected at 5% significance level

Since the τ_μ -test fails to reject the null of a unit root in the data series, we now test for the joint null of a unit root and no intercept term in equation (4.4) using the ϕ_1 -statistic. As can be seen from Table 4.2, the sample estimate of the ϕ_1 -statistic is 1.89 which is lower than the 95% critical value (which is 4.71 for 100 observations). Hence, the joint null of a unit root and no intercept term in equation (4.4) can not be rejected at 5% significance level.

On the basis of the above findings, regression equation (4.5) is estimated and the τ -statistic is calculated. The estimated value of the test statistic is -1.11 which lower (in absolute terms) than the 95% critical value of the test statistic (which is 1.95). Hence it is concluded that the current account balance series of Japan is characterized by the presence of a unit root. Further, ADF test on the first difference of the series does not show any evidence of unit roots in the differenced data. This suggests that there is only one unit root in the current account balance series of Japan.

Relative Government Consumption

Government consumption of Japan relative to rest of the world is shown in Figure 4.10. The graph shows no evidence of a linear time trend . Accordingly, first the τ_μ -statistic is calculated. The sample estimate of the statistic is found to be -0.81 which is lower than the 95% critical value (in absolute terms) of the test statistic (which is -2.89 for 100 observations). This implies that the null of a unit root in the series can not be

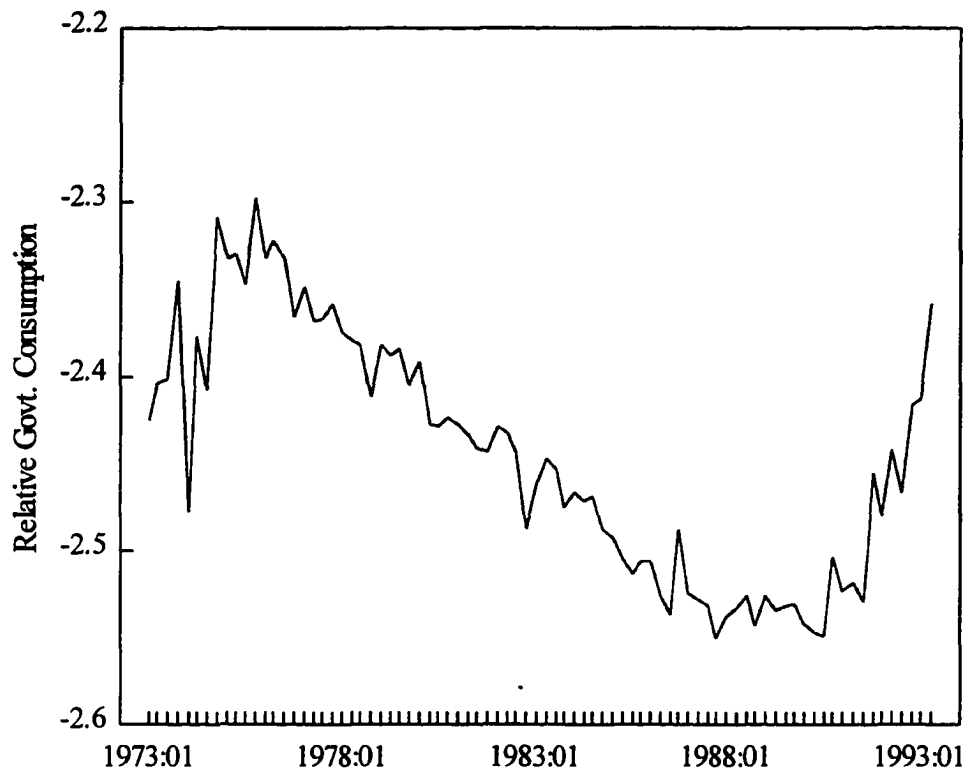


Figure 4.10 Relative Government Consumption: Japan

rejected at conventional significance level. Then the appropriateness of including an intercept term in the regression equation is tested using the ϕ_1 -statistic. The estimated ϕ_1 -statistic is 0.44 which is lower than the 95% critical value of the test statistic. This suggests that the intercept term should not be included in the regression equation. So, the equation is estimated without an intercept term, and the τ -test is performed. The sample value of the test statistic is found to be -0.50 which is clearly lower (in absolute value) than the 95% critical value of the τ -statistic. This suggests that the null of a unit root in the data series can not be rejected. The ADF test on the first difference of the series convincingly rejects the null of a unit root. Hence we may conclude that there is evidence in favor of only one unit root in the level of the data series.

It is evident from the test results presented above that real income, real exchange rate, current account balance, and relative government consumption data all contain a unit root. The implication of this finding is that all of these variables are affected by shocks whose effects on the respective variables do not fade away over time. These are essentially what we call permanent shocks. Our focus in the following chapters is on the methods to identify the permanent and transitory components in income and real exchange rate, and then use the decomposed series to test the hypotheses generated by the intertemporal models of current account determination.

Endnotes

1. Equations (4.3) through (4.5) were estimated with 4 lags. Plots of the residuals of equations (4.3) through (4.5) were tested visually for evidence of structural breaks. Such inspection of the residuals did not reveal signs of structural change. Autocorrelation and partial autocorrelation functions of the residuals did not show significant serial correlation. Formally, the Ljung-Box statistics were calculated for each residual for lags up to 24. The test statistics clearly showed that the autocorrelations at various lags were all statistically insignificant. Thus, diagnostic checking of the residuals showed no evidence of structural breaks or significant violation of the iid assumption.

CHAPTER 5. TRANSITORY AND PERMANENT COMPONENTS IN INCOME AND REAL EXCHANGE RATE

Transitory and Permanent Components in Time Series

It is evident from chapter 4 that real income (of the US, rest of world vis-à-vis the US, Japan, and rest of world vis-à-vis Japan) and multilateral effective real exchange rates of the US and Japan are characterized by the presence of a unit root. In addition to its relevance from the statistical point of view (estimation and inference), presence of a unit root in the data series has important conceptual implications. Specifically, if there is a unit root in a series Y_t then exogenous shocks to the variable Y_t have permanent effects in the sense that the effects of exogenous shocks to Y_t do not fade away even in the long-run. To illustrate, let us consider the simple case where the stochastic variable Y_t may be represented as a random walk plus a noise process:

$$y_t = y_{t-1} + \varepsilon_t + \upsilon_t - \upsilon_{t-1},$$

where ε_t and υ_t are random shocks. In the above representation, the effect of ε_t on Y_t is permanent while that of υ_t is transitory in nature. This can be seen by writing the general solution for Y_t , given the initial condition that at $t = 0$ the values of Y_t and υ_t are Y_0 and υ_0 respectively

$$Y_t = Y_0 - \upsilon_0 + \sum_{i=1}^t \varepsilon_i + \upsilon_t.$$

It is evident from the general solution for Y_t that the shock υ_t affects Y_t only

contemporaneously with no effect on Y_{t+s} ($s > 0$). Similarly, $v_{t;j}$ ($j > 0$) has no effect on Y_t . On the other hand, a given realization of ε_t affects the current value of Y_t as well as all subsequent values Y_{t+s} ($s > 0$). More generally, let us consider the following time series model:

$$Y_t = \alpha_0 + \sum_{i=1}^k \alpha_i Y_{t-i} + u_t,$$

where

$$u_t = \sum_{i=0}^p \beta_i \varepsilon_{t-i} + \sum_{i=0}^q \gamma_i \eta_{t-i}.$$

So, Y_t may be expressed as $A(L)Y_t = B(L)\varepsilon_t + C(L)\eta_t$, where $A(L)$ and $B(L)$ are appropriate polynomials in the lag operator L . Now Y_t may be solved as:

$$Y_t = \frac{B(L)}{A(L)} \varepsilon_t + \frac{C(L)}{A(L)} \eta_t = \frac{\varepsilon_t}{D(L)} + \frac{\eta_t}{E(L)}.$$

If the roots of $D(L)$ and $E(L)$ are all within the unit circle (in modulus) then Y_t may be expressed as an infinite sum of current and past realizations of ε_t and η_t , and Y_t is covariance stationary. However, if only one of the polynomials, for example $D(L)$, has one root equal to unity then Y_t may be expressed as:

$$Y_t = \frac{\varepsilon_t}{(1-L)F(L)} + \frac{\eta_t}{E(L)} = \sum_{i=1}^{\infty} \varepsilon_{t-i} + \frac{\eta_t}{E(L)}.$$

Since $D(L)$ has only one unit root, all roots of $F(L)$ are within the unit circle (in modulus).

In this case, the effect of η_t may persist more than one period. Given enough time, the

effect of η_t shock on Y_t will fade away, and in the long-run there will not remain effect of this shock on Y_t . On the other hand, it is evident that the effect of an ε_t shock will persist forever. In this sense, an ε_t shock will have a permanent effect on Y_t .

As has been discussed in chapter 4, income and real exchange rate series are characterized by the presence of a unit root. This implies that some shocks to these variables have permanent effects. In addition to the permanent shocks, there may be transitory shocks to these variables. The remainder of this chapter focuses on how one may attempt to identify and separate the transitory and permanent shocks, and the movements in the observed variables associated with the two types of shocks. Specifically, two methods of identifying the transitory and permanent components of the variables are discussed. The first is an univariate suggested by Beveridge and Nelson (1981). The other is a bivariate method proposed by Blanchard and Quah (1989) who described and implemented the method to identify transitory and permanent components in US real income.

Decomposition of Time Series into Transitory and Permanent Components

Despite their elegance, the intertemporal models of current account determination have not been tested empirically due to our inability to separate the transitory and permanent components in economic data series. However, some recent developments in time series techniques provide ways to implement such decomposition of real world data. This section provides a brief discussion of the two methods used in this study. The first method is a univariate technique suggested by Beveridge and Nelson (1981). The other

method is due to Blanchard and Quah (1989). This method uses a bivariate vector autoregression (VAR) of two variables and imposes some identifying restrictions on the coefficients of the VAR representation to identify and estimate the transitory and the permanent components in a variable containing a unit root.

Univariate Decomposition: The Beveridge-Nelson Method

Beveridge and Nelson (1981) developed an elegant method of decomposing a time series with a unit root into a permanent component (a stochastic trend component) and a transitory component (a cyclical component that is stationary in nature). Their idea and decomposition method is reviewed below.

Let z_t be a variable with one unit root, i.e., $z_t \sim I(1)$, and therefore $w_t = z_t - z_{t-1}$ is stationary (i.e., $w_t \sim I(0)$). Since w_t is covariance stationary, the Wold representation theorem may be used to write:

$$w_t = \mu + \sum_{i=0}^{\infty} \lambda_i \varepsilon_{t-i}; \quad \lambda_0 = 1,$$

where ε_t are iid with zero mean and constant variance. The expectation of z_{t+k} conditional on data for z through time t , denoted by $\hat{z}(k)$, is given by:

$$\hat{z}(k) = E(z_{t+k} | z_t, z_{t-1}, \dots),$$

which is equivalent to:

$$\hat{z}(k) = z_t + E(w_{t+1} + w_{t+2} + \dots + w_{t+k} | w_t, w_{t-1}, \dots).$$

Therefore, we have:

$$\hat{z}(k) = z_t + \hat{w}_t(1) + \hat{w}_t(2) + \dots + \hat{w}_t(k). \quad (5.1)$$

Now, $w_{t+i} = \mu + \lambda_0 \varepsilon_{t+i} + \lambda_1 \varepsilon_{t+i-1} + \lambda_2 \varepsilon_{t+i-2} + \dots + \lambda_i \varepsilon_t + \lambda_{i+1} \varepsilon_{t-1} + \lambda_{i+2} \varepsilon_{t-2} + \dots$,

Therefore, we may write:

$$\hat{w}_t(i) = \mu + \lambda_i \varepsilon_t + \lambda_{i+1} \varepsilon_{t-1} + \lambda_{i+2} \varepsilon_{t-2} + \dots$$

This gives:

$$\hat{w}_t(i) = \mu + \sum_{j=0}^{\infty} \lambda_{i+j} \varepsilon_{t-j}. \quad (5.2)$$

Substituting (5.2) in (5.1) yields:

$$\hat{z}_t(k) = z_t + k\mu + \left(\sum_{i=1}^k \lambda_i \right) \varepsilon_t + \left(\sum_{i=2}^{k+1} \lambda_i \right) \varepsilon_{t-1} + \dots$$

Taking a long forecast horizon, one may write:

$$\hat{z}_t(k) \cong z_t + k\mu + \left(\sum_{i=1}^{\infty} \lambda_i \right) \varepsilon_t + \left(\sum_{i=2}^{\infty} \lambda_i \right) \varepsilon_{t-1} + \dots \quad (5.3)$$

Thus, the forecast profile is asymptotic to a linear function of the forecast horizon k with slope equal to μ (the rate of drift in the series) and a *level* (algebraically the intercept) which itself is a stochastic process. Beveridge and Nelson (1981) interpret this level as the permanent or (stochastic) trend component. So, the permanent component in z_t is given by:

$$\bar{z}_t = z_t + \left(\sum_{i=1}^{\infty} \lambda_i \right) \varepsilon_t + \left(\sum_{i=2}^{\infty} \lambda_i \right) \varepsilon_{t-1} + \dots$$

This permanent component of z_t is a random walk process with drift μ (as has been shown in Beveridge and Nelson (1981)), and may be interpreted as the current observed value of z plus all forecastable future changes in the series beyond the mean rate

of drift. In other words, it is the value the series would have if it remained in its long-run path in the current time period.

Empirical implementation of their decomposition method required Beveridge and Nelson (1981) to confine their attention to linear processes of rational form:

$$w_t = \mu + \frac{\theta(L)}{\phi(L)} \varepsilon_t,$$

where $\theta(L)$ and $\phi(L)$ are polynomials of order p and q in the lag operator L . Within this class of processes, we may write:

$$\phi(L)w_t = \phi(L)\mu + \theta(L)\varepsilon_t.$$

The above equation may be estimated to obtain the parameters (θ , ϕ , and μ) and the innovations ε_t . Beveridge and Nelson (1981) suggest truncating the AR and MA components at some suitably large values of p and q . Obviously the above procedure involves rather heavy computational burden as equation (5.3) needs to be evaluated numerically for each period in the sample. This prompted Cuddington and Winters (1987) to suggest an alternative procedure to obtain the decomposition. Their method involves the tedious Beveridge and Nelson computation for one period only to obtain the level benchmark. Once the benchmark level is estimated, decomposition can be obtained rather easily for other time periods of the sample. Miller (1988) follows the Cuddington and Winters (1987) method to derive an alternative expression for $\bar{z}(t)$, the stochastic trend component, that does not involve the complex Beveridge and Nelson computation even for obtaining the benchmark level. When appropriately corrected, Miller's method

provides an exact formula if w_t is a pure AR(p) process. If, on the other hand, w_t contains a moving average (MA) component then this method too requires truncation of an infinite sum. Newbold (1990) shows that these computational complexities are simply unnecessary. He suggests an alternative computational algorithm which is described below.

Subtracting the mean from the ARMA(p,q) process w_t , let us define $Y_t = w_t - \mu$, so that $\hat{y}_t(j) = \hat{w}_t(j) - \mu$, where $\hat{y}_t(j)$ is the j-period ahead forecast of Y_t at period t. Then the expression for \bar{z}_t becomes:

$$\bar{z}_t = z_t + \lim_{k \rightarrow \infty} \left[\sum_{j=1}^k \hat{w}_t(j) - k\mu \right].$$

Alternatively, we may write:

$$\bar{z}_t = z_t + \lim_{k \rightarrow \infty} \left[\sum_{j=1}^k \hat{y}_t(j) \right] = z_t + c_t.$$

Newbold (1990) then proves that c_t may be written as:

$$c_t = \sum_{j=1}^q \hat{y}_t(j) + (1 - \phi_1 - \dots - \phi_p)^{-1} \sum_{j=1}^p \sum_{i=j}^p \phi_i \hat{y}_t(q - j + 1),$$

where

$$\hat{y}_t(i) = y_{t+i} \text{ for } i \leq 0.$$

The method of Newbold (1990) has been is used here to obtain the univariate decomposition of the non-stationary variables into respective transitory and permanent components¹.

Bivariate Decomposition: The Blanchard-Quah Method

Blanchard and Quah (1989) describes a bivariate method of identifying permanent and transitory movements in real GNP. Using the information contained in the joint process followed by real income and unemployment rate, and by imposing restrictions on the coefficients of the bivariate moving average representation of the first differences of real GNP and (the level of) unemployment rate, they show how one may recover the unobserved transitory and permanent shocks from observed data. Unlike the Beveridge-Nelson (1981) method, this method does not restrict the permanent component to be random walk.

The issue of decomposition of an integrated sequence of random variable has been discussed in Quah (1992). He shows that if Y is integrated then, under certain conditions, it is possible to obtain a unique permanent and transitory decomposition of the integrated series Y such that $Y = Y_p + Y_T$ where $Y_p \sim I(1)$ is the permanent component and $Y_T \sim I(0)$ is the transitory component. Permanent in this context indicates only that disturbances to Y_p have long-run effects on Y (and not that the increments of Y_p are uncorrelated). Transitory means that the effects of shocks to this component disappear gradually over time. The essence of the theorem proving such a result may be summarized as follows: let Y be integrated and w be a given random sequence such that $z = [w, \Delta Y]'$ is jointly covariance stationary and linearly regular with spectral density matrix full rank at origin. Further, in Wold decomposition $z = C \cdot v$ each entry in C is $1/2$ -summable. Then it is possible to obtain a permanent and transitory decomposition of

Y if and only if ΔY is not Granger causally prior to w . If such a decomposition exists, then it is unique. This section presents a brief discussion of the Blanchard-Quah method of decomposing a nonstationary time series into transitory and permanent components. The discussion is presented in terms of identifying the transitory and permanent components of real income. The same methodology is applied to decompose the real exchange rates into transitory and permanent components.

Following Blanchard and Quah(1989), let us assume that variables x_t and Y_t are both characterized by the presence of a unit root. Let x_t denote real exchange rate and Y_t denote real income. In vector notation let $z_t = [\Delta x_t, \Delta Y_t]'$, where $\Delta = (1-L)$ denotes the first difference of the variable in question. It assumed that there are two types of disturbances affecting the system: the first type of shock has no long-run effect on real income. That is, effects of this shock dies down gradually over time and, in the long-run it has no effect on Y_t . The second type of disturbance may have long-run effects on the variables. This is the shock that is reflected by the presence of a unit root in the GNP data. Let us denote the two type of shocks by $[\varepsilon_{\tau} \quad \varepsilon_{\text{pl}}]'$. The exact source of the two types of shocks are not important; all that is needed is that the two types of shocks have distinct effects in terms of how long they can affect the variables. Each of the two shocks may individually be a composite of more than one shock, but the dynamic effects of the individual component of ε_{jt} ($j = \tau, \text{p}$) are similar². Each shock is permitted to be serially correlated. However, they are uncorrelated at all leads and lags.

Under usual regularity conditions, both components of z_t can always be uniquely represented as an invertible distributed lag of serially uncorrelated disturbances. The vector autoregression (VAR) exists with square summable MA representation if the vector process z_t is stationary, and there is no cointegration between x_t and Y_t ³. We shall need the additional technical condition that the innovations in the bivariate Wold decomposition of the first difference of x_t and Y_t are linear combinations of the two underlying disturbances. The joint process followed by x_t and Y_t may be described as follows.

The assumptions about the joint process described by z_t imply that z_t follows a stationary process given by:

$$z_t = C(0)\varepsilon_t + C(1)\varepsilon_{t-1} + C(2)\varepsilon_{t-2} + \dots = \sum_{j=0}^{\infty} C(j)\varepsilon_{t-j}. \quad (5.4).$$

We assume that $\text{Var}(\varepsilon_t) = I_2$, where I is an identity matrix. This is simply a convenient normalization assumption. Let us refer to the process described above for z_t as *the working model*. Our goal is to identify the above *working model* from actual data. The assumption that ε_T has no long-run effect on Y_t imposes a restriction on the coefficients of the sequence of $C(t-j)$ matrices. This restriction may be written as :

$$C_{21}[L=1] = 0.$$

This long-run restriction will be used to identify the working model. The restriction that $C_{21}[L=1] = 0$ implies that not only is ΔY_t unaffected by ε_T but also that ε_T has no long-run effect on the level of Y_t itself⁴.

Since z_t is stationary, it has a Wold moving average representation given by:

$$z_t = v_t + A(1)v_{t-1} + A(2)v_{t-2} + \dots = \sum_{j=0}^{\infty} A(j)v_{t-j}, \quad (5.5)$$

where $A(0) = I_2$, and $\text{Var}(v_t) = \Sigma$, where Σ is a 2×2 matrix.

The moving average representation (5.5) is unique and can be obtained by first estimating and then inverting the vector autoregressive (VAR) representation of z_t . From the two representations of z_t above, we see that v_t , the vector of innovations, ε_t and the vector of original disturbances, are related as follows:

$$v_t = C(0)\varepsilon_t, \text{ and } C(j)\varepsilon_{t-j} = A(j)v_{t-j}.$$

Substituting for $v_t = C(0)\varepsilon_t$ in the right hand side of the second of the two relations above, we may write:

$$C(j) = A(j)C(0). \quad (5.6)$$

So if we can identify $C(0)$ then we can obtain $C(j)$ from our knowledge of $A(j)$ — which we can obtain by inverting the VAR representation of z_t . Also, knowledge of $C(0)$ allows us to obtain the original disturbance vector ε_t as $C^{-1}(0) \cdot v_t$. Once we identify $C(0)$, $C(j)$ and ε_t , we can obtain our transitory and permanent decomposition of Y_t from equation (5.4).

For identification, we use the following restrictions: from long-run effect of ε_T on Y_t we have the first restriction:

$$C_{21}[L = 1] = 0.$$

This is equivalent to: $\sum_{j=0}^{\infty} C_{21}(j) = 0$.

However, from equation (5.6) we have $C(j) = A(j)C(0)$. An equivalent way to express the above restriction is as follows:

$$\sum_{j=0}^{\infty} C_{21}(j) = \sum_{j=0}^{\infty} A_{21}(j)C_{11}(0) + \sum_{j=0}^{\infty} A_{22}(j)C_{21}(0) = 0. \quad (5.8)$$

Other restrictions come from the variance-covariance relationships. Recalling that $C(0)\varepsilon_t = v_t$ where $\text{Var}(\varepsilon_t) = I_2$ and $\text{Var}(v_t) = \Sigma$, we have the following restrictions from the variance-covariance matrix:

$$C(0)C'(0) = \Sigma,$$

where $\Sigma = \begin{bmatrix} \sigma_{11} & \sigma_{12} \\ \sigma_{21} & \sigma_{22} \end{bmatrix}$

This gives the following restrictions:

$$C_{11}^2(0) + C_{12}^2(0) = \sigma_{11}, \quad (5.9)$$

$$C_{11}(0)C_{21}(0) + C_{12}(0)C_{22}(0) = \sigma_{12}, \quad (5.10)$$

$$C_{21}^2(0) + C_{22}^2(0) = \sigma_{22}. \quad (5.11)$$

The long-run restriction (equation (5.8)) and the variance-covariance relationships (equations (5.9) through (5.11)) allow us to identify the $C(0)$ matrix.

The decomposition strategy may be summarized as follows: first we estimate a vector autoregressive representation of z_t as :

$$z_t = B(L)z_{t-1} + v_t, \quad (5.12)$$

where $B(L)$ is a matrix polynomial in the lag operator, L .

Then the estimated VAR model (equation (5.12)) is inverted to obtain the bivariate

moving average (BMAR) representation :

$$z_t = A(L)v_t, \quad (5.13)$$

where $A(L) = [I - B(L)L]^{-1}$.

Once the BMAR (equation (5.13)) is obtained, we use our restrictions to identify and estimate $C(0)$ and then use this matrix to obtain $C(j) = A(j) C(0)$ and $\varepsilon_t = C^{-1}(0) \cdot v_t$. Once we have our *working model* identified, the permanent and transitory decomposition is given by equation (5.4). In empirical implementation of this decomposition method, the bivariate VAR in the first differences of real income and real exchange rate is estimated with 4 lags. Formally, Sim's lag-length test is used in an attempt to choose the shortest acceptable lag for the VAR. Since we are using quarterly data, lag-length tests are performed on lags that are multiples of four. Specifically, starting with 12 lags in the VAR, first the joint significance of lags 9 through 12 is tested. If these four lags are found statistically insignificant then test is performed on the joint significance of lags 5 through 8. For both US and Japanese data, Sim's test suggests that lags 5 through 12 in the VAR are jointly insignificant. Accordingly, the decomposition is performed with 4 lags in the bivariate VAR in ΔY and ΔR .

For visual comparison, plots of the transitory components of different data series obtained under the two decomposition methods are presented in appendix D. There, only the transitory components are presented. The permanent component in a particular data series is obtained as the difference between the observed data and the transitory component in the respective series. Therefore, the plots of the transitory components of a

series under the two decomposition methods provide sufficient information about the difference between the two decomposition methods.

Endnotes

1. The decomposition is performed with the help of the statistical package RATS. Philip Meguire of North Carolina State University has kindly provided an upgraded version of the decomposition algorithm written for RATS. In practice, one fits an ARMA(p,q) model on the first difference of the I(1) variable using Box-Jenkins method. This allows the choice of appropriate values of p and q that are used in the Beveridge-Nelson decomposition via this algorithm.

2. See Blanchard and Quah (1989) on this issue.

3. If x_t and Y_t are each difference stationary, and a cointegrating relation exists between them, (i.e., x_t and Y_t are CI(1,1)) then a finite order VAR representation in the first difference between them does not exist. See Campbell (1987), and Lastrapes (1992) on this issue. In empirical implementation of the Blanchard-Quah method, tests are performed to examine if there exists any cointegrating relationship between real income and real exchange rate. The tests show that both for the US and Japan there is no cointegration between real income and real exchange rate. Huizinga (1987) reaches similar conclusion on the issue of cointegration between real income and real exchange rate.

4. We can show this as follows:

$$\Delta Y_t = (1 - L)Y_t = C_{21}(L)\varepsilon_{Tt} + C_{22}(L)\varepsilon_{pt}.$$

This gives:

$$Y_t = (1 - L)^{-1} C_{21}(L)\varepsilon_{Tt} + (1 - L)^{-1} C_{22}(L)\varepsilon_{pt} = \sum_{k=0}^{\infty} \sum_{j=0}^k C_{21}(j)\varepsilon_{T,t-j} + \sum_{k=0}^{\infty} \sum_{j=0}^k C_{22}(j)\varepsilon_{p,t-j}$$

Here $\sum_{j=0}^{\infty} C_{21}(j)$ is the effect of ε_{Tt} on Y_t after infinitely many time periods. Therefore, the first term in the equation above represents the effect of ε_{Tt} on Y_t after long enough time periods. The restriction that $\sum_{j=0}^{\infty} C_{21}(j) = 0$, therefore, is equivalent to saying that ε_{Tt} has no long-run effect on the level of Y_t .

CHAPTER 6. LONG RUN RELATIONSHIP: COINTEGRATION AND ERROR CORRECTION

Cointegration and Error Correction

In order to introduce the concept of cointegration and its relation to error correction models, it is useful to revisit the concept of integration in the context of time series data. A series x_t with no deterministic component which has a stationary and invertible autoregressive moving average (ARMA) representation after differencing d times, but which is not stationary after differencing $(d-1)$ times, is said to be integrated of order d , denoted $x_t \sim I(d)$. Informally, a series is said to be integrated if it accumulates some past effects; such a series is non-stationary because its future path depends upon all such past influences, and is not tied to some mean to which it must eventually return. A series that is integrated of order d needs to be differenced d times to achieve stationarity. However, a linear combination of a number of series, each integrated of order d , may have a lower order of integration than any one of the individual series. In this case, the variables are said to be cointegrated.

Formally, the concept of cointegration may be defined as follows: let x_t denote an n -dimensional vector such that the components of the x_t are integrated of order d . The elements of x_t are said to be cointegrated of order $(d - b)$, denoted $x_t \sim CI(d, b)$ if:

(i) $x_t \sim I(d)$; and (ii) there exists a non-zero vector α such that $\alpha' x_t \sim I(d - b)$, $d \geq b > 0$.

The vector α is called the cointegrating vector.

The notion of cointegration allows us to describe the existence of potential equilibrium or stationary relation between two or more time series, each of which is individually non-stationary. The concept of equilibrium in economics refers to the idea that the variables hypothesized to be linked by some theoretical economic relationship should not diverge from each other in the long-run. Such variables may drift apart in the short-run, or because of seasonal effects. However, if there exists any equilibrium relation among them then these variables can not diverge indefinitely and without bounds. In other words, the divergence from a stable equilibrium state has to be statistically bounded and, at some point, diminishing over time. In this context, cointegration may be viewed as a statistical expression of the nature of the relationship among the variables that are tied together through a long-run equilibrium relationship.

To illustrate, let us consider the case where two series $\{x_t\}$ and $\{y_t\}$ are each integrated of order 1 and evolve according to the following data generating process¹:

$$x_t + \beta y_t = u_t, \quad u_t = u_{t-1} + \varepsilon_{1t}, \quad (6.1)$$

$$x_t + \alpha y_t = e_t, \quad e_t = \rho e_{t-1} + \varepsilon_{2t} \text{ with } |\rho| < 1, \quad (6.2)$$

It is assumed that $[\varepsilon_{1t}, \varepsilon_{2t}]'$ identically and independently distributed (iid) as bivariate normal with $E(\varepsilon_{1t}) = E(\varepsilon_{2t}) = 0$, $\text{var}(\varepsilon_{1t}) = \sigma_{11}$, $\text{var}(\varepsilon_{2t}) = \sigma_{22}$, and $\text{cov}(\varepsilon_{1t}, \varepsilon_{2t}) = \sigma_{12}$.

Assuming $\alpha \neq \beta$, the above system may be solved as:

$$x_t = \alpha(\alpha - \beta)^{-1} u_t - \beta(\alpha - \beta)^{-1} e_t, \quad (6.3)$$

$$y_t = -(\alpha - \beta)^{-1} u_t + (\alpha - \beta)^{-1} e_t. \quad (6.4)$$

Since $\{u_t\}$ is a random walk process, and $\{x_t\}$ and $\{y_t\}$ are linear functions of $\{u_t\}$, it follows that $x_t \sim I(1)$, and $y_t \sim I(1)$. However $x_t + \alpha y_t = e_t \sim I(0)$. In this example, even though both $\{x_t\}$ and $\{y_t\}$ are $I(1)$, a linear combination of them has an order of integration that is lower than that of each $\{x_t\}$ and $\{y_t\}$. Hence $\{x_t\}$ and $\{y_t\}$ are cointegrated with a cointegrating vector $[1 : \alpha]'$, and $x_t + \alpha y_t$ may be interpreted as an equilibrium relation. In the long-run, the variables $\{x_t\}$ and $\{y_t\}$ move towards the equilibrium even though this relation need not hold exactly even as $t \rightarrow \infty$.

An important property that characterizes a set of cointegrated variables is that such a set of variables has, among other representations, an error correction representation. To see this, let us express equations (6.1) and (6.2) in first difference and use $x_t + \alpha y_t$ to obtain:

$$\Delta x_t + \beta \Delta y_t = \varepsilon_{1t},$$

$$\Delta x_t + \alpha \Delta y_t = \varepsilon_{2t} - (1 - \rho)x_{t-1} - \alpha(1 - \rho)y_{t-1}.$$

In matrix notation, the two equations above may be written as:

$$\begin{bmatrix} 1 & \beta \\ 1 & \alpha \end{bmatrix} \begin{bmatrix} \Delta x_t \\ \Delta y_t \end{bmatrix} = \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} - (1 - \rho)x_{t-1} - \alpha(1 - \rho)y_{t-1} \end{bmatrix}.$$

This gives:

$$\begin{bmatrix} \Delta x_t \\ \Delta y_t \end{bmatrix} = (\alpha - \beta)^{-1} \begin{bmatrix} (\alpha \varepsilon_{1t} - \beta \varepsilon_{2t}) + (1 - \rho)(x_{t-1} + \alpha y_{t-1}) \\ (\varepsilon_{2t} - \varepsilon_{1t}) - (1 - \rho)x_{t-1} - \alpha(1 - \rho)y_{t-1} \end{bmatrix}.$$

Therefore, we may write:

$$\begin{bmatrix} \Delta x_t \\ \Delta y_t \end{bmatrix} = (\alpha - \beta)^{-1} \begin{bmatrix} \beta(1-\rho) & \beta\alpha(1-\rho) \\ -(1-\rho) & -\alpha(1-\rho) \end{bmatrix} \begin{bmatrix} x_{t-1} \\ y_{t-1} \end{bmatrix} + \begin{bmatrix} v_{1t} \\ v_{2t} \end{bmatrix},$$

where

$$\begin{bmatrix} v_{1t} \\ v_{2t} \end{bmatrix} = \begin{bmatrix} \alpha\varepsilon_{1t} - \beta\varepsilon_{2t} \\ \varepsilon_{2t} - \varepsilon_{1t} \end{bmatrix}.$$

This may be re-written as:

$$\Delta x_t = \beta\delta[x_{t-1} + \alpha y_{t-1}] + v_{1t} = \theta_1[x_{t-1} + \alpha y_{t-1}] + v_{1t},$$

$$\Delta y_t = -\delta[x_{t-1} + \alpha y_{t-1}] + v_{2t} = \theta_2[x_{t-1} + \alpha y_{t-1}] + v_{2t},$$

where $\delta = (\alpha - \beta)^{-1}(1 - \rho)$, which is an error correction representation. From this, one can see that δ is non-zero if and only if $\rho \neq 1$. If $\rho = 1$, then both u_t and e_t are random walk processes and in that case $\{x_t\}$ and $\{y_t\}$ are not cointegrated. If they are cointegrated then δ has to be less than unity (in absolute value) and this leads to $\delta \neq 0$ in the error correction model and the error correction representation becomes valid. Conversely, if an error correction model exists then the assumption that $\alpha \neq \beta$ implies that $|\rho| < 1$ and $x_t + \alpha y_t$ is a stationary process and hence x_t and y_t are cointegrated.

The cointegration relation is interpreted as a long-run equilibrium relation while the error correction term is interpreted as temporary divergence from the long-run equilibrium. The coefficients θ_1 and θ_2 are adjustment coefficients and are interpreted as the speeds at which the variables x_t and y_t adjust toward the long-run equilibrium path in response to last period's deviation from such equilibrium. It is noteworthy that θ_1 and θ_2 have opposite signs implying different directions of movements in x_t and y_t in response

to deviations from long-run equilibrium path.

Testing for Cointegration

Since the publication of Engle and Granger (1987), numerous methods have been developed that attempt to test for cointegration among a set of integrated variables².

Commonly used testing procedures fall into one of the following two categories: those based on static regression, and those that are based on the cointegrating rank.

Single Equation Method: The Engle-Granger Procedure

Tests of cointegration based on static regression focus on distinguishing between no cointegration, and at least one cointegration vector. Such tests do not allow estimation of more than one cointegrating vector. The test suggested by Engle and Granger (1987) falls within this category. They suggest a two-step procedure and can be implemented as described below.

All variables that are potentially cointegrated are tested first for the order of integration. Once it is found that all of the integrated variables have the same order of integration, one of the variables is regressed on the remaining variables. The choice of the regressand is arbitrary as long as its coefficient in the cointegrating vector is non-zero. Thus, if the n -dimensional vector y_t has all components that are integrated of the same order, for example $I(1)$, then the first step is to estimate the following regression using ordinary least squares (OLS) estimator:

$$y_{1t} = \theta' y_{2t} + \varepsilon_t,$$

where

$$y_t = \begin{bmatrix} y_{1t} & y_{2t} \\ 1 \times 1 & (n-1) \times 1 \end{bmatrix}.$$

If the variables in y_t are cointegrated then there must exist at least one $(n-1) \times 1$ vector θ such that $\hat{\varepsilon}_t = y_{1t} - \hat{\theta}'y_{2t}$ is stationary ($I(0)$). This idea forms the basis of the tests for cointegration using static regression framework.

Thus, testing for cointegration amounts to testing for stationarity of the estimated residual: $\hat{\varepsilon}_t = y_{1t} - \hat{\theta}'y_{2t}$. The commonly used tests for unit roots such as the Augmented Dickey-Fuller (ADF) test (Dickey and Fuller 1979), or the Phillips-Perron test (Phillips and Perron 1988) can be applied in the univariate case. The null hypothesis that is tested is that of no cointegration against the alternative of at least one cointegrating vector.

However, the critical values of these test can not be used because of the facts that $\hat{\theta}$ is estimated using sample data, as well as due to the multivariate nature / dimension of the vector y_{2t} . Appropriate critical values for the Augmented Dickey-Fuller t-test and Phillips-Perron $z(\pi)$ and $z(t_\pi)$ tests are tabulated in Engle and Yoo (1987) and Phillips and Ouliaris (1990), for y_{2t} with dimensions ranging between 1 and 5. MacKinnon (1990) provides results for an extensive set of simulations.

Stock (1987) shows that OLS estimate of the vector θ is *super consistent* in the sense that as sample size increases $\hat{\theta}$ converges to true θ at a rate that is faster than standard OLS estimator for stationary variables. However, the static regression-based OLS estimator of the cointegrating vector suffers from sizable finite sample bias. Monte

Carlo experiments suggest that a large number of observations (relative to typical sample sizes available for real world economic data) may be needed before the biases become small. These biases are strongly correlated with $(1 - R^2)$, where R^2 is the coefficient of determination. This suggests that cointegrating regression with values of R^2 well below unity should be viewed with great caution³.

However, in the context of multivariate regression, a high value of R^2 is not sufficient to guarantee that biases are small. This is because the R^2 of an equation can not fall when additional variables are added to it. Thus, the inference that high values of R^2 imply low biases, especially where a high R^2 may have been achieved by an *ad hoc* addition of regressors, is not valid⁴. Another weakness of this class of tests is that the test statistics do not have well defined limiting distributions, and testing for cointegration is not a straight forward procedure (Hall 1989). For instance, testing for unit roots in the individual variables is a prerequisite for the cointegration test, yet the critical values are not adjusted accordingly and there is no theory that allows any such adjustment (Campbell and Perron, 1991).

The problem of finite sample bias of static OLS estimates of the cointegrating vector may partially be overcome through dynamic specification of the regression equation. Cointegration tests based on *proper dynamic specification* of the regression equation have higher power compared with their static counterparts. However, the choice of an appropriate dynamic structure (lag structure) is critical in this context. An underspecified regression equation may yield results that are inferior to those obtained

from a static equation framework.

Further, the superior properties of the estimates and tests from dynamic models depend critically on the validity of the assumption that the regressors are at least weakly exogenous. Perhaps the most important weakness of this class of testing and estimation procedure is that they can not address the issue of multiple cointegrating relations among the variables in question. Specifically, when there are multiple cointegrating vectors, the Engle-Granger procedure yields inconsistent parameter estimates.

While the estimation and testing based on a single equation model are convenient and often efficient, for some purposes only a systems approach is desirable. Violation of weak exogeneity assumption for the regressors, and the presence of multiple cointegrating relations among the variables are specific cases where the systems approach is distinctly superior to the single equation model. The systems methods are based on the cointegrating rank of the system of variables. Within this class of methods the maximum likelihood procedure by Johansen (1988) is the most commonly used method in the literature.

Systems Method: The Johansen Procedure

Johansen's maximum likelihood procedure (Johansen 1988) can be illustrated as follows⁵: let us assume that y_t is an n -dimensional vector whose elements are integrated of order 1 ($y_t \sim I(1)$). Let us consider the following autoregressive representation of y_t :

$$y_t = \mu + \Pi_1 y_{t-1} + \Pi_2 y_{t-1} + \dots + \Pi_k y_{t-k} + \varepsilon_t, \quad t = 1, 2, \dots, T, \quad (6.5)$$

where ε_t is an n -dimensional vector of independently and identically distributed (iid)

normal variates with mean zero and covariance matrix Ω . Equation (6.5) may be reparameterized to obtain:

$$\Delta y_t = \mu + \sum_{i=1}^{k-1} \Gamma_i \Delta y_{t-i} + \Pi y_{t-k} + \varepsilon_t, \quad (6.6)$$

where

$$\Gamma_i = -(I - \Pi_1 - \Pi_2 - \dots - \Pi_i), \quad (i = 1, 2, \dots, k-1), \text{ and } \Pi = -\left(I - \sum_{i=1}^k \Pi_i\right).$$

Now the issue of potential cointegration may be investigated by comparing both sides of equation (6.6). Since $y_t \sim I(1)$, $\Delta y_t \sim I(0)$, and so are Δy_{t-i} . That gives the left hand side of equation (6.6) stationary. Since Δy_{t-i} are all stationary, the right hand side of equation (6.6) will be stationary if Πy_{t-k} is also stationary. Stationarity of Πy_{t-k} requires that Π has less than full column rank. Johansen's test for cointegration centers around testing whether Π has less than full rank. Depending on the rank of Π , three distinct cases can be identified:

1. $\text{rank}(\Pi) = n$ (full rank) which implies that all elements of y_t are stationary.
2. $\text{rank}(\Pi) = 0$ (Π is a null matrix) implying that there is no linear combination of y_t which is stationary.
3. $\text{rank}(\Pi) = r$, $0 < r < n$, in which case Π may be written as the product of $(n \times r)$ matrices α and β such that $\Pi = \alpha\beta'$.

In the first case, the issue of cointegration is not relevant since all elements of y_t are stationary. In case 2 clearly there is no linear combination of y_t that is stationary

implying that the elements of y_t are not cointegrated. It is only in case 3 that all or some variables in y_t are cointegrated. This can be seen by observing that for Πy_{t-k} to be stationary $\beta' y_{t-k}$ must be stationary. Therefore the r linearly independent columns of β provide r cointegrating vectors. However, the partition $\Pi = \alpha\beta'$ is not unique as, for any non-singular matrix G , $\Pi = (\alpha G)(G^{-1}\beta')$ is also a valid partition. This implies that data can provide information only about the space spanned by α and β . In practice, one can normalize each column provided it is known *a priori* that this particular element of the column of β (with respect to which the normalization is to be done) is non-zero.

Testing for cointegration, and estimation of α and β , may be carried out as follows: first Δy_t may be regressed on lagged differences Δy_{t-i} to obtain a set of residuals, R_{0t} . Similarly, y_{t-k} may be regressed on the same lagged differences Δy_{t-i} to obtain a second set of residuals, R_{kt} . Now the log-likelihood function may be written (up to a factor of proportionality) in terms of these residuals as:

$$\ln L(\alpha, \beta, \Omega) = -\frac{T}{2} \ln |\Omega| - \frac{1}{2} \sum_{t=1}^T [(R_{0t} + \alpha\beta'R_{kt})\Omega^{-1}(R_{0t} + \alpha\beta'R_{kt})] \quad (6.7)$$

Assuming β fixed, the above log-likelihood function may be maximized with respect to α and Ω by regressing R_{0t} on $-\beta'R_{kt}$ to yield:

$$\hat{\alpha}(\beta) = s_{0k}\beta(\beta's_{kk}\beta)^{-1}, \quad (6.8)$$

$$\hat{\Omega}(\beta) = s_{00}(-s_{0k})\beta(\beta's_{kk}\beta)^{-1}\beta's_{k0}, \quad (6.9)$$

where

$$s_{ij} = \frac{1}{T} \sum_{t=1}^T R_{it} R'_{jt}, \quad i, j = 0, k.$$

Therefore, maximizing the log-likelihood function (6.7) is equivalent to minimizing the expression

$$\left| s_{00} - s_{0k} \beta (\beta' s_{kk} \beta)^{-1} \beta' s_{0k} \right|.$$

It can be shown that minimizing the above expression is equivalent to minimizing the following expression with respect to β :

$$\frac{|\beta' s_{kk} \beta - \beta' s_{k0} s_{00}^{-1} s_{0k} \beta|}{|\beta' s_{kk} \beta|}.$$

The maximum likelihood estimator of β is obtained by solving the equation:

$$|\lambda s_{kk} - s_{k0} s_{00}^{-1} s_{0k}| = 0.$$

and obtaining n estimated eigenvalues $(\hat{\lambda}_1 > \hat{\lambda}_2 > \dots > \hat{\lambda}_n)$, and corresponding n eigenvectors $(\hat{v}_1, \hat{v}_2, \dots, \hat{v}_n)$. The matrix of eigenvectors, V , is normalized such that $\hat{V}' s_{kk} V = I$, and the cointegrating vectors are given by the r statistically significant eigenvectors. That is:

$$\hat{\beta}' = (\hat{v}_1, \hat{v}_2, \dots, \hat{v}_r).$$

Once β is estimated, α can be obtained from equation (6.8).

Testing for the existence of potential cointegrating relationships among the variables in y_t involves testing for statistically significant eigenvalues (λ_i) . The eigenvectors (v_i) corresponding to the statistically significant eigenvalues (λ_i) are the

coefficients of the variables in the cointegrating relationship. Johansen (1988) suggests the following two likelihood ratio tests, depending on the null and alternative hypotheses considered:

Trace Test:

Null hypothesis (H_0): There are at most q cointegrating relations ($r \leq q$).

Alternative hypothesis (H_A): $r = n$.

$$\text{Test statistic: } \lambda_{tr} = -T \sum_{i=q+1}^n \ln(1 - \hat{\lambda}_i). \quad (6.10)$$

Maximum Eigenvalue Test:

Null hypothesis (H_0): There are at most q cointegrating relations ($r \leq q$).

Alternative hypothesis (H_A): $r = q+1$.

$$\text{Test statistic } \lambda_{max} = -T \ln(1 - \hat{\lambda}_{q+1}). \quad (6.11)$$

Either of these test statistics does not follow any standard distribution as the estimated eigenvalues correspond to $(n-r)$ non-stationary common trends rather than stationary linear combination of the data. Empirical distributions are multivariate versions of the Dickey-Fuller distribution and are derived in terms of a multivariate Brownian motion. The empirical distributions of these tests have been calculated and reported in Johansen and Juselius (1990) for values of $(n - r)$ between 1 to 5⁶.

Once tests for cointegration have been performed and the cointegrating vector obtained, one can test for linear restrictions on the cointegrating vectors, β and / or the adjustment vectors α . Testing procedure involves estimating the likelihood ratio statistic:

$$\lambda_{LR} = -T \sum_{i=1}^r \ln\left(\frac{1+\tilde{\lambda}_i}{1-\hat{\lambda}_i}\right), \quad (6.12)$$

where r is the number of significant eigenvalues obtained by applying the trace and the maximum eigenvalue tests defined by equations (6.10) and (6.11), and $\hat{\lambda}_i$ and $\tilde{\lambda}_i$ are the estimated eigenvalues from the unrestricted and restricted (under the null hypothesis being tested) models respectively. Johansen (1988) shows that the test statistic λ_{LR} is asymptotically distributed as χ^2 with rs degrees of freedom, where s is the number of linear restrictions under the null hypothesis.

Empirical Results

Johansen's maximum likelihood method is employed to test for the existence of and estimate potential long-run relationships among current account balance, permanent components of income and real exchange rate, government consumption ($I(1)$ variables). Due to data limitations, real income and government consumption variables are defined in relative form⁷. Relative income is defined as (natural log of) home country's income minus (natural log of) rest of the world income. Relative government consumption is defined in similar fashion.

This section presents the empirical results of the cointegration analysis. First, the results for the US vs. rest of the world are presented. Then the results for Japan vs. rest of the world (vis-à-vis Japan) are presented. For each country, results are presented both for the Beveridge-Nelson (B-N), and the Blanchard-Quah (B-Q) decomposition of income and real exchange rate.

Since quarterly data is used in the analysis, potentially one should try lag-lengths in multiples of four. However, as the available number of observations is not very large, attempt is made to test and estimate cointegrating relations with the minimum acceptable lag-length in the VAR system. Initially, starting with a lag-length of 8 in the VAR, Sim's lag-length test is used to choose a shorter acceptable lag structure in testing and estimating the cointegrating relationships among the variables. In all cases, lag-length of 4 is found acceptable. After performing the cointegration test and estimating the cointegrating vectors, the residuals are used to test for normality and iid properties using the Jarque-Bera (1980) test and the Ljung-Box Q-statistic, respectively. The empirical results of cointegration tests for US data are presented below. Then similar results for Japanese data will be presented.

Cointegration Analysis: The United States

Diagnostic Checking of Model Specification

The empirical results discussed are conditional on proper specification of the lag structure so that the error terms in Johansen's procedure satisfy the assumptions of normality and iid. Therefore, it is important to check the residuals of the Johansen model for normality and iid properties. The test of normality of the residuals is performed with Jarque-Bera test (J-B test), while the iid assumption is tested by using the Ljung-Box Q-statistics. In implementing the Johansen (1988) procedure, the following set of equations is estimated:

$$\Delta y_t = \mu + \sum_{i=1}^k \Gamma_i \Delta y_{t-i} + \Pi y_{t-(k+1)} + \Phi z_t + \varepsilon_t,$$

where

$$y_t = \begin{bmatrix} CA_t \\ Y_{pr,t} \\ R_{pr,t} \\ G_t \end{bmatrix}, \mu = \begin{bmatrix} \mu_{ca} \\ \mu_Y \\ \mu_R \\ \mu_G \end{bmatrix}, \varepsilon_t = \begin{bmatrix} \varepsilon_{ca,t} \\ \varepsilon_{Y_{pr,t}} \\ \varepsilon_{R_{pr,t}} \\ \varepsilon_{G,t} \end{bmatrix} \text{ and}$$

z_t represents the stationary components of income and real exchange rate included in estimating the equation. Here CA denotes current account balance, Y_{pr} denotes relative permanent income, R_{pr} denotes permanent component in real exchange rate, and G denotes relative government consumption.

The normality and iid properties are tested for the elements of ε_t vector defined above. The results of these tests for US data are presented in Table 6.1⁸. Under the assumption of normality of the residuals, the Jarque-Bera statistic (τ) follows $\chi^2(2)$ distribution. The 95% critical value for $\chi^2(2)$ is 5.99. The sample estimates of the J-B test for the residuals are all less than 5.90. In fact, it can be seen from Table 6.1, except for the equation representing the current account balance, the estimated values of the test statistic are very low compared to the 95% critical value of the Jarque-Bera statistic. Since the sample estimates of the J-B statistic are lower than $\chi_{0.95}^2(2)$ for data from both decomposition methods and for all equations, the normality assumption is not contradicted by the residuals. The Q-statistics with lags up to 16 are calculated.

Table 6.1. Test of Normality and iid for the Residuals: US Data

Results from Beveridge-Nelson Decomposition				
Test Statistic	Residual from Equation			
	$\varepsilon_{ca,t}$	$\varepsilon_{Y_{pr},t}$	$\varepsilon_{R_{pr},t}$	$\varepsilon_{G,t}$
J-B Test: (τ)	5.89	0.59	0.29	0.49
Ljung-Box Test: Q(4)	3.98	0.98	0.12	0.39
Ljung-Box Test: Q(16)	17.91	23.56	12.88	11.73

Results from Blanchard-Quah Decomposition				
J-B Test (τ)	2.01	1.42	0.004	0.22
Ljung-Box Test: Q(4)	4.02	0.45	0.17	0.74
Ljung-Box Test: Q(16)	15.97	19.68	11.63	13.54

With either method of decomposition, the Q-statistics at all lags up to 16 are insignificant implying that the iid assumption of the Johansen method is not violated.

Results of Cointegration Test

Now we present the results of cointegration test based on Johansen's procedure (1988). The methodology used here follows the procedure outlined and implemented by Johansen and Juselius (1990). Here the results of cointegration tests for US data are presented. In what follows, cointegration test results with data from Beveridge-Nelson (B-N) decomposition method is presented first. Then the results for data from the Blanchard-Quah (B-Q) decomposition are presented.

Johansen Procedure with Beveridge-Nelson Decomposition

Results of Johansen's procedure, using the data from the Beveridge-Nelson decomposition, are summarized in Tables 6.2 and 6.3. These results are obtained by estimating the following set of equations;

$$\Delta y_t = \mu + \sum_{i=1}^k \Gamma_i \Delta y_{t-i} + \Pi y_{t-(k+1)} + \Phi z_t + \varepsilon_t, \quad k = 4,$$

where

$$y_t = \begin{bmatrix} CA_t \\ Y_{pr,t} \\ R_{pr,t} \\ G_t \end{bmatrix}, \mu = \begin{bmatrix} \mu_{ca} \\ \mu_Y \\ \mu_R \\ \mu_G \end{bmatrix}, \varepsilon_t = \begin{bmatrix} \varepsilon_{ca,t} \\ \varepsilon_{Y_{pr},t} \\ \varepsilon_{R_{pr},t} \\ \varepsilon_{G,t} \end{bmatrix}, \text{ and } z_t \text{ is as defined earlier.}$$

It should be mentioned at this point that the trace test is a test of the null hypothesis that there are at most q cointegrating vectors ($r \leq q$) against the general alternative that $r \geq q+1$, where r is the number of cointegrating vectors. But it does not help to pin down the number of such relations. For that, the maximum eigenvalue test may be used which tests the null that $r \leq q$ against the alternative that $r = q + 1$. Further, the tests are to be performed sequentially starting with the null that $r = 0$. The test result at a particular stage of the sequence is conditional on the results from earlier stages of the test.

It is evident from Table 6.2 that using the trace test (λ_{tr}), and starting with null of no cointegration ($r = 0$), we can successively reject the null hypotheses of no cointegration ($r = 0$), against the alternative that there is at least one cointegrating

Table 6.2. Johansen's Trace Test: US Data and B-N Decomposition

<u>Statistic and Critical Values</u>				
Null	Alternative	Statistic	$\lambda_{tr}(0.95)$	$\lambda_{tr}(0.90)$
$r = 0$	$r \geq 1$	86.70	53.12	49.648
$r \leq 1$	$r \geq 2$	44.98	34.91	32.00
$r \leq 2$	$r \geq 3$	20.52	19.96	17.85
$r \leq 3$	$r = 4$	3.44	9.24	7.53

Table 6.3. Johansen's Maximum eigenvalue Test: US Data and B-N Decomposition

<u>Statistics and Critical Values</u>				
Null	Alternative	Statistic	$\lambda_{\max}(0.95)$	$\lambda_{\max}(0.90)$
$r = 0$	$r = 1$	41.73	28.14	25.56
$r \leq 1$	$r = 2$	24.45	22.00	19.77
$r \leq 2$	$r = 3$	17.08	15.67	13.75
$r \leq 3$	$r = 4$	3.44	9.24	7.53

vector($r \geq 1$), $r \leq 1$ against the alternative that $r \geq 2$, and $r \leq 2$ against the general alternative that $r \geq 3$ at 5% significance level. However, the trace test can not reject the null that $r \leq 3$ at 5% or 10% significance level. So, the trace test (λ_{tr}) suggests that there are at most three cointegrating relations among US current account balance, relative permanent income, permanent component of real exchange rate, and relative government

consumption. The results of maximum eigenvalue test are reported in Table 6.3. Using the maximum eigenvalue test (λ_{\max}), we can reject the null hypotheses of $r = 0$ (against $r = 1$), $r \leq 1$ (against $r = 2$), $r \leq 2$ (against $r = 3$) at 5% significance level. However, we can not reject the null that $r \leq 3$ (against $r = 4$). Hence, given the result of λ_{tr} -test and the λ_{\max} -test, we may conclude that there are three cointegrating relations ($r = 3$) among the variables in question.

Corresponding to variable ordering [CA, Y_{pr} , R_{pr} , G, constant], the estimated cointegrating vectors are:

$$\beta' = \begin{bmatrix} -1.00 & -2.23 & 1.75 & -0.78 & -2.07 \\ -1.00 & -5.77 & 0.13 & -0.71 & -2.39 \\ -1.00 & -1.30 & -0.28 & -0.35 & -0.42 \end{bmatrix}$$

where CA denotes current account balance, Y denotes relative income, R denotes real exchange rate, G denotes relative government consumption, and subscript pr indicates permanent component. Thus, normalizing with respect to current account balance the estimated cointegrating vectors may be written as:

$$\text{relation 1: } CA^{us} = -2.07 - 2.23Y_{pr} + 1.75R_{pr} - 0.78G$$

$$\text{relation 2: } CA^{us} = -2.39 - 5.77Y_{pr} + 0.13R_{pr} - 0.71G$$

$$\text{relation 3: } CA^{us} = -0.42 - 1.30Y_{pr} - 0.28R_{pr} - 0.42G$$

When there is only one cointegrating vector, the coefficient of a variables is interpreted as the long-run effect of the variables in question on the variable with respect to which the normalization is done. However, if there are more than one cointegrating

vectors, there is no obvious interpretation for the coefficients in the estimated vectors. Nevertheless, it may be observed that in all three relations both real income and government consumption have negative effects on current account balance. The sign of relative permanent income is consistent with many traditional models of current account determination while the intertemporal model predicts that such permanent income change should not have any effect on current account balance. The sign of relative government consumption is generally consistent with most models. In two of the three relations, permanent increase in relative price in rest of the world leads to improvement in US current account balance. Our model in Chapter 3 predicted no effect of permanent real exchange rate change on current account. That prediction is dependent on the homotheticity of preferences of the agents. In a more general setting, however, the intertemporal models do not assign any specific sign to this coefficient. Depending on the parameters of the model, both positive and negative signs can be consistent with this class of models. The interesting question is whether the effects of permanent changes in income and real exchange rate are statistically significantly different from zero. That issue is addressed later in this section.

Corresponding to the cointegrating vectors above, the adjustment vectors are:

$$\hat{\alpha}_1 = [-0.075 \quad 0.013 \quad -0.011 \quad 0.030]$$

$$\hat{\alpha}_2 = [-0.091 \quad 0.047 \quad 0.018 \quad -0.041]$$

$$\hat{\alpha}_3 = [0.348 \quad 0.025 \quad -0.086 \quad 0.026]$$

As mentioned earlier, the coefficients of an adjustment vector are interpreted as the speeds of adjustment by the associated variables in response to deviation from the long-run equilibrium. The estimated adjustment vectors suggest that in the case of US data it is the current account balance that adjusts most in response to deviations from equilibrium.

Test of Exclusion Restriction on the Coefficients

From the perspective of the present study, we are interested in testing whether permanent changes in real income and real exchange rate have significant effects on current account balance. Intertemporal models of current account determination, as discussed in chapter 3, suggest that permanent changes in income should leave current account balance unaffected. On the other hand, permanent movements in real exchange rate may or may not have significant effects on current account balance. Under the assumption of homotheticity of the intertemporal utility function, and starting from an initial condition of balance, a permanent real exchange rate shock leaves the current account balance unaffected. We may test whether these propositions of the intertemporal models are supported by actual data. These propositions may be tested by imposing restrictions on the cointegrating vectors that the coefficients corresponding to relative permanent income and permanent component of real exchange rate are zero.

Table 6.4 represents the results of such tests of restrictions on the coefficients of the cointegration vector. Test of the restriction that relative permanent income has no effect on current account balance gives the test statistic $\chi_c^2 = 18.68$. Under the null

Table 6.4. Test of Restrictions on Cointegration Vectors: US Data

<u>Test Results based on Beveridge-Nelson Decomposition</u>			
Null Hypothesis	Estimated Test Statistic: χ_c^2	Distribution of χ_c^2 under H_0	95% Critical Value
Y_{pr} has no effect	18.68	$\chi^2(3)$	7.82
R_{pr} has no effect	20.23	$\chi^2(3)$	7.82
G has no effect	15.90	$\chi^2(3)$	7.82

hypothesis, test statistic χ_c^2 follows χ^2 -distribution with 3 degrees of freedom (since we have three cointegrating vectors). The tabulated value of $\chi_{0.95}^2(3)$ is 7.82. Since sample values of the test statistic exceed the tabulated value, the null hypothesis of no effect of relative permanent income on current account balance is rejected at 5% significance level.

Similarly, tests of the restriction that permanent real exchange rate changes have no effect on current account balance gives estimated test statistic $\chi_c^2 = 20.23$. Under the null hypothesis, test statistic χ_c^2 follows χ^2 -distribution with 3 degrees of freedom. The tabulated value of $\chi_{0.95}^2(3)$ is 7.82. So, the null hypothesis of no effect of permanent real exchange rate on current account balance can be rejected at 5% significance level.

Tests of the restriction that relative government consumption has no effect on current account yields estimated test statistic $\chi_c^2 = 15.90$. This estimated value of χ_c^2 exceeds the tabulated value of $\chi_{0.95}^2(3)$. Therefore, the null hypothesis can be rejected at 5% significance level .

Thus, we observe that permanent changes in real exchange rate do have statistically significant effect on US current account balance. However, these test results are not necessarily contradictory to the conclusions of the intertemporal models. It is only under specific assumption about preference pattern (specifically, homotheticity of preferences) that we are able to predict that permanent real exchange rate changes shall leave current account balance unaffected. On the other hand, according to our model, permanent income changes are not expected to affect current account balance. However, permanent real income changes do have significant effects on US current account balance. Thus, the empirical evidence is against the conclusion of the intertemporal models regarding the effects of permanent income changes on current account balance.

Johansen Procedure with Blanchard-Quah Decomposition

The cointegration test results, with permanent and transitory decomposition obtained by using the Blanchard-Quah method, are summarized in Table 6.5 and 6.6. Table 6.5 reports the results of the trace test, while Table 6.6 contains the results of the maximum eigenvalue test. As before, in implementing the Johansen (1988) procedure, the following model is estimated:

$$\Delta y_t = \mu + \sum_{i=1}^k \Gamma_i \Delta y_{t-i} + \Pi y_{t-(k+1)} + \Phi z_t + \varepsilon_t, \quad k = 4,$$

where

$$y_t = \begin{bmatrix} CA_t \\ Y_{pr,t} \\ R_{pr,t} \\ G_t \end{bmatrix}, \mu = \begin{bmatrix} \mu_{ca} \\ \mu_Y \\ \mu_R \\ \mu_G \end{bmatrix}, \varepsilon_t = \begin{bmatrix} \varepsilon_{ca,t} \\ \varepsilon_{Y_{pr,t}} \\ \varepsilon_{R_{pr,t}} \\ \varepsilon_{G,t} \end{bmatrix},$$

and z_t is as defined earlier.

Table 6.5. Johansen's Trace test: US data and B-Q Decomposition

<u>Statistics and Critical Values</u>				
Null	Alternative	Statistic	$\lambda_{tr}(0.95)$	$\lambda_{tr}(0.90)$
$r = 0$	$r \geq 1$	86.62	53.12	49.648
$r \leq 1$	$r \geq 2$	44.69	34.91	32.00
$r \leq 2$	$r \geq 3$	19.36	19.96	17.85
$r \leq 3$	$r = 4$	3.21	9.24	7.53

Table 6.6. Johansen's Maximum eigenvalue test: US data and B-Q Decomposition

<u>Statistics and Critical Values</u>				
Null	Alternative	Statistic	$\lambda_{max}(0.95)$	$\lambda_{max}(0.90)$
$r = 0$	$r = 1$	38.93	28.14	25.56
$r \leq 1$	$r = 2$	25.33	22.00	19.77
$r \leq 2$	$r = 3$	16.15	15.67	13.75
$r \leq 3$	$r = 4$	3.21	9.24	7.53

Now, it is evident from Table 6.5 that the trace test (λ_{tr}) leads to rejection of null hypotheses of no cointegration ($r = 0$) against the general alternative $r \geq 1$, and at most one cointegrating vector ($r \leq 1$) against the general alternative $r \geq 2$ at 5% significance level. The null that there are at most 2 cointegrating vectors ($r \leq 2$) against the general alternative that $r \geq 3$ can not be rejected at 5% significance level. However, the null can be rejected at 10% significance level. It is well known by now that the cointegration tests generally have low power, and for that reason some researchers allows significance level of 10% when the λ_{tr} is close to $\lambda_{tr}(0.95)$ but is less than the 95% critical value of the statistics. On the basis of the result at 10% significance level, the trace test (λ_{tr}) suggests that there are at most three cointegrating vectors. The null that $r \leq 3$ against the general alternative that $r = 4$ can not be rejected at 5% or 10% significance level. Thus, the trace test (λ_{tr}) leads to the conclusion that there are at most 2, or possibly 3 cointegrating vectors.

Results of the maximum eigenvalue test are presented in Table 6.6. We can see from Table 6.6 that the maximum eigenvalue test clearly rejects the null hypotheses that $r = 0$ (against the alternative $r = 1$), $r \leq 1$ (against the alternative $r = 2$), and $r \leq 2$ (against the alternative $r = 3$) at conventional significance level. However, the null that $r \leq 3$ against the alternative $r = 4$ can not be rejected at 5% significance level. Hence, the maximum eigenvalue test suggest the presence of 3 cointegrating vectors. Given these results, we would accept three cointegrating vectors in the data set.

The estimated cointegrating vectors are:

$$\beta' = \begin{bmatrix} -1.00 & -0.62 & 1.88 & -0.70 & -1.55 \\ -1.00 & -17.95 & 2.18 & -1.87 & -8.45 \\ -1.00 & -2.21 & -0.17 & -0.43 & -0.84 \end{bmatrix},$$

where the variables are ordered as $[CA, Y_{pr}, R_{pr}, G, \text{constant}]$, and as before, Y denotes relative income, R denotes real exchange rate, G denotes relative government consumption, and subscript pr indicates permanent component. Normalizing with respect to current account balance, the estimated cointegrating relations may be written as:

$$\text{relation 1: } CA^{us} = -1.55 - 0.62Y_{pr} + 1.88R_{pr} - 0.70G,$$

$$\text{relation 2: } CA^{us} = -8.45 - 17.95Y_{pr} + 2.18R_{pr} - 1.87G.$$

$$\text{relation 3: } CA^{us} = -0.84 - 2.21Y_{pr} - 0.17R_{pr} - 0.43G$$

The estimated cointegration vectors have signs that are similar to those obtained with data from Beveridge-Nelson decomposition. In all three relations, permanent income shows negative effect on current account balance. This is more in accord with traditional models. In two of the three estimated relations, increase in relative price in rest of the world leads to improvement in US current account balance. Only in one case the reverse is true. In all three relations, current account deteriorates with increase in government consumption which is similar to the result obtained with Beveridge-Nelson decomposition.

Results of tests of restrictions that permanent components in (relative) real income and real exchange rate, and relative government consumption have no effect on the current account balance of the United States are presented in Table 6.7.

Table 6.7. Test of Restriction on Cointegration Vectors: US Data

<u>Test Results based on Blanchard-Quah Decomposition</u>			
Null Hypothesis	Estimated Test Statistic: χ_c^2	Distribution of χ_c^2 under H_0	95% Critical Value
Y_{pr} has no effect	15.83	$\chi^2(3)$	7.82
R_{pr} has no effect	21.68	$\chi^2(3)$	7.82
G has no effect	13.09	$\chi^2(3)$	7.82

It is evident from Table 6.7 that the test of the restriction that relative permanent income has no effect on current account balance gives the test statistic $\chi_c^2 = 15.83$. Under the null hypothesis, test statistic χ_c^2 follows χ^2 -distribution with 3 degrees of freedom (since we have three cointegrating vectors under both decomposition methods). The tabulated value of $\chi_{0.95}^2(3)$ is 7.82. Since sample value of the test statistic exceeds the tabulated value, the null hypothesis of no effect of relative permanent income on current account balance is rejected at 5% significance level.

Similarly, test of the restriction that permanent real exchange rate changes have no effect on current account balance yields an estimated test statistic $\chi_c^2 = 21.68$. Under the null hypothesis, test statistic χ_c^2 follows χ^2 -distribution with 3 degrees of freedom. The tabulated value of $\chi_{0.95}^2(3)$ is 7.82. Since the sample value of the test statistic exceeds the tabulated value, the null hypothesis of no effect of changes in permanent real exchange rate on current account balance can be rejected at 5% significance level.

Test of the restriction that relative government consumption has no effect on current account balance yields an estimated test statistic $\chi_c^2 = 13.09$. This estimated value of χ_c^2 exceeds the tabulated value of $\chi_{0.95}^2(3)$ (which is 7.82). Therefore, the null hypothesis that relative government consumption has no effect on current account balance can be rejected at 5% significance level for US data.

Comparing these results of the tests of restrictions on the coefficients of the cointegrating vectors with those using the Beveridge-Nelson decomposition, we observe that they are qualitatively similar. Under either decomposition method, both permanent real income and permanent real exchange rate have statistically significant effects on US current account balance. These results are not supportive of the general predictions of the intertemporal models.

Cointegration Analysis: Japan

Diagnostic Checking of Model Specification

As has been done in the case of US data, diagnostic checking has been done to ensure that the lag structure used in the Johansen procedure does not violate the assumption of iid and normality of the error term. Results of diagnostic tests on the residuals from the equations for Japan are summarized in Table 6.8. In implementing the Johansen (1988) procedure, the following set of equations is estimated:

$$\Delta y_t = \mu + \sum_{i=1}^k \Gamma_i \Delta y_{t-i} + \Pi y_{t-(k+1)} + z_t + \varepsilon_t,$$

where

$$y_t = \begin{bmatrix} CA_t \\ Y_{pr,t} \\ R_{pr,t} \\ G_t \end{bmatrix}, \mu = \begin{bmatrix} \mu_{ca} \\ \mu_Y \\ \mu_R \\ \mu_G \end{bmatrix}, \varepsilon_t = \begin{bmatrix} \varepsilon_{ca,t} \\ \varepsilon_{Y_{pr,t}} \\ \varepsilon_{R_{pr,t}} \\ \varepsilon_{G,t} \end{bmatrix}, \text{ and } z_t \text{ is stationary part of } Y_t \text{ and } R_t.$$

Here CA denotes current account balance, Y_{pr} denotes relative permanent income, R_{pr} denotes permanent component in real exchange rate, and G denotes relative government consumption.

It is evident from Table 6.8 that the J-B statistics are insignificant for data from both methods of decomposition and for all equations. The Ljung-Box Q-statistics calculated with lags up to 16 are all found to be insignificant at 5% level. Since the

Table 6.8. Test of Normality and iid for Residuals: Japanese Data

Test Statistic	Residual from Equation			
	$\varepsilon_{ca,t}$	$\varepsilon_{Y_{pr,t}}$	$\varepsilon_{R_{pr,t}}$	$\varepsilon_{G,t}$
J-B Test: (τ)	0.81	1.02	0.014	0.86
Ljung-Box Test: Q(4)	3.77	0.15	1.56	2.33
Ljung-Box Test: Q(16)	11.36	11.48	20.25	9.00
<u>Results from Blanchard-Quah Decomposition</u>				
J-B Test (τ)	0.087	0.030	0.541	1.41
Ljung-Box Test: Q(4)	8.15	0.46	0.04	5.09
Ljung-Box Test: Q(16)	25.71	12.48	19.46	16.71

estimated values of the statistics are all insignificant at 5% level, we may infer that the estimated residuals do not contradict the iid assumption. Given these test results (both the J-B test and the Ljung-Box test), it is reasonable to perform the Johansen procedure with 4 lags in the VAR system.

Results of Cointegration Tests

Johansen's Procedure: Beveridge-Nelson Decomposition

The results of cointegration tests using the Beveridge-Nelson decomposition of income and real exchange rate are summarized in Table 6.9 and Table 6.10. The former contains the results of the trace test (λ_{tr}), while the latter reports the results of the maximum eigenvalue test (λ_{max}). It is evident from Table 6.9 that, based on the trace test (λ_{tr}), it is possible to reject the null hypotheses that $r = 0$ (against $r \geq 1$), $r \leq 1$ (against $r \geq 2$), and $r \leq 2$ (against $r \geq 3$) at 5% significance level. However, it is not possible to reject the null hypothesis of $r \leq 3$ against the general alternative that $r = 4$. Thus the λ_{tr} -test suggests that there are at most 3 cointegrating vectors.

Table 6.9. Johansen's Trace Test: Japan and B-N Decomposition

<u>Statistics and Critical Values</u>				
Null	Alternative	Statistic	$\lambda_{tr}(0.95)$	$\lambda_{tr}(0.90)$
$r = 0$	$r \geq 1$	70.29	53.12	49.65
$r \leq 1$	$r \geq 2$	40.28	34.91	32.00
$r \leq 2$	$r \geq 3$	20.70	19.96	17.85
$r \leq 3$	$r = 4$	5.28	9.24	7.53

Table 6.10. Johansen's Maximum eigenvalue Test: Japan and B-N Decomposition

<u>Statistics and Critical Values</u>				
Null	Alternative	Statistic	$\lambda_{\max}(0.95)$	$\lambda_{\max}(0.90)$
$r = 0$	$r = 1$	30.01	28.14	25.56
$r \leq 1$	$r = 2$	19.58	22.00	19.77
$r \leq 2$	$r = 3$	15.43	15.67	13.75
$r \leq 3$	$r = 4$	5.28	9.24	7.53

Table 6.10 shows that on the basis of the maximum eigenvalue test (λ_{\max}) it is possible to reject the null of no cointegration ($r = 0$) against the alternative that there is one cointegrating vector ($r = 1$). However, we can not reject the null that $r \leq 1$ against the alternative that $r = 2$ at 5% significance level. Therefore, the λ_{\max} -test leads to the conclusion that there is only one cointegrating relation between Japanese current account balance, relative permanent income, permanent component of real exchange rate, and relative government consumption.

The estimated cointegrating vector is given by:

$$\beta' = [-1.0 \quad 0.58 \quad 0.75 \quad 0.14 \quad 0.66]$$

where the variables are ordered as [CA, Y_{pr} , R_{pr} , G, constant], and Y denotes relative income, R denotes real exchange rate, G denotes relative government consumption, and the subscript pr indicates permanent component (as has been defined before).

Thus, the estimated long-run relation may be written as:

$$CA^J = 0.66 + 0.58Y_{pr} + 0.75R_{pr} + 0.14G$$

It may be observed that unlike the US data, the sign of permanent real income is not in accord with traditional models of current account determination, as is the sign of relative government consumption. The long-run effect of increases in world relative price on current account balance of Japan is positive which is in accord with traditional theory and does not contradict the intertemporal models.

Corresponding to the estimated cointegrating vector above, the estimated adjustment vector is:

$$\hat{\alpha} = [-0.018 \quad 0.003 \quad -0.159 \quad 0.126]$$

The adjustment vector suggests that relative price adjusts most in response to deviations from long-run equilibrium, while relative income makes the least adjustment.

Results of tests of exclusion restrictions on the variables are presented in Table 6.11 below.

Table 6.11. Test of Restrictions on Cointegrating Vector: Japanese Data

<u>Test Results based on Beveridge-Nelson Decomposition</u>			
Null Hypothesis	Estimated Test Statistic: χ_c^2	Distribution of χ_c^2 under H_0	95% Critical Value
Y_{pr} has no effect	2.23	$\chi^2(1)$	3.84
R_{pr} has no effect	6.97	$\chi^2(1)$	3.84
G has no effect	0.82	$\chi^2(1)$	3.84
Y_{pr} and G have no effect	2.51	$\chi^2(2)$	5.99

Test of the restriction that permanent changes in income have no effect on current account balance gives an estimated value of test statistic $\chi_c^2 = 2.23$. Under the null hypothesis, this test statistic follows χ^2 distribution with 1 degree of freedom (since there is only one cointegrating vector). The tabulated value of $\chi_{0.95}^2(1)$ is 3.84. Hence, the null hypothesis can not be rejected at 5% significance level. Test of similar restriction on permanent real exchange rate gives an estimated $\chi_c^2 = 6.97$. Under the null hypothesis of no effect of permanent real exchange rate changes on current account balance, the test statistics follows $\chi^2(1)$. Since the sample estimate of the test statistic exceed the 95% tabulated value, the null that permanent real exchange rate has no effect on current account is rejected at 5% significance level.

Test of the restriction that relative government consumption has no effect on current account gives $\chi_c^2 = 0.82$. This sample estimate is lower than $\chi_{0.95}^2(1)$, implying that the null can not be rejected at 5% significance level. Test of the null that both relative permanent income and relative government consumption have no effect on current account balance gives a test statistic $\chi_c^2 = 2.5$. Under the null, this test statistic is distributed as χ^2 distribution with 2 degrees of freedom. The tabulated value of $\chi_{0.95}^2(2)$ is 5.99. Since the sample estimate of the test statistic is lower than the 95% tabulated value, the null can not be rejected at 5% significance level. The test result regarding the effect of permanent income on current account balance is supportive of the predictions of the intertemporal models of current account determination.

Finally, with B-N decomposition, the restricted cointegrating vector is:

$$\beta' = [-1.00 \quad 1.27 \quad -1.31]$$

with variable ordering as follows: [CA, R_{pr} , constant], where notations are as before. In other words, the long-run relationship may be written as:

$$CA^J = -1.31 + 1.27R_{pr}.$$

Thus, in the long-run, permanent increase (i.e., increase in world relative price) in real exchange rate improves the current account balance of Japan.

Johansen Procedure: Blanchard-Quah Decomposition

Johansen's cointegration test results, using data from the Blanchard-Quah decomposition method, are presented in Table 6.12 and 6.13. It is evident from Table 6.12 that trace test (λ_{tr}) allows us to reject at 5% significance level the null hypotheses of $r = 0$ (against $r \geq 1$), $r \leq 1$ (against $r \geq 2$), and $r \leq 2$ (against $r \geq 3$). However, the null that $r \leq 3$ against $r = 4$ can not be rejected at 5% or 10% significance level. Thus, the λ_{tr} -test suggest that there are at most three cointegrating relations.

Table 6.13 shows the results of the maximum eigenvalue test. It is clear from the table that, based on the maximum eigenvalue test (λ_{max}), we can reject the null that $r = 0$ against the alternative that $r = 1$ at 5% significance level. However, we can not reject the null hypothesis that $r \leq 1$ against the alternative that $r = 2$ at 5% or 10% significance level. This, along with the λ_{tr} -test leads us to conclude that there is only one cointegrating relation among the variables in question.

Table 6.12. Johansen's Trace Test: Japan and B-Q Decomposition

<u>Statistics and Critical Values</u>				
Null	Alternative	Statistic	$\lambda_{tr}(0.95)$	$\lambda_{tr}(0.90)$
$r = 0$	$r \geq 1$	71.31	53.12	49.65
$r \leq 1$	$r \geq 2$	38.85	34.91	32.00
$r \leq 2$	$r \geq 3$	20.79	19.96	17.85
$r \leq 3$	$r = 4$	4.55	9.24	7.53

Table 6.13. Johansen's Maximum Eigenvalue Test: Japan and B-Q Decomposition

<u>Statistics and Critical Values</u>				
Null	Alternative	Statistic	$\lambda_{max}(0.95)$	$\lambda_{max}(0.90)$
$r = 0$	$r = 1$	32.46	28.14	25.56
$r \leq 1$	$r = 2$	18.06	22.00	19.77
$r \leq 2$	$r = 3$	16.24	15.67	13.75
$r \leq 3$	$r = 4$	4.55	9.24	7.53

The estimated cointegrating vector is:

$$\beta' = [-1.0 \quad 5.41 \quad -11.72 \quad -1.69 \quad 18.81]$$

with the variables ordered as [CA, Y_{pr} , R_{pr} , G, constant], where the notations are as

before. Thus, the estimated long-run relation may be written as:

$$CA^J = 18.81 + 5.41Y_{pr} - 11.72R_{pr} - 1.69G$$

Corresponding to the estimated cointegrating vector above, the estimated adjustment vector is:

$$\hat{\alpha} = [0.022 \quad -0.0006 \quad 0.0012 \quad -0.0059]$$

The adjustment vector suggests that the current account balance adjusts most in response to deviations from long-run equilibrium, while other variables make very little adjustment.

Results of tests of exclusion restrictions on the variables are presented in Table 6.14 below. Test of the restriction that permanent changes in relative income has no effect on current account balance gives an estimated value of test statistic $\chi_c^2 = 1.50$. Under the null hypothesis, this test statistic follows χ^2 distribution with 1 degree of freedom (since there is only one cointegrating vector). The tabulated value of $\chi_{0.95}^2(1)$ is 3.84. Hence the null hypothesis can not be rejected at conventional significance level.

Table 6.14. Test of Restrictions on Cointegrating Vector: Japanese Data

Test Results based on Blanchard-Quah Decomposition

Null Hypothesis	Estimated Test Statistic: χ_c^2	Distribution of χ_c^2 under H_0	95% Critical Value
Y_{pr} has no effect	1.50	$\chi^2(1)$	3.84
R_{pr} has no effect	13.29	$\chi^2(1)$	3.84
G has no effect	0.77	$\chi^2(1)$	3.84
Both Y_{pr} and G have no effect	3.06	$\chi^2(2)$	5.99

Test of similar restriction on permanent real exchange rate gives an estimated $\chi_c^2 = 13.29$. Under the null hypothesis of no effect of permanent real exchange rate changes on current account, the test statistics follows $\chi^2(1)$. Since the sample estimate of the test statistic is greater than the tabulated value, the null hypothesis is rejected at 5% significance level. Restriction that relative government consumption has no effect on current account gives $\chi_c^2 = 0.77$. This estimate is lower than $\chi_{0.95}^2(1)$, implying that the null can not be rejected at 5% significance level. Test of the null that both relative permanent income and relative government consumption have no effect on current account balance gives a test statistic $\chi_c^2 = 3.07$. Under the null, this test statistic is distributed as χ^2 distribution with 2 degrees of freedom. The tabulated value of $\chi_{0.95}^2(2)$ is 5.99. Since the sample estimate of the test statistic is lower than the 95% tabulated value, the null can not be rejected at 5% significance level. Thus the test result regarding the effect of permanent income on current account balance is supportive of the predictions of the intertemporal models of current account determination.

The estimated restricted cointegrating vector using the B-Q decomposition data is:

$$\beta' = [-1.00 \quad 3.46 \quad -3.74].$$

Therefore, the restricted long-run relationship may be written as:

$$JCA = -3.74 + 3.46R_{pr}$$

As in the case with data from Beveridge-Nelson decomposition, a permanent increase in world relative price improves the current account balance of Japan.

It is evident from the discussion above that the empirical results are not conclusive. For US data, movements in the permanent component of real income have statistically significant effect on US current account balance. The directions of such effects are more in accord with the predictions of the traditional models. However, for Japanese data, we find that permanent income changes have effects on Japanese current account balance that are not statistically different from zero. Permanent movements in real exchange rate have statistically significant effect in both US and Japanese data. However, this result in itself neither supports nor contradicts the predictions of the intertemporal models of current account determination. It is worth noting that these results relate to long-run relationship among current account balance, income, real exchange rate, and government consumption. However, the observed long-run relationships, as revealed by the cointegration analysis, may be consistent with many different short-run dynamic adjustments among the variables. Such short-run dynamics among the variables of the model is the focus of our discussion in the next chapter.

Endnotes

1. This example is borrowed from Banerjee, Dolado, Galbraith and Hendry (1993).
2. Detailed discussion of the various methods of testing for cointegration can be found in Campbell and Perron (1991).
3. See Banerjee, Hendry and Smith (1986), and Hendry and Neale (1987) for details.
4. See Banerjee, Dolado, Galbraith and Hendry (1993) for detailed discussion on this issue.
5. For discussion of the issues involved in this method see Johansen and Juselius (1990).
6. See Osterwald-Lenum (1992) for more extensive estimates of the test statistics under various model specification.
7. This strategy of using relative income and relative government consumption imposes the restriction that the coefficients of the home country and rest of the world (for both income and government consumption) are equal in magnitude but opposite in sign — an empirically testable assumption. It is used only because of data limitation.
8. The Ljung-Box Q-test was performed with up to 16 lags. The conclusions from the test are the same for all lags. Therefore, in the text, results are presented for lags 4 (Q(4)) and 16 (Q(16)).

CHAPTER 7. SHORT RUN DYNAMICS: IMPULSE RESPONSE AND INNOVATION ACCOUNTING

Error Correction

The results pertaining to relationships among current account balance, permanent components in income and real exchange rate, and government consumption reported in chapter 6 are in essence long-run relationships. As we have discussed earlier, these are equilibrium relations that tie the variables of the system together so that the deviations from such relations do not grow indefinitely. Existence of an equilibrium relation does not imply that the relation should hold exactly at any particular point in time. Rather, the system may deviate from the equilibrium path most of the time. However, the variables of the system should adjust in response to deviations from the equilibrium path (which holds only as a long-run relation). A long-run relationship is compatible with many short-run adjustment processes. This chapter discusses the short-run dynamics of the system where transitory and permanent components of income and real exchange rate are considered explicitly. The methodology used to investigate the short-run dynamics is the conventional innovation accounting and impulse response functions.

In estimating the impulse response functions and in obtaining the variance decomposition, the cointegrating relationships among the $I(1)$ components of the variables are explicitly taken into account. The entire system of equations is estimated with restriction imposed on the cointegrating vectors, where the imposed restrictions are

those that could not be rejected (those restrictions are discussed in chapter 6). The VAR system estimated is as follows:

$$Z_t = \mu + A(L)Z_{t-1} + \hat{\alpha}\hat{\beta}'\hat{\omega}_{t-k} + \varepsilon_t$$

where $Z_t = [\Delta CA_t \quad \Delta Y_{pr,t} \quad Y_{tr,t} \quad \Delta R_{pr,t} \quad R_{tr,t} \quad \Delta G_t]'$, $A(L)$ is a 6×6 matrix of k^{th} order polynomial in the lag operator L , $\hat{\beta}'$ is the estimated cointegrating vector, $\hat{\alpha}$ is the corresponding estimated adjustment vectors, and $\hat{\omega}$ is the estimated vector of cointegrating residuals. However, instead of a full VAR with error correction, a near-VAR system with error correction is estimated. The only difference between the system estimated here and a full VAR with error correction is that in the estimated system relative government consumption is assumed to be an exogenous variable, i.e., relative government consumption affects, but is not affected by, other variables in the system. Therefore, other variables of the system do not enter in the equation for relative government consumption. The above system of near-VAR is estimated and then used to investigate the short-run dynamic behavior of the variables.

Since the coefficients of the variables in a VAR system do not have any meaningful interpretation, they are not reported here. The only coefficients worthy of mention are those of the cointegrating residuals. So, instead of all the coefficients, those of the error correction terms in various equations are presented. It may be recalled that the coefficient of an error correction term may be interpreted as speed of adjustment by the respective dependent variable in response to deviations from long-run equilibrium.

Adjustment to Long-run Equilibrium: The United States

The estimated coefficients of the error correction terms from the Johansen procedure (reported in chapter 6) for US data are presented in Table 7.1. In the table the variables e_1 , e_2 , and e_3 are the three cointegrating residuals corresponding to the three cointegration vectors reported in the previous chapter. Under either decomposition method, the coefficient of current account balance is significant in two of the three cases. When data from the Beveridge-Nelson decomposition is used, the coefficient of the error

Table 7.1. Coefficients of Error Correction Terms: US Data

Equation	<u>B-N decomposition</u>			<u>B-Q decomposition</u>		
	e_1	e_2	e_3	e_1	e_2	e_3
ΔCA	-0.0975 (-3.07)	-0.0230 (-0.34)	0.59 (4.05)	-0.0435 (-1.56)	-0.0788 (-1.93)	0.5967 (2.99)
ΔY_{pr}	0.0109 (1.69)	0.0351 (2.49)	0.0298 (1.02)	0.0049 (0.70)	0.0163 (1.63)	0.0284 (0.60)
Y_{tr}	0.0046 (1.45)	0.0106 (1.57)	0.0080 (0.58)	-0.0003 (-0.624)	0.0013 (1.84)	-0.0025 (-0.72)
ΔR_{pr}	-0.0369 (-2.18)	0.0456 (1.27)	0.0209 (0.28)	-0.0383 (-1.88)	0.0507 (1.73)	-0.2812 (-2.00)
R_{tr}	-0.0117 (-2.24)	0.014 (1.27)	0.0085 (0.37)	0.0025 (0.71)	0.0010 (2.20)	-0.0242 (-1.01)

Notes: (i). Figures in parentheses denote t-ratios of the estimated coefficients

(ii). e_1 , e_2 , and e_3 are the residuals from the three cointegrating relations found in US data.

correction term corresponding the permanent income is significant in only one relation, while with data from the Blanchard-Quah decomposition all of these coefficients are insignificant. All coefficients of error correction terms corresponding to transitory income are statistically insignificant implying negligible adjustment by transitory income in response to deviations from long-run equilibrium. Permanent component of real exchange rate has one significant adjustment coefficient under each decomposition. The same is true about the transitory component in real exchange rate. Thus, under both decomposition methods, the current account adjusts in most cases at a faster rate while income seems to respond the least in response to deviations from long-run equilibrium.

Adjustment to Long-run Equilibrium: Japan

The coefficients of the error correction terms in the estimated VAR for Japan are presented in Table 7.2. We can observe from the table that with B-N decomposition, the coefficients of transitory components in income and real exchange rate are statistically significant at 5% level. The speed of adjustment in response to deviations from long-run equilibrium is higher for transitory real exchange rate than any other variable. The adjustment speed for current account balance is significant at 10% level. On the other hand, permanent components in income and real exchange rate have statistically insignificant speeds of adjustments. With B-Q decomposition, we observe that only current account balance has statistically significant coefficient implying significant speed of adjustment. All other variables have coefficients that are not statistically insignificant.

Table 7.2. Coefficients of Error Correction Term: Japanese Data

Equation	B-N decomposition	B-Q decomposition
ΔCA	-0.1127 (-1.86)	-0.0615 (-2.89)
ΔY_{pr}	-0.0118 (-0.96)	0.0025 (0.44)
Y_{tr}	0.0146 (2.02)	0.0009 (0.80)
ΔR_{pr}	0.0235 (0.36)	0.0127 (0.61)
R_{tr}	-0.1239 (-2.42)	0.0010 (0.25)

Note: Figures in parentheses denote t-ratios.

Impulse Response Function

Impulse response functions show the dynamic responses of the variables in the VAR system to various shocks. Essentially, these functions are obtained via the vector moving average representation (VMA) of the estimated VAR system. This is an alternative representation of the VAR system where the variables of the system are expressed in terms of current and past values of the shocks to different variables. Consequently, as discussed by Sims (1980), the VMA representation allows us to trace out the time paths of the effects of various shocks on the variables in the system. Using this feature of the VMA representation, we can investigate the responses of different variables to a shock to one particular variable. The responses of different variables to an initial shock to a variable are called the impulse response functions.

The practical problem involved in obtaining the impulse response functions is the fact that the coefficients of the VMA representation are not known and can not be estimated directly. They are obtained by estimating and then inverting the VAR system. Since an estimated VAR is under-identified, the VMA representation can be obtained only by imposing some additional structure on the system. There is more than one way of imposing such structure. The Choleski decomposition method has been used in this study to obtain the impulse response functions and variance decompositions (innovation accounting). Two different orderings of the variables are used in the analysis. They are:

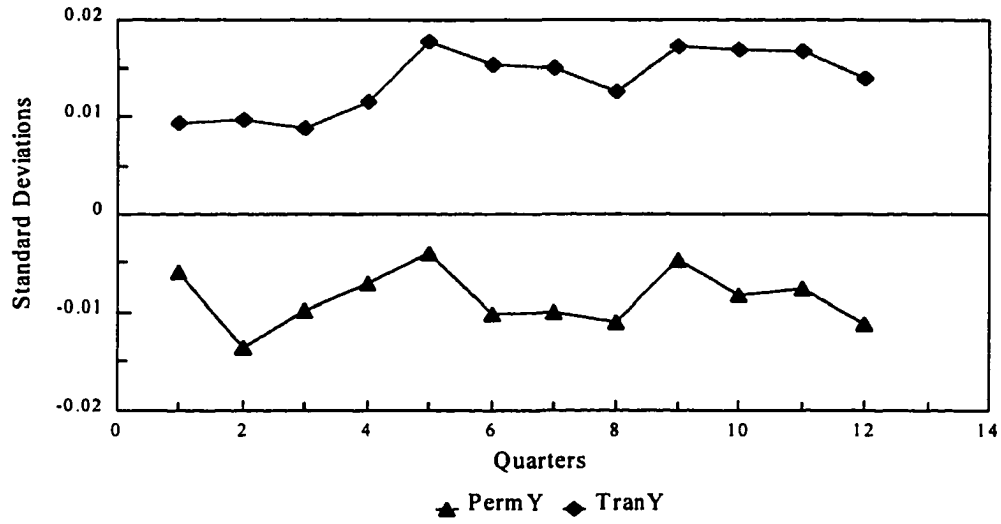
1. **Ordering 1:** Government consumption → permanent income → transitory income → permanent real exchange rate → transitory real exchange rate → current account balance.
2. **Ordering 2:** Government consumption → transitory income → permanent income → transitory real exchange → permanent real exchange rate → current account balance.

The two orderings above are variations of the ordering: *government consumption* → *income* → *real exchange rate* → *current account balance*. Such an ordering is in the tradition of the open economy macroeconomics literature. Since our interest in this study is in the current account balance, only impulse responses of current account balance are presented. Instead of presenting the responses of the first difference, the impulse response functions are plotted in terms of the level of the current account balance.

Impulse Response of Current Account Balance: The United States

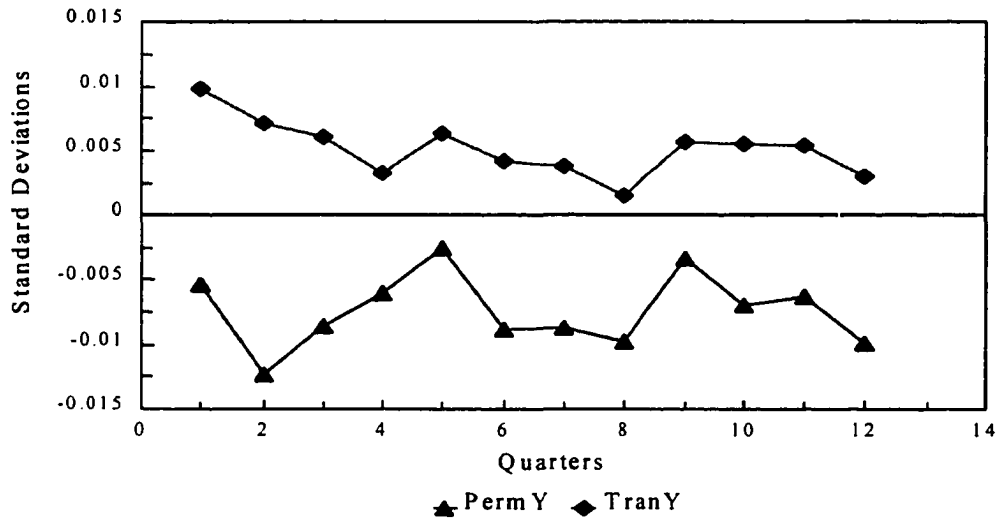
Response to Income Shocks: B-N Decomposition

Responses of US current account balance to transitory and permanent income shocks, where the decomposition is obtained by the Beveridge-Nelson method, are presented in Figures 7.1 and 7.2. Figure 7.1 reflects variable ordering 1, while Figure 7.2 reflects variable ordering 2. Under ordering 1, the instantaneous response of US current account to a transitory income increase is to jump to a surplus. This surplus reaches a peak after 4 quarters. From that point onward, no systematic pattern can be identified although current account remains in surplus. The instantaneous effect of a permanent income increase is a deterioration in the current account that worsens during the second quarter. Then it improves over the next three quarters, although it remains in deficit during the 12 quarters over which the impulse response functions are obtained. The magnitudes of the responses to the two types of income shocks are not very different. When variable ordering 2 is used, a transitory income shock leads to a current account surplus instantaneously which declines steadily over the next three quarters. After that, we do not observe any systematic behavior in the response current account even though it remains in surplus. On the other hand, the response of the current account balance to a permanent income shock under ordering 2 is similar to that under ordering 1. Such an income shock leads to current account deficit instantaneously that worsens in the next quarter before improving over the next three quarters. After that, fluctuations in current account balance show irregular pattern while remaining in deficit.



The graph represents the level of current account balance

Figure 7.1 Responses to Income Shocks: Variable Ordering 1



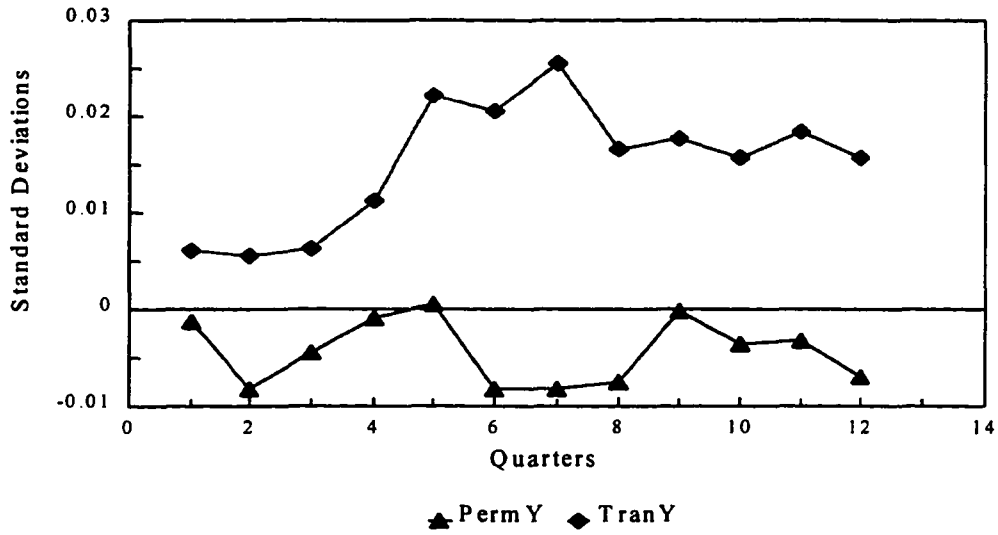
The graph represent the level of current account balance

Figure 7.2 Responses to Income Shocks: Variable Ordering 2

The magnitudes of the responses of current account to the two types of income shocks are very similar to those under variable ordering 1. On the one hand, we see that under both orderings a transitory income shock leads to a current account surplus which is consistent with the intertemporal models. However, one feature of the impulse response functions is that transitory and permanent income shocks affect the current account balance in the opposite direction: the former leads to a surplus while the latter leads to a deficit in the current account.

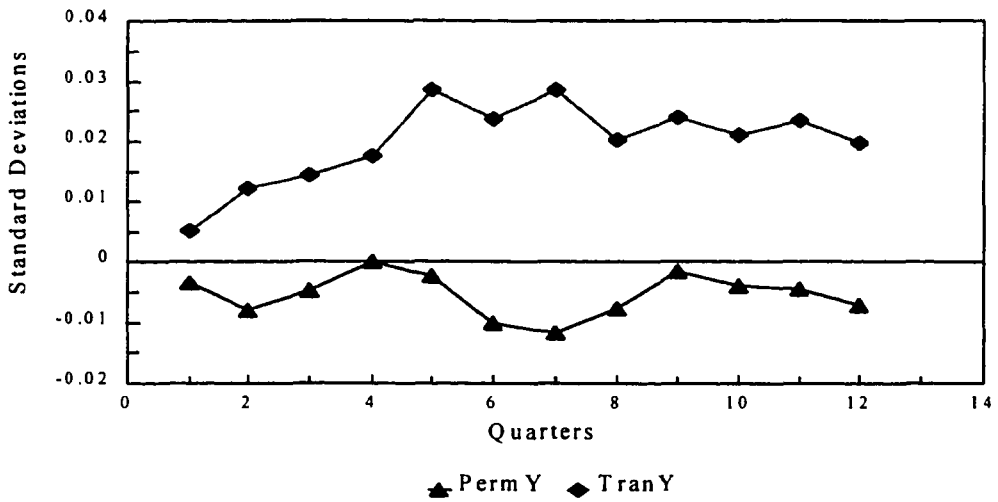
Response to Income Shocks: B-Q Decomposition

Responses of US current account balance to the two types of income shocks, where the decomposition of income is obtained by using the Blanchard-Quah method, are plotted in Figure 7.3 under variable ordering 1, and in Figure 7.4 under variable ordering 2. In these two diagrams, we can observe that the two orderings generate very similar responses of the current account balance to the two types of income shocks. Under both variable orderings, the instantaneous response of the current account balance to a transitory income increase is to jump to surplus that improves over the next 4 quarters. From that point onward, current account balance fluctuates randomly while it continues to remain in surplus. On the other hand, the instantaneous response of current account to a permanent income shock is for the former to move to deficit that worsens in the next quarter. From quarter three, the deficit starts improving which continues for the next three quarters. From that point onward, no systematic pattern can be detected. The magnitudes of the responses to the two types of income shocks reflect a relatively stronger effect of



The graph represents the level of current account balance

Figure 7.3 Responses to Income Shocks: Variable Ordering 1



The graph represents the level of current account balance

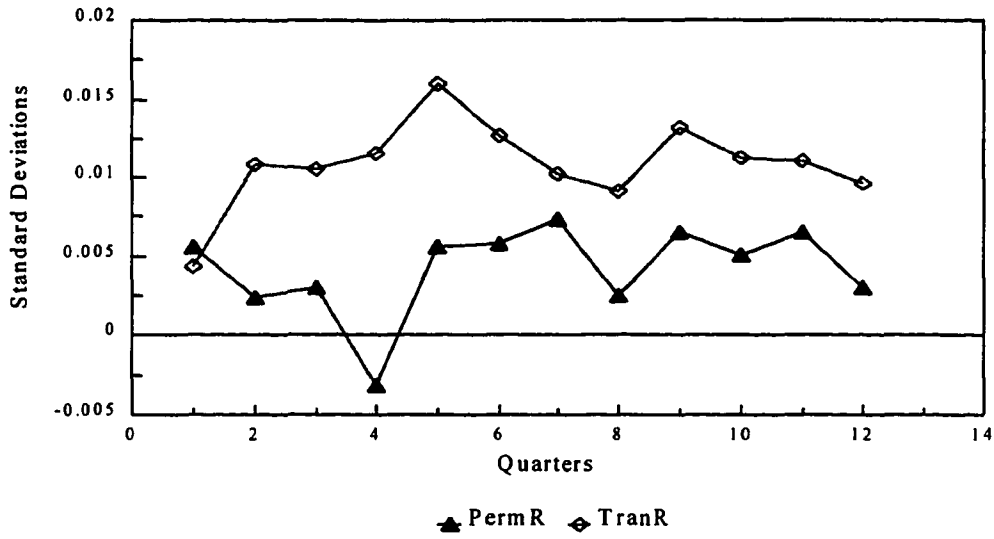
Figure 7.4 Responses to Income Shocks: Variable Ordering 2

transitory income than permanent income on US current account balance. Thus, with data from the Blanchard-Quah decomposition, we see some support (though weak) for the intertemporal models (of current account determination) in the response of US current account balance to the two types of income shocks.

Response to Real Exchange Rate Shock: B-N Decomposition

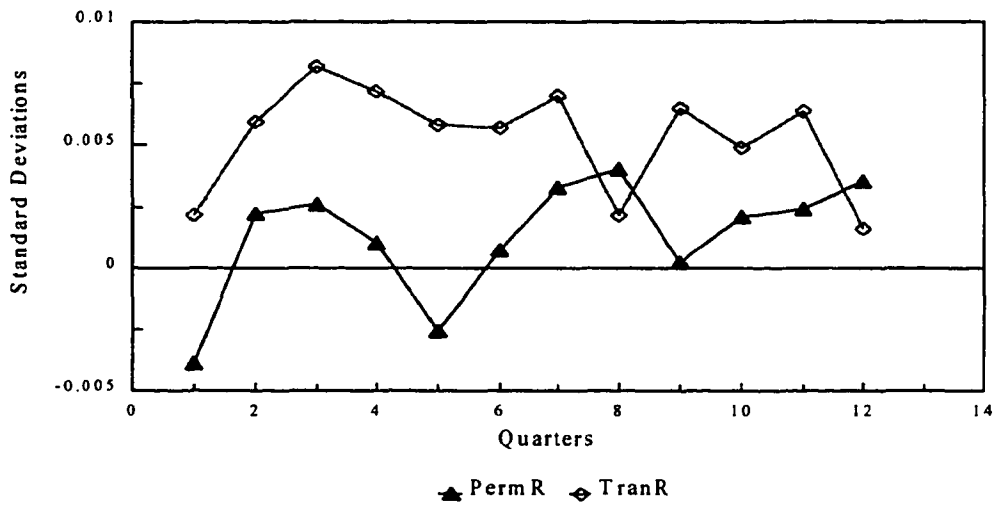
The responses of US current account balance to transitory and permanent real exchange rate shocks using B-N decomposition are plotted in Figures 7.5 and 7.6, using variable orderings 1 and 2, respectively. It is evident from Figure 7.5 that under variable ordering 1 a transitory real exchange rate shock (deterioration) leads to instantaneous surplus in the current account balance which improves over the next 4 quarters. Then the surplus starts shrinking. Similarly, the instantaneous effect of a permanent real exchange rate shock on the current account positive. However, after the first period, the surplus in the current account falls. By period 4, the current account balance goes to deficit. From that point onward, the current account balance fluctuates without showing any systematic pattern.

From Figure 7.6 we observe that if variable ordering 2 is used, a transitory deterioration in US real exchange rate leads to surplus in the current account that grows moderately during the first three quarters. The surplus then shrinks over the next three quarters. From quarter six, it starts fluctuating without showing any systematic pattern. The instantaneous effect of a permanent real exchange rate shock (increase in world relative price) is to generate a current account deficit that moves back to surplus in the



The graph represents the level of current account balance

Figure 7.5 Responses to Real Exchange Rate Shocks: Variable Ordering 1



The graph represents the level of current account balance

Figure 7.6 Responses to Real Exchange Rate Shocks: Variable Ordering 2

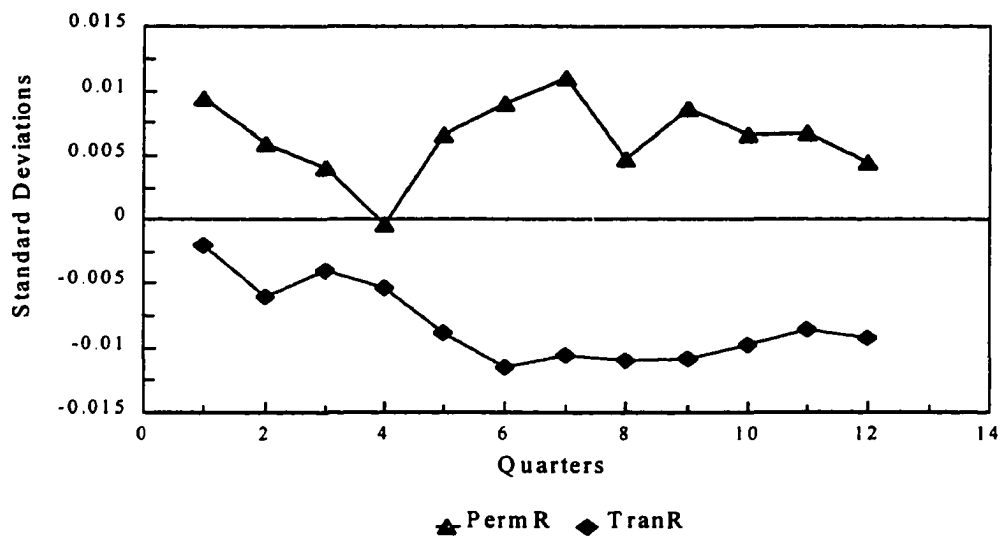
very next quarter. The general pattern of this response function does not suggest a systematic effect of a permanent real exchange rate shock on US current account balance. Further, we can observe that the magnitudes of the effects of a transitory shocks are larger than those of a permanent shock to real exchange rate.

Response to Real Exchange Rate Shock: B-Q Decomposition

Responses of US current account to transitory and permanent real exchange rate shocks, where the decomposition is obtained by B-Q method, are presented in Figures 7.7 and 7.8, using variable orderings 1 and 2, respectively. Figure 7.7 uses variable ordering 1, and shows that a transitory real exchange rate deterioration leads to current account deficit instantaneously that worsens over the next 6 quarters before it flattens out. A permanent real exchange rate shock (increase in world relative price) leads to an instantaneous improvement in the current account that declines over the next three quarters. In general, current account movements do not show any systematic response.

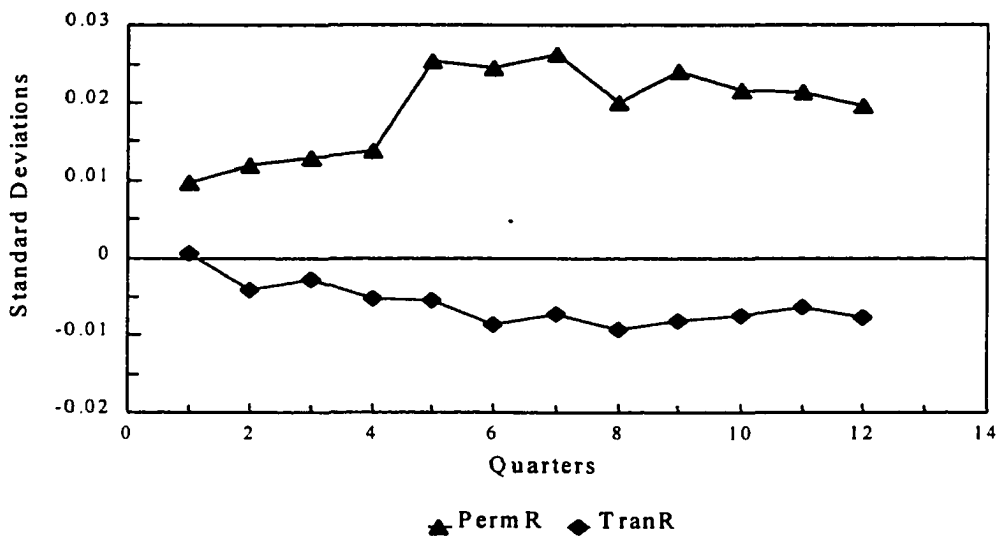
When variable ordering 2 is used in obtaining the impulse response functions, the response of current account to a transitory real exchange rate shock is very similar to that under ordering 1: transitory increase in world relative price leads to a current account deficit while a permanent increase in world relative price leads to current account surplus. The initial surplus in current account grows over the next four quarters. After that the movements in the current account balance do not reveal any systematic pattern although it continues to remain in the surplus.

Response to Own and Government



The graph represents the level of current account balance

Figure 7.7 Responses to Real Exchange Rate Shocks: Variable Ordering 1

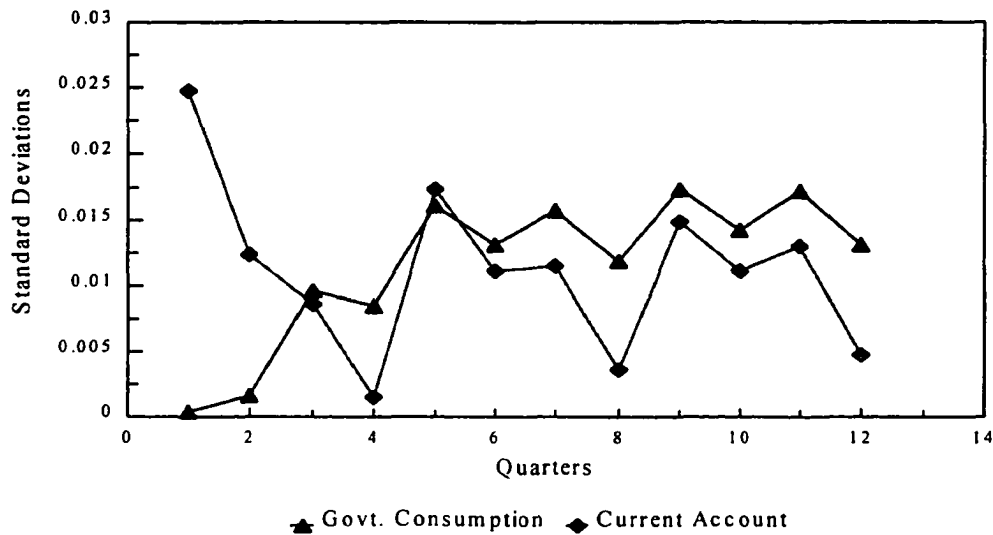


The graph represents the level of current account balance

Figure 7.8 Responses to Real Exchange Rate Shocks: Variable Ordering 2

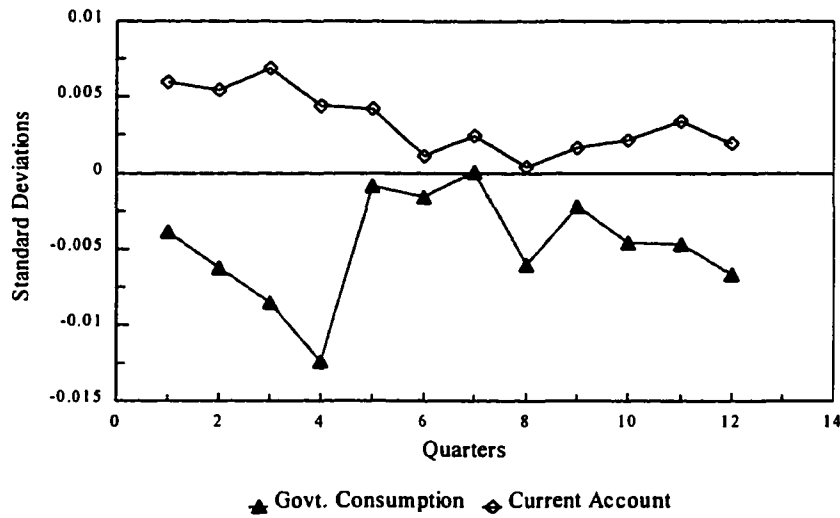
Consumption Shock

Responses of US current account balance to its own shock and shocks in government consumption are plotted in Figures 7.9 and 7.10, the former using the B-N decomposition and the latter using the B-Q decomposition. It appears from these two diagrams that government consumption shock has some positive effect on current account over first few quarters if data from the B-N decomposition are used. If data from the B-Q decomposition are used, such a shock has negative effect on current account balance. Effects of own shock to current account balance does not show any systematic pattern under either method of decomposition.



The graph represents the level of current account balance

Figure 7.9 Responses to Own and Govt. Consumption Shocks: Variable Ordering 1



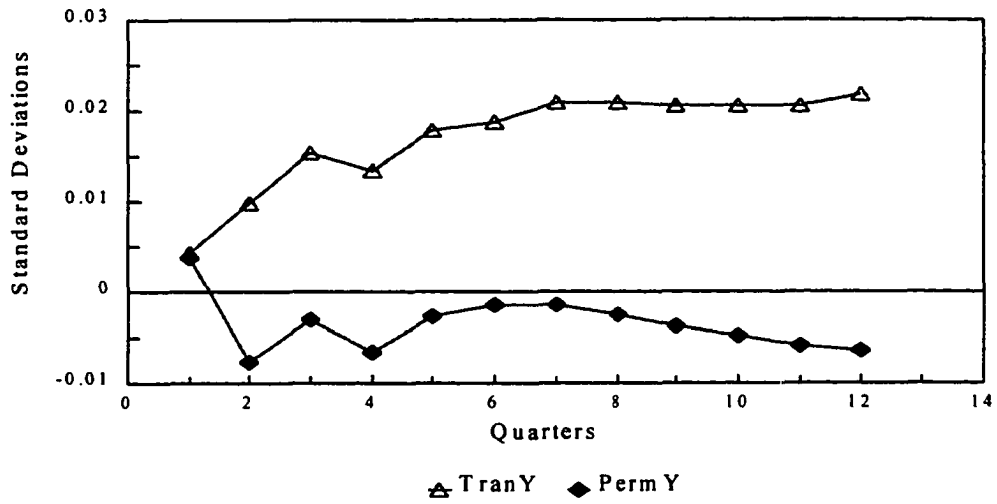
The graph represents the level of current account balance

Figure 7.10 Responses to Own and Govt. Consumption Shocks: Variable Ordering 2

Impulse Response of Current Account Balance: Japan

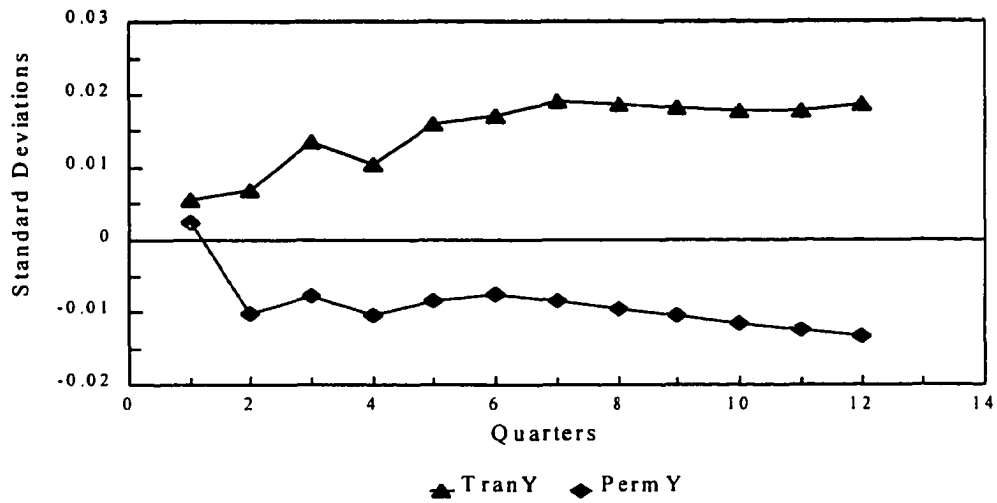
Response to Income Shock: B-N Decomposition

Responses of Japanese current account balance to transitory and permanent income shocks with data from the B-N decomposition are shown in Figure 7.11 and 7.12. The former is obtained under variable orderings 1, while the latter uses the variable ordering 2. In Figure 7.11 we observe that a transitory income increase generates a current account surplus that grows over the first few quarters. From quarter 6 onwards the growth in the surplus comes to an end and current account remains stable over the remaining periods for which the impulse response functions are obtained. A permanent



The graph represents the level of current account balance

Figure 7.11 Responses to Income Shocks: Variable Ordering 1



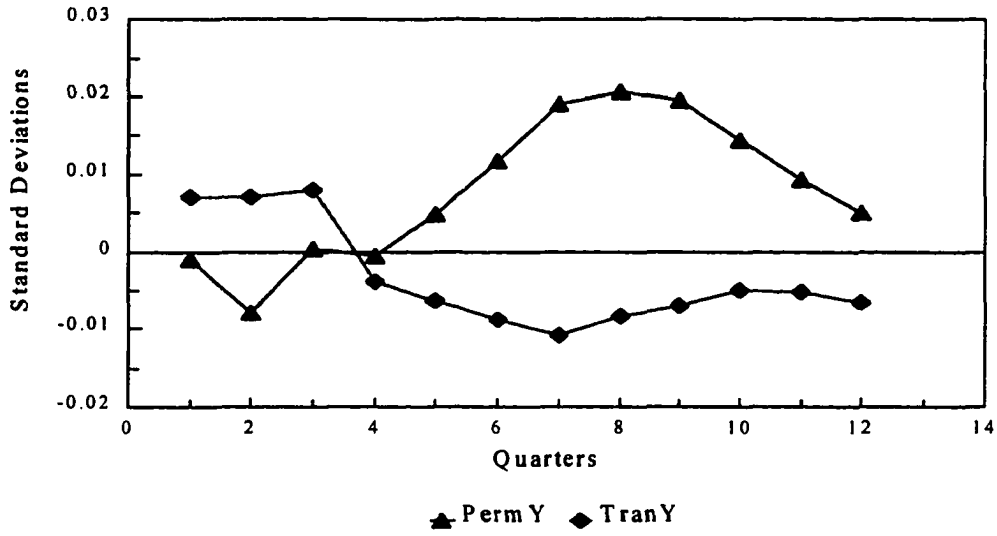
The graph represents the level of current account balance

Figure 7.12 Responses to Income Shocks: Variable Ordering 2

income increase, on the other hand, does not seem to have a strong effect on current account balance. After generating a surplus instantaneously, it goes to deficit in the second quarter. The deficit tends to shrink as time passes. Looking at Figure 7.12 we observe that the response of current account to a transitory income shock under variable ordering 2 is very similar to that under variable ordering 1. A permanent income shock initially generates a small surplus in the current account balance. In the next quarter, the surplus disappears and current account moves to deficit. In the third quarter, the deficit shrinks. Thus, the movements in current account balance does not suggest any systematic pattern. The magnitudes of the responses to transitory and permanent income shocks suggest a stronger effect of the former than the latter on Japanese current account balance. Thus the responses of current account to the two types of income shocks are insensitive to variable orderings, and are broadly supportive of the predictions of the intertemporal models of current account determination.

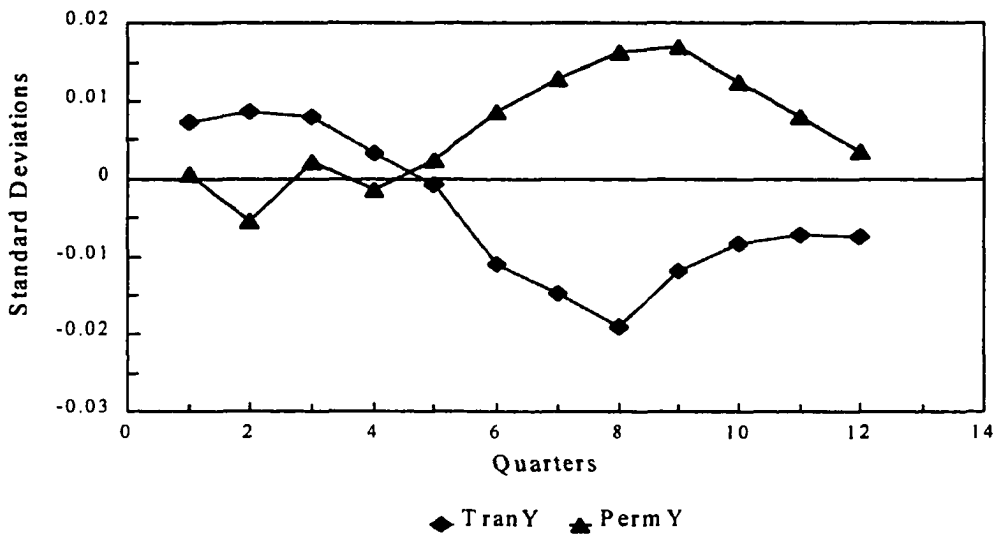
Response to Income Shock: B-Q Decomposition

Impulse response (to transitory and permanent income shocks) functions using data from the B-Q decomposition are presented in Figures 7.13 and 7.14, under variable orderings 1 and 2, respectively. From these two figures, we can observe that when data from the B-Q decomposition are used, under either variable ordering, a transitory income shock initially leads to current account surplus that remains steady for three quarters. Then the surplus starts disappearing and the current account goes to deficit. A permanent income shock, on the other hand, initially does not show any systematic effect on current



The graph represents the level of current account balance

Figure 7.13 Responses to Income Shocks: Variable Ordering 1



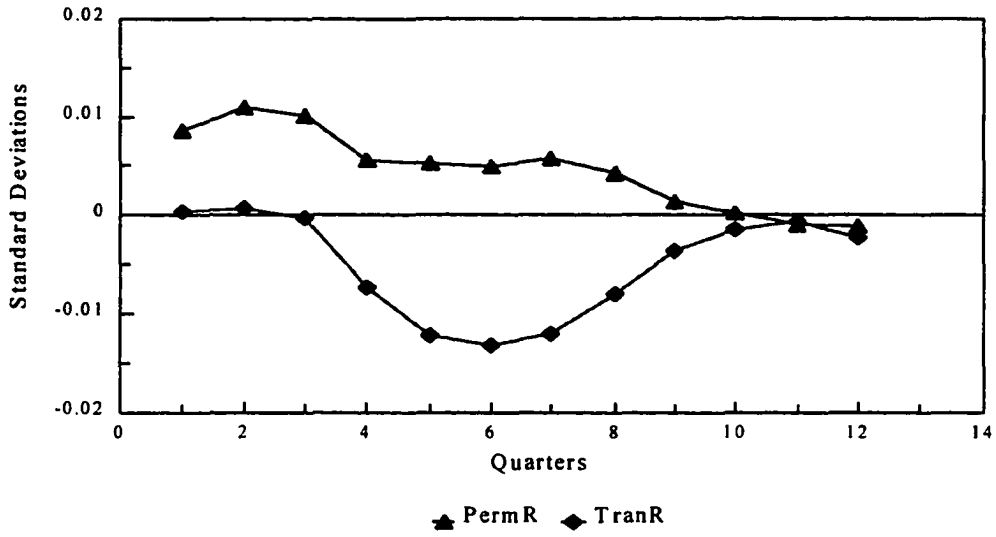
The graph represents the level of current account balance

Figure 7.14 Responses to Income Shocks: Variable Ordering 2

account balance. However, after 4 quarters the current account moves to surplus that reaches a peak around quarter 8 and then it starts declining steadily. At the end of quarter 12 the surplus in current account balance induced by a permanent income increase virtually disappears. It may be recalled that in chapter 6, statistical test suggested no long-run effect of permanent income changes on current account balance. The impulse response functions in Figures 7.13 and 7.14 reinforce the results of those statistical tests. Thus the responses of Japanese current account to transitory and permanent income changes are in more supportive of the intertemporal models of current balance than US data.

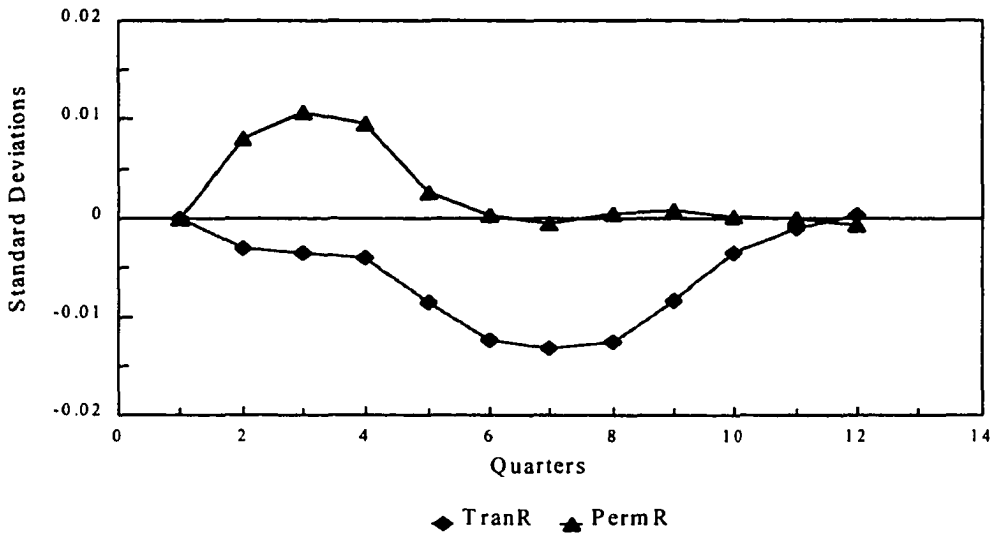
Response to Real Exchange Rate Shock: B-N Decomposition

Figures 7.15 and 7.16 show the responses of Japanese current account to transitory and permanent real exchange rate shocks, where the decomposition is obtained by the B-N method. Figure 7.15 reflects variable ordering 1 while Figure 7.16 reflects ordering 2. These two figures show that under either variable ordering a transitory real exchange rate shock initially has little or no effect on the current account balance. However after 3 to 4 quarters, such a shock leads to current account deficit that reaches its lowest around quarter 6. From that point onwards it starts to shrink and by quarter 10 to 11 the deficit disappears. Thus, irrespective of variable ordering, data from B-N decomposition shows that a transitory real exchange rate shock has some effect (although with some lag) on Japanese current account balance. On the other hand, under both variable orderings, a permanent real exchange rate shock leads to an initial improvement



The graph represents the level of current account balance

Figure 7.15 Responses to Real Exchange Rate Shocks: Variable Ordering 1



The graph represents the level of current account balance

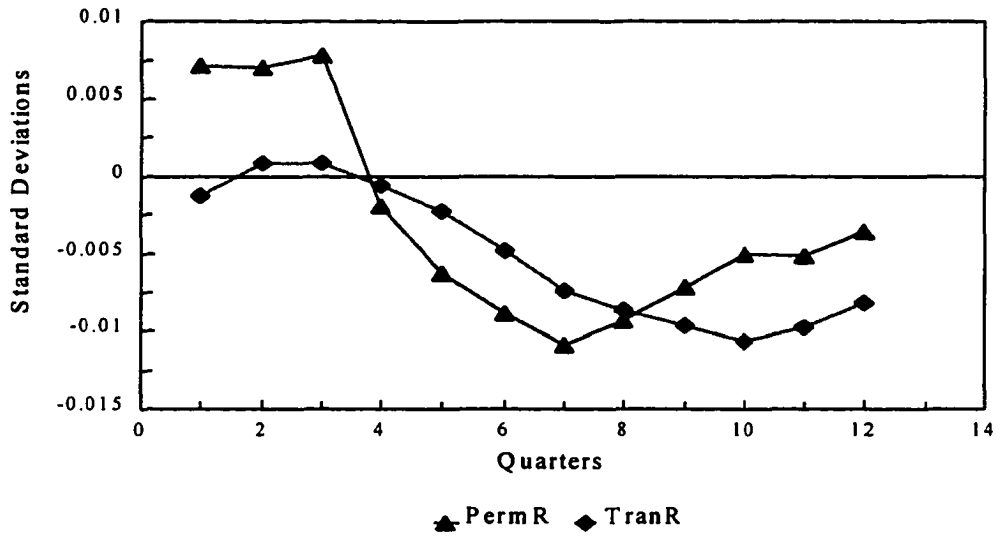
Figure 7.16 Responses to Real Exchange Rate Shocks: Variable Ordering 2

in the current account balance. At first the surplus in the current account increases; however, after 3 to 4 quarters it begins to diminish systematically and continuously. The decline in the surplus is faster under variable ordering 2 than under variable ordering 1.

Response to Real Exchange Rate Shock: B-Q Decomposition

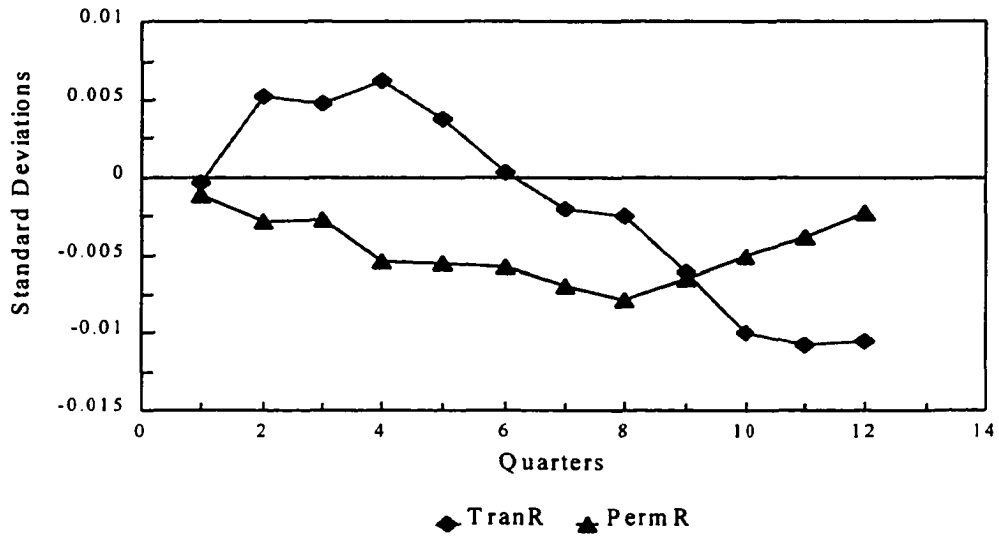
Responses of Japanese current account to transitory and permanent real exchange rate shocks with data obtained from the B-Q decomposition are presented in Figures 7.17 and 7.18. The former corresponds to variable ordering 1, while the latter reflects variable ordering 2. From Figure 7.17 we see that, under variable ordering 1, initially a transitory real exchange rate shock does not show any effect for 4 quarters. Then it lead to deficit in current account which grows steadily until quarter 10. Then the deficit starts shrinking. Under variable ordering 2, a transitory real exchange rate shock initially leads to a current account surplus for first four quarters. Then the surplus starts shrinking, and by quarter 6 the initial surplus totally disappears. The current account then goes to deficit that grows until quarter 10. Then the deficit shows a tendency shrink.

When variable ordering 1 is used, a permanent real exchange depreciation initially generates a current account surplus. Then from the 4th quarter onwards, current account goes into deficit. This deficit worsens for the next four quarters before it starts to shrink. On the other hand, under variable ordering 2, a permanent real exchange rate depreciation leads to a deficit in the current account that worsens over 8 quarters before showing signs of improvement. Thus, we see quite different short-run response of the current account balance to a real exchange rate shock under the two decomposition methods. With data



The graph represents the level of current account balance

Figure 7.17 Responses to Real Exchange Rate Shocks: Variable Ordering 1



The graph represents the level of current account balance

Figure 7.18 Responses to Real Exchange Rate Shocks: Variable Ordering 2

from the B-N decomposition, a permanent real exchange rate increase leads to temporary surplus in the current account, but does not induce any deficit. If data obtained from the B-Q decomposition are used, a permanent real exchange rate deterioration may or may not generate any surplus in current account. However, after some lag it generates deficit that persists for quite some time. Similarly, if data from the B-N decomposition are used, a transitory real exchange rate deterioration necessarily generates current account deficit. On the contrary, with B-Q decomposition, such a shock generates a temporary surplus before inducing any deficit when variable ordering 2 is used. These results, thus, do not lead to unequivocal conclusions.

Responses to Own and Government Consumption Shocks

Figure 7.19 shows the response of Japanese current account to own shock and shock to relative government consumption where the data from the B-N decomposition is used. It shows that when the B-N decomposition is used, neither government consumption shock nor current account's own shock has any systematic effect on current account balance. A government consumption shock almost has no effect on Japanese current account balance. Figure 7.20 shows the response to Japanese current account balance to the same two shocks with data from the the Blanchard-Quah decomposition. It is evident from Figure 7.20 that when data from B-Q decomposition are used, government consumption shock seems to be associated with positive movements in current account balance of Japan. On the other hand, current account balances's own shock does not seem to have any systematic effect on Japanese current account balance.

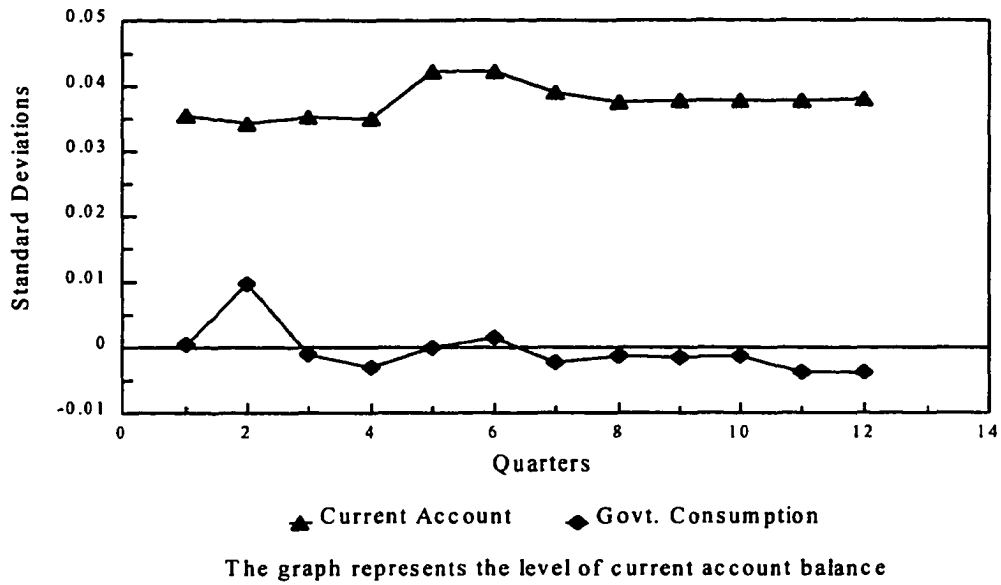


Figure 7.19 Responses to Own and Govt. Consumption Shocks: Variable Ordering 1

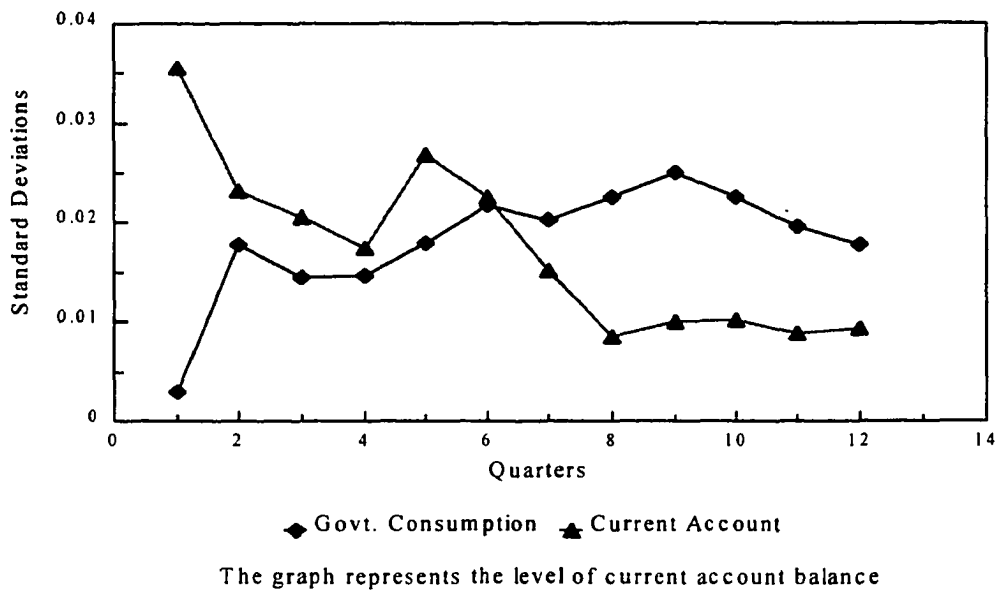


Figure 7.20 Responses to Own and Govt. Consumption Shocks: Variable Ordering 2

Variance Decomposition

In this section we present the results of variance decomposition to see if the transitory and permanent components of income and real exchange rate have markedly different contributions in explaining the forecast error variances. Variance decomposition is the technique of isolating the contributions of different variables in the estimated system in explaining k-period ahead forecast error. The results presented here are obtained by using the Choleski factorization method, where we have used both of the variable orderings mentioned earlier. It is well known that the results of such a procedure may be conditional on the ordering of the variables used. In general, if the correlation between any two innovations is rather high then the decomposition results are highly sensitive to the ordering used. However, in such a situation we may follow the guideline suggested by Litterman and Weiss (1983). If the correlation between two innovations are high, one should run a pair of decompositions with the two variables placed next to each other, but only interchanging their positions. Usually the contribution of the variable appearing in the first position tends to have a higher contribution in the decomposition result. If this is true for both orderings then we can not draw any definitive conclusion. On the other hand, if one variable does relatively better when placed in the second position than the other then this suggests that this variable has a relatively stronger causative influence than the other variable. If the variable appearing in the second position does better under both orderings, then some linear combination of the two variables is the true causative factor.

Now let us examine the results of the variance decomposition procedure. The decomposition results are presented for 4 different forecast horizons: 1, 4, 8, 12. The only exception is current account balance for which results are presented for forecast horizons 1, 2, 4, 8, 12. Variance decompositions are presented for data obtained from both decomposition methods, and for both variable orderings. For ease of understanding, the columns of the tables containing the decomposition results are arranged such that they reflect the variable ordering used. First we present the results for the US and then for Japan. For each country, first the results for data obtained from the B-N decomposition is presented first under variable ordering 1 and then under ordering 2. Then results are presented for data from the B-Q decomposition following the same sequence of orderings. The discussion focuses on the decomposition of forecast error variance of current account balance.

Forecast Error Variance Decomposition: The United States

Innovation Accounting: Beveridge-Nelson Decomposition

Variance decompositions with data from the B-N decomposition are presented in Tables 7.3 and 7.4, under variable orderings 1 and 2, respectively. It is evident from these two tables that the contributions of different variables in explaining near-term forecast error variance are not sensitive to the variable ordering used. In explaining one-period ahead forecast error variance, the contribution of transitory income is the highest: in fact, it is almost three times that of permanent income, and almost equal to the sum of the contributions of permanent income, transitory and permanent components of real

Table 7.3. Variance Decomposition: US and B-N Decomposition

		<u>Forecast Horizon</u>		<u>Innovation in</u>			
	k	G. Cons	PermY	TranY	PermR	TranR	CA
G. Cons.	All	100.00	0.00	0.00	0.00	0.00	0.00
PermY	1	0.06	99.94	0.00	0.00	0.00	0.00
	4	10.17	65.95	18.93	3.24	1.11	0.59
	8	9.37	61.93	18.36	5.62	3.24	1.49
	12	8.92	61.69	18.03	5.63	3.76	1.97
TranY	1	0.04	0.32	99.63	0.00	0.00	0.00
	4	10.51	7.71	73.74	3.02	1.72	3.30
	8	10.76	8.30	68.37	4.44	1.78	6.35
	12	10.54	8.39	66.39	4.94	2.48	7.26
PermR	1	0.00	2.28	0.45	97.27	0.00	0.00
	4	12.22	2.08	4.62	66.30	14.24	0.55
	8	12.73	4.46	7.87	60.84	13.04	1.06
	12	13.12	4.35	7.81	60.02	13.18	1.52
TranR	1	0.02	1.88	0.41	96.95	0.74	0.00
	4	13.36	1.78	4.00	64.96	15.10	0.81
	8	13.64	4.56	6.42	59.91	14.00	1.47
	12	13.94	4.56	6.38	58.97	14.02	2.13
CA	1	0.02	4.43	11.32	3.98	2.44	77.81
	2	0.18	7.68	9.34	3.98	5.78	73.03
	4	5.32	8.17	8.51	6.43	4.96	66.62
	8	7.97	7.86	8.14	9.47	5.15	61.42
	12	9.27	9.18	7.86	9.16	5.20	59.32
Correlation Matrix							
G. Cons	12	1.00					
PermY	12	-0.15	1.00				
TranY	12	-0.25	-0.12	1.00			
PermR	12	0.36	-0.46	-0.52	1.00		
TranR	12	0.64	-0.38	-0.44	0.50	1.00	
CA	12	-0.34	-0.20	-0.21	-0.38	-0.07	1.00

Note: Variance decomposition is obtained using variable ordering 1.

Table 7.4. Variance Decomposition: US and B-N Decomposition

	<u>Forecast Horizon</u>		<u>Innovation in</u>				
	k	G. Cons	TranY	PermY	TranR	PermR	CA
G. Cons	All	100.00	0.00	0.00	0.00	0.00	0.00
TranY	1	0.04	99.96	0.00	0.00	0.00	0.00
	4	10.51	74.03	7.42	2.71	2.04	3.30
	8	10.76	68.68	7.99	4.14	2.07	6.35
	12	10.54	66.69	8.09	4.69	2.73	7.26
PermY	1	0.06	0.32	99.62	0.00	0.00	0.00
	4	10.17	20.01	64.87	3.27	1.08	0.59
	8	9.37	19.61	60.67	5.58	3.27	1.49
	12	8.92	19.37	60.34	5.61	3.78	1.97
TranR	1	0.02	0.32	1.98	97.69	0.00	0.00
	4	13.36	4.03	1.75	64.23	15.83	0.81
	8	13.64	6.24	4.74	59.19	14.72	1.47
	12	13.94	6.21	4.73	58.20	14.78	2.13
PermR	1	0.00	0.34	2.39	96.53	0.73	0.00
	4	12.22	4.69	2.01	64.03	16.50	0.55
	8	12.73	7.70	4.62	58.74	15.14	1.06
	12	13.12	7.66	4.51	57.89	15.31	1.52
CA	1	0.02	12.10	3.65	4.51	1.92	77.81
	2	0.18	9.72	7.30	4.74	5.02	73.03
	4	5.32	8.93	7.75	6.95	4.43	66.62
	8	7.97	8.19	7.81	10.14	4.48	61.42
	12	9.27	7.72	9.33	9.89	4.48	59.32
Correlation Matrix							
G. Cons.	12	1.00					
TranY	12	-0.25	1.00				
PermY	12	-0.15	-0.11	1.00			
TranR	12	0.35	-0.53	-0.46	1.00		
PermR	12	0.66	-0.43	-0.38	0.53	1.00	
CA	12	-0.34	-0.21	-0.20	-0.37	-0.15	1.00

Note: Variance decomposition is obtained using variable ordering 2.

exchange rate, and government consumption. This is true for both variable orderings. So, for a very short-term forecast horizon, transitory income has the highest explanatory power. At longer-term forecast horizons importance of other variables increase: for 8 period ahead forecast error variance, contributions of permanent and transitory income are very similar. On the other hand, between transitory and permanent components of real exchange rate, the variable appearing at the first position seems to have relatively higher explanatory power. This suggests that the evidence regarding the relative performances of these two variables are inconclusive. Further, for a very short forecast horizon (one or two period ahead forecasts) contribution of relative government consumption is very low. However, for 8 to 12 period ahead forecasts, relative government consumption's contribution rises to about 10% which is very similar to those of transitory and permanent components of real income.

Innovation Accounting: Blanchard-Quah Decomposition

Variance decomposition results using data from the B-Q method are presented in Table 7.5 and Table 7.6, using variable orderings 1 and 2, respectively. These two tables show that ordering of variables has insignificant effect on variance decomposition results. Under either of the two orderings considered, contribution of the permanent component in income in one-period forecast error variance is below 1 percent, while that of transitory income is about 5 percent. For 8 period ahead forecast error variance, contribution of transitory income is about 16% (under either variable ordering) which is almost twice that of permanent income. This shows that for near-term forecasting transitory income

Table 7.5. Variance Decomposition: US and B-Q Decomposition

	<u>Forecast Horizon</u>		<u>Innovation in</u>				
	k	G. Cons	PermY	TranY	PermR	TranR	CA
G. Cons	All	100.00	0.00	0.00	0.00	0.00	0.00
PermY	1	0.27	99.73	0.00	0.00	0.00	0.00
	4	9.83	84.36	0.88	1.23	2.92	0.77
	8	9.15	79.58	2.07	3.19	4.12	1.88
	12	9.81	77.85	2.25	3.54	4.30	2.24
TranY	1	1.09	12.63	86.28	0.00	0.00	0.00
	4	13.27	1.43	64.70	9.94	7.30	3.36
	8	17.00	6.56	54.92	9.16	4.74	7.62
	12	16.41	7.73	53.34	9.04	5.64	7.83
PermR	1	0.89	0.23	82.76	16.12	0.00	0.00
	4	21.31	0.66	60.36	11.41	3.84	2.41
	8	19.87	6.07	53.32	12.17	3.28	5.30
	12	20.52	6.25	52.58	11.79	3.28	5.57
TranR	1	0.55	36.25	23.39	2.98	36.84	0.00
	4	3.09	21.85	42.50	1.80	22.55	8.21
	8	4.46	29.20	36.69	2.61	20.07	6.98
	12	6.32	27.56	36.67	2.83	19.45	7.17
CA	1	0.03	0.24	5.19	12.76	0.60	81.17
	2	0.04	5.15	6.28	10.67	2.11	75.76
	4	6.01	6.46	7.27	10.67	2.24	67.35
	8	9.76	7.92	15.84	14.75	2.31	49.43
	12	10.11	10.81	15.36	14.66	2.26	46.78
Correlation Matrix							
G. Cons	12	1.00					
PermY	12	-0.18	1.00				
TranY	12	0.25	-0.64	1.00			
PermR	12	0.31	-0.73	0.34	1.00		
TranR	12	-0.38	-0.03	0.11	-0.34	1.00	
CA	12	-0.24	-0.34	-0.45	0.38	-0.28	1.00

Note: Variance decomposition is obtained using variable ordering 1.

Table 7.6. Variance Decomposition: US and B-Q Decomposition

	<u>Forecast Horizon</u>		<u>Innovation in</u>				
	k	G. Cons	TranY	PermY	TranR	PermR	CA
G. Cons	All	100.00	0.00	0.00	0.00	0.00	0.00
TranY	1	1.09	98.91	0.00	0.00	0.00	0.00
	4	13.27	54.61	11.51	11.83	5.41	3.36
	8	17.00	47.14	14.35	7.96	5.94	7.62
	12	16.41	45.59	15.48	8.85	5.83	7.83
PermY	1	0.27	12.73	86.99	0.00	0.00	0.00
	4	9.83	9.54	75.70	2.00	2.15	0.77
	8	9.15	11.54	70.12	2.48	4.83	1.88
	12	9.81	11.74	68.37	2.55	5.29	2.24
TranR	1	0.55	44.47	15.17	39.82	0.00	0.00
	4	3.09	56.45	7.91	22.20	2.15	8.21
	8	4.46	48.98	16.91	20.40	2.28	6.98
	12	6.32	47.08	17.14	19.96	2.32	7.17
PermR	1	0.89	69.34	13.65	1.21	14.91	0.00
	4	21.31	51.53	9.50	4.41	10.84	2.41
	8	19.87	43.08	16.31	4.23	11.21	5.30
	12	20.52	42.64	16.19	4.22	10.85	5.57
CA	1	0.03	3.81	1.62	0.06	13.31	81.17
	2	0.04	7.94	3.49	2.48	10.30	75.76
	4	6.01	7.85	5.88	2.75	10.16	67.35
	8	9.76	16.11	7.65	2.42	14.64	49.43
	12	10.11	16.77	9.41	2.44	14.48	46.78
Correlation Matrix							
G. Cons	12	1.00					
TranY	12	0.02	1.00				
PermY	12	0.02	-0.51	1.00			
TranR	12	-0.35	0.44	-0.32	1.00		
PermR	12	0.35	-0.17	-0.44	-0.46	1.00	
CA	12	-0.24	-0.56	-0.36	-0.30	0.52	1.00

Note: Variance decomposition is obtained using variable ordering 1.

definitely has greater explanatory power than permanent income. On the other hand, the reverse is true about the relative contributions of permanent and transitory components in real exchange rate. At one-period ahead forecast horizon, permanent component of real exchange rate contributes about 13% of forecast error variance which is greater than the sum of the contributions of permanent and transitory income, and relative government consumption. Contribution of transitory real exchange rate is about 20 percent of that of permanent real exchange rate. The same pattern continues at longer higher forecast horizons. At very short forecast horizons, relative contribution of government consumption are is low. However, in longer term forecasts (e.g. 8-period forecast horizons) government consumption's share rises to about 10 percent. Thus, transitory income, and permanent real exchange rate are more important as explanatory variables in near-term forecasting of current account balance. On the other hand, the importance of government consumption in explaining the forecast error variance of current account balance increases with the length of the forecast horizon.

Forecast Error Variance Decomposition: Japan

Innovation Accounting: Beveridge-Nelson Decomposition

Variance decomposition results using data from the Beveridge-Nelson decomposition are presented in Table 7.7 and Table 7.8, using the variable orderings 1 and 2, respectively. It is evident from these two tables that the ordering of the transitory and permanent components of income and real exchange rate has very little effect on the relative contributions of the different variables of the model. Under both orderings,

permanent component of income contributes more in explaining forecast error variances at horizons 2 to 12 period ahead. At 2-period ahead forecast horizon, permanent income contributes between 9 to 10 percent of the forecast error variance. The contribution of transitory income is between 2 to 3 percent for the same forecast horizon and this remains under 6 percent for 12-period ahead forecast horizon.

Looking at the shares of transitory and permanent components of real exchange rate, we observe that at very short-term forecast horizon (one to two period ahead forecast) transitory real exchange rate contributes very little to explain the forecast error variance. At similar forecast horizon, permanent real exchange rate explains about 5 percent of the forecast error variance. At 8 to 12 period ahead forecast horizons contributions of transitory and permanent components are very similar, each explaining about 5 percent of the forecast error variance.

Looking at the same table we see that relative government consumption seems to have the maximum explanatory power over 4 to 12 period ahead forecast horizons. In particular, its contribution is higher than any other variable in explaining the forecast error variance at any horizon between 4 to 12 periods ahead . At shorter forecast horizons (for instance, at one to two period ahead forecast horizons), however, it is dominated by both permanent real income and permanent component in real exchange rate. So, at shorter forecast horizon, permanent components in income and real exchange rate seem to more important than other variables, while at longer forecast horizon, relative government consumption has the highest explanatory power.

Table 7.7. Variance Decomposition: Japan and B-N Decomposition

	<u>Forecast Horizon</u>		<u>Innovation in</u>				
	k	G. Cons	PermY	TranY	PermR	TranR	CA
G. Cons	All	100.00	0.00	0.00	0.00	0.00	0.00
PermY	1	0.16	99.84	0.00	0.00	0.00	0.00
	4	7.06	57.72	26.71	1.81	2.81	3.88
	8	11.94	53.12	25.01	1.73	3.58	4.63
	12	12.58	52.17	24.80	1.81	4.05	4.59
TranY	1	0.92	10.97	88.11	0.00	0.00	0.00
	4	2.60	10.53	81.95	2.28	1.98	0.66
	8	11.69	8.61	69.96	3.54	5.35	0.85
	12	12.62	8.37	69.07	3.45	5.66	0.83
PermR	1	3.52	0.14	0.28	96.07	0.00	0.00
	4	10.86	0.97	1.91	70.09	10.75	5.42
	8	12.20	1.20	5.33	60.11	14.57	6.60
	12	12.82	1.49	5.24	56.57	17.55	6.34
TranR	1	0.48	2.38	5.41	16.53	75.20	0.00
	4	1.36	6.94	6.63	19.14	64.50	1.43
	8	3.64	6.81	14.24	14.35	59.55	1.41
	12	3.95	6.73	11.84	13.74	62.35	1.38
CA	1	0.02	1.04	1.39	5.42	0.01	92.13
	2	5.23	9.01	3.00	4.93	0.02	77.82
	4	10.85	9.57	4.42	5.40	2.70	67.06
	8	11.20	9.62	5.27	5.08	4.57	64.26
	12	11.26	9.59	5.23	5.45	5.73	62.74
Correlation Matrix							
G. Cons	12	1.00					
PermY	12	-0.12	1.00				
TranY	12	-0.13	0.02	1.00			
PermR	12	0.18	-0.43	-0.42	1.00		
TranR	12	-0.35	-0.31	-0.30	0.05	1.00	
CA	12	0.21	-0.21	-0.35	-0.23	-0.30	1.00

Note: Variance decomposition is obtained using variable ordering 1.

Table 7.8. Variance Decomposition: Japan and B-N Decomposition

	<u>Forecast Horizon</u>		<u>Innovation in</u>				
	k	G. Cons	TranY	PermY	TranR	PermR	CA
G. Cons	1	100.00	0.00	0.00	0.00	0.00	0.00
TranY	1	0.92	99.08	0.00	0.00	0.00	0.00
	4	2.60	90.51	1.96	2.72	1.55	0.66
	8	11.69	76.39	2.18	7.43	1.46	0.85
	12	12.62	75.23	2.21	7.61	1.50	0.83
PermY	1	0.16	11.06	88.78	0.00	0.00	0.00
	4	7.06	33.25	51.18	1.50	3.13	3.88
	8	11.94	31.06	47.07	2.37	2.94	4.63
	12	12.58	30.72	46.25	2.94	2.92	4.59
TranR	1	0.48	2.82	4.97	91.73	0.00	0.00
	4	1.36	2.67	10.91	81.01	2.63	1.43
	8	3.64	12.63	8.41	70.22	3.68	1.41
	12	3.95	10.75	7.82	72.21	3.88	1.38
PermR	1	3.52	0.39	0.03	17.31	78.75	0.00
	4	10.86	2.07	0.81	21.06	59.78	5.42
	8	12.20	5.08	1.44	23.10	51.57	6.60
	12	12.82	5.12	1.61	25.76	48.35	6.34
CA	1	0.02	2.10	0.32	0.80	4.63	92.13
	2	5.23	1.88	10.13	0.70	4.25	77.82
	4	10.85	4.53	9.45	1.66	6.45	67.06
	8	11.20	5.92	8.97	3.40	6.24	64.26
	12	11.26	5.86	8.96	5.00	6.18	62.74
Correlation Matrix							
G. Cons	12	1.00					
TranY	12	-0.14	1.00				
PermY	12	-0.11	-0.05	1.00			
TranR	12	-0.35	-0.39	-0.27	1.00		
PermR	12	0.29	-0.38	-0.32	-0.03	1.00	
CA	12	0.21	-0.35	-0.17	-0.34	-0.15	1.00

Note: Variance decomposition is obtained using variable ordering 2.

Innovation Accounting: The Blanchard-Quah Decomposition

Results of variance decomposition using data from the Blanchard-Quah decomposition method are presented in Tables 7.9 and 7.10, under variable orderings 1 and 2, respectively. These two tables show that transitory income changes explain more of the forecast error variance than any other variable for all forecast horizons except the 2-period ahead forecast. At 4 period-ahead horizon, transitory income explains about 9 percent of the forecast error variance which is about the same as the contribution of the permanent component in real income. However, for 12-period ahead forecast the contribution of the former rises to about 12 percent of the forecast error variance, while that the latter remains stagnant at about 8 percent. This pattern is true irrespective of the variable ordering considered.

The contributions of transitory and permanent components of real exchange rate are dependent on the variable ordering used: the variable appearing first shows more explanatory power, although individually each explains less than 5 percent of the forecast error variance. It may be noted that under variable ordering 1 transitory real exchange rate explains less than 1 percent of the forecast error variance, while its contribution is about 2 percent under ordering 2. This shows that transitory real exchange rate is not vary important variable in explaining forecast error variance of Japanese current account balance. Government consumption explains between 8 to 9 percent of the 12-period ahead forecast horizon under either variable ordering.

Table 7.9. Variance Decomposition: Japan and B-Q Decomposition

	<u>Forecast Horizon</u>		<u>Innovation in</u>				
	k	G. Cons	PermY	TranY	PermR	TranR	CA
G. Cons	All	100.00	0.00	0.00	0.00	0.00	0.00
PermY	1	0.22	99.78	0.00	0.00	0.00	0.00
	4	15.25	69.53	1.54	6.31	0.85	6.52
	8	15.10	64.34	4.64	7.40	1.05	7.48
	12	14.93	63.33	4.87	7.23	1.31	8.32
TranY	1	2.03	6.03	91.94	0.00	0.00	0.00
	4	5.82	11.35	74.15	1.41	0.21	7.06
	8	10.76	12.22	60.23	1.17	0.59	15.04
	12	10.88	12.84	56.56	2.34	1.71	15.67
PermR	1	1.39	4.76	90.00	3.85	0.00	0.00
	4	12.70	7.42	67.40	5.13	0.60	6.75
	8	12.94	8.80	64.60	5.36	0.79	7.50
	12	12.96	8.98	63.56	5.76	0.96	7.77
TranR	1	2.17	46.49	45.14	3.76	2.44	0.00
	4	6.76	49.06	26.25	0.61	4.25	13.08
	8	14.21	38.07	19.89	2.68	3.95	21.21
	12	14.64	38.42	17.54	3.07	5.14	21.19
CA	1	0.01	0.86	2.41	0.32	0.50	95.90
	2	4.92	7.31	1.97	4.26	0.53	81.01
	4	7.92	8.04	8.96	3.86	0.65	70.56
	8	7.68	8.42	11.26	3.42	0.90	68.32
	12	8.69	8.68	11.85	3.94	0.93	65.90
Correlation Matrix							
G. Cons	12	1.00					
PermY	12	0.08	1.00				
TranY	12	-0.20	-0.56	1.00			
PermR	12	0.70	0.20	-0.31	1.00		
TranR	12	0.19	0.30	-0.25	-0.13	1.00	
CA	12	-0.09	-0.40	-0.52	-0.05	-0.08	1.00

Note: Variance decomposition is obtained using variable ordering 1.

Table 7.10. Variance Decomposition: Japan and B-Q Decomposition

	<u>Forecast Horizon</u>		<u>Innovation in</u>				
	k	G. Cons	TranY	PermY	TranR	PermR	CA
G.Cons	All	100.00	0.00	0.00	0.00	0.00	0.00
TranY	1	2.03	97.97	0.00	0.00	0.00	0.00
	4	5.82	84.04	1.46	0.73	0.89	7.06
	8	10.76	69.30	3.15	0.62	1.13	15.04
	12	10.88	65.47	3.94	2.73	1.32	15.67
PermY	1	0.22	6.14	93.64	0.00	0.00	0.00
	4	15.25	5.49	65.58	2.07	5.08	6.52
	8	15.10	7.83	61.15	3.42	5.03	7.48
	12	14.93	8.34	59.86	3.46	5.08	8.32
TranR	1	2.17	67.24	24.38	6.20	0.00	0.00
	4	6.76	44.42	30.89	3.04	1.81	13.08
	8	14.21	33.41	24.55	4.51	2.11	21.21
	12	14.64	29.80	26.16	5.56	2.65	21.19
PermR	1	1.39	94.70	0.06	2.34	1.52	0.00
	4	12.70	71.40	3.42	4.15	1.58	6.75
	8	12.94	68.78	4.63	4.25	1.90	7.50
	12	12.96	67.35	5.19	4.78	1.94	7.77
CA	1	0.01	3.00	0.26	0.00	0.82	95.90
	2	4.92	2.82	6.46	3.16	1.63	81.01
	4	7.92	8.48	8.52	2.76	1.75	70.56
	8	7.68	11.77	7.92	2.58	1.73	68.32
	12	8.69	12.39	8.14	3.00	1.87	65.90
Correlation Matrix							
G. Cons		1.00					
TranY		-0.24	1.00				
PermY		0.15	-0.58	1.00			
TranR		0.53	-0.07	-0.02	1.00		
PermR		0.66	-0.57	0.63	0.24	1.00	
CA		-0.09	-0.58	-0.30	-0.06	-0.08	1.00

Note: Variance decomposition is obtained using variable ordering 2.

Thus, the variance decomposition results for US data suggest that transitory and permanent components in income and real exchange rate have quite different explanatory power. At a very short forecast horizon, transitory income's contribution in explaining forecast error variance of current account far exceeds that of permanent income. This is true irrespective of the variable ordering used in obtaining the variance decomposition. At longer forecast horizons, the contribution of transitory income grows, while that of permanent income grows at a faster rate. Consequently, the gap between the relative contributions of transitory and permanent components of income narrows with the length of the forecast horizon. The dominance of transitory over permanent income in explaining forecast error variance of US current account balance is more pronounced if data from the Blanchard-Quah decomposition is used in the analysis.

Relative importance of transitory and permanent components of real exchange rate in explaining the forecast error variance of US current account balance is less conclusive. When data from the Beveridge-Nelson decomposition is used, no definitive conclusion can be reached on this issue. The variable appearing first in the Choleski factorization seems to have the higher explanatory power, leading to no precise conclusion. However, when data from the Blanchard-Quah decomposition are used, we find that the permanent component contributes about 3 to 4 times more than the transitory component in real exchange rate. The contribution of relative government consumption is very small at very short forecast horizons irrespective of the decomposition method and variable ordering used. However, for 12 period ahead forecast error variance, its contribution

remains stable around 10 percent. Therefore, we find that for near term forecast, transitory real income and permanent real exchange rate are the two most important variable in explaining the forecast error variance of US current account balance.

Variance decomposition results for Japan show that transitory real exchange rate has insignificant contribution towards explaining forecast error variance of Japanese current account balance. Permanent component in real exchange rate contributes about 5 percent of the forecast error variance of current account, which is rather low (although this is higher than the contribution of transitory component in real exchange rate). If data from the Beveridge-Nelson decomposition are used then permanent income seems to have significantly more explanatory power than transitory income. However, when data from Blanchard-Quah decomposition is used, it is the transitory component in income that has a relatively higher explanatory power. Thus, the results are dependent on the method used to decompose income. The contribution of permanent income is similar under either decomposition method while that of the transitory income is higher if the Blanchard-Quah is used in the decomposition stage. Contribution of relative government consumption is rather low at short forecast horizons, while it explains about 10 percent of the forecast error variance of current account for 12 period ahead forecasts.

CHAPTER 8. SUMMARY AND CONCLUSIONS

It is generally recognized that the current account balance of a country evolves according to the intertemporal saving-investment decisions by economic agents. Therefore, any theoretical analysis of the causes and consequences of movements in the current account balance should necessarily include national saving and investment behavior in the analysis. One purpose of theoretical model building is to increase our understanding of the real world economic issues. From this perspective, the usefulness of theoretical analysis will depend, at least partially, upon its ability to explain observed movements of important macroeconomic variables. This study is an effort to examine whether real world data support the predictions of the intertemporal models of current account determination. The motivation behind the study is to see if the intertemporal models can explain the observed movements in current account balance, or we need to explore alternative and new explanation.

In this study we illustrate how we can approach the issue of identifying transitory and permanent components in economic variables that are characterized by the presence of unit roots. Isolation of the stationary component from a nonstationary variable may be an important step in many econometric analysis.

The empirical results of our analysis is not very conclusive. Results of cointegration analysis with Japanese data are more supportive of the predictions of the intertemporal models of current account determination than those for US data. We find

that permanent changes in income have insignificant effects on Japanese current account balance. On the other hand, we find that similar income changes in US data have statistically significant effects on US current account balance. Real exchange rate movements have significant effects on the current account balance of both the US and Japan. For both countries, irrespective of the decomposition technique used, a permanent increase in the relative price in rest of the world improves the current account balance. These results directly contradict the results of two recent empirical studies, namely those by Yellen and Rose (1989), and Rose (1991). They find no significant relationship between trade balance and exchange rate. However, the results of Yellen and Rose (1989), and Rose (1991) may be, at least partially, due to the choice of Engle-Granger (1987) single equation procedure, and Stock-Watson (1988) principal component procedure. It has been shown by Gonzalo (1991) that these two tests of cointegration are among the least powerful tests to detect cointegration.

Results of error correction models suggest existence of long-run equilibrium relationship among current account balance, government consumption, and transitory and permanent components of income and real exchange rate. The estimated error correction models show that in most cases the current account balance adjusts in response to deviations from long-run equilibrium. Transitory and permanent components in real exchange rate show some sign of adjustment, specially in the case of Japanese data. Income seems to respond the least in response to deviations from long-run equilibrium.

Impulse response functions and variance decomposition show that in the short-run transitory and permanent changes in income have differential influence on current account balance. In the very short run, transitory income seems affect the US current account more than permanent changes in income. This is true for data obtained from both decomposition methods. For Japanese data, impulse response functions seem to suggest that transitory income changes have greater effect of current account balance. On the other hand, the results of variance decomposition are sensitive to the decomposition technique used. If data from the Beveridge-Nelson decomposition is used, permanent income seems to have more power in explaining near term forecast error variance. The opposite is true for data obtained from the Blanchard-Quah decomposition. The effects of relative influences of transitory and permanent changes in real exchange rate on current account balance more inconclusive. For US data the results are sensitive to decomposition technique and variable ordering used. For Japanese data, no systematic pattern can be identified from the impulse response and variance decomposition analysis.

Finally, it should be mentioned that the findings of this study are conditional on the decomposition method used. The methods used to identify the transitory and permanent components in a time series data are rather new. There are other methods that have been suggested by others, and there is no theoretical reason to discriminate among these techniques. It is only through additional works in this line that we shall be able to make more definitive judgment about the techniques of isolating the stationary component from a unit root process.

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APPENDIX A

The periodic utility function is:

$$u_t = u(c_t) = a_0 + a_1 c_t - \frac{a_2}{2} c_t^2 \quad [a.1]$$

and the periodic budget constraint is $A_{t+1} = R[A_t + y_t - c_t]$. Given the restriction on borrowing, the difference equation describing the budget equation may be solved, via successive recursion, to obtain:

$$\sum_{j=0}^{\infty} \left(\frac{1}{R}\right)^j E_t c_{t+j} = A_t + \sum_{j=0}^{\infty} \left(\frac{1}{R}\right)^j E_t y_{t+j} \quad [a.2]$$

which states that the expected present value of consumption at time t equals the expected present value of income plus current asset holding. The problem is to choose the sequence $\{c_t\}_{t=0}^{\infty}$, subject to the budget constraint, to maximize the objective function:

$$U = E_0 \sum_{t=0}^{\infty} \beta^t u(c_t) \quad [a.3]$$

where $u = u(c_t)$ is given by [a.1]. The Euler equation for the optimization problem is:

$$E_t R \beta u'(c_{t+1}) = (\beta R)^{-1} u'(c_t) \quad [a.4]$$

where $c_t = y_t + A_t - R^{-1} A_{t+1}$. Given the specific utility function, the Euler equation reduces to:

$$E_t c_{t+1} = \alpha + (\beta R)^{-1} c_t \quad \text{where } \alpha = \frac{a_1 [1 - (\beta R)^{-1}]}{a_2}. \quad [a.5]$$

Successive recursion on the above equation yields:

$$E_t c_{t+j} = \alpha \left[\frac{1 - \gamma^j}{1 - \gamma} \right] + \gamma^j c_t \quad \text{where } \gamma = (\beta R)^{-1}. \quad [a.6]$$

Substituting [a.6] in [a.5], we get:

$$\sum_{j=0}^{\infty} \left(\frac{1}{R}\right)^j \left[\alpha \left(\frac{1 - \gamma^j}{1 - \gamma} \right) + \gamma^j c_t \right] = A_t + \sum_{j=0}^{\infty} \left(\frac{1}{R}\right)^j E_t y_{t+j}$$

$$\text{Let } \Delta = \sum_{j=0}^{\infty} \left(\frac{1}{R}\right)^j E_t y_{t+j}.$$

Therefore, we have:

$$A_t + \Delta = \frac{\alpha}{1-\gamma} \sum_{j=0}^{\infty} \left(\frac{1}{R}\right)^j (1-\gamma^j) + \sum_{j=0}^{\infty} \left(\frac{1}{R}\right)^j \gamma^j c_t$$

Rearranging, we get:

$$\sum_{j=0}^{\infty} \left(\frac{\gamma}{R}\right)^j c_t = A_t + \Delta - \frac{\alpha}{1-\gamma} \sum_{j=0}^{\infty} \left(\frac{1}{R}\right)^j (1-\gamma^j)$$

$$\text{Now, } \sum_{j=0}^{\infty} \left(\frac{\gamma}{R}\right)^j = \frac{1}{1-\frac{\gamma}{R}} = \frac{\beta R^2}{\beta R^2 - 1}.$$

Therefore, we can write:

$$c_t = \left(1 - \frac{1}{\beta R^2}\right) \left[A_t + \Delta - \frac{\alpha}{1-\gamma} \sum_{j=0}^{\infty} \left(\frac{1}{R}\right)^j (1-\gamma^j) \right] \quad [\text{a.7}]$$

$$\begin{aligned} \text{Now } \frac{\alpha}{1-\gamma} \sum_{j=0}^{\infty} \left(\frac{1}{R}\right)^j (1-\gamma^j) &= \frac{\alpha}{1-\gamma} \left\{ \sum_{j=0}^{\infty} \left[\frac{1}{R^j} - \left(\frac{\gamma}{R}\right)^j \right] \right\} \\ &= \frac{\alpha}{1-\gamma} \left[\frac{1}{1-\frac{1}{R}} - \frac{\beta R^2}{\beta R^2 - 1} \right] \\ &= \frac{\alpha}{1-\gamma} \left[\frac{R}{R-1} - \frac{\beta R^2}{\beta R^2 - 1} \right] \end{aligned}$$

Substituting the above expression in [a.7], we get:

$$\begin{aligned} c_t &= \left(1 - \frac{1}{\beta R^2}\right) [A_t + \Delta] - \left(\frac{\beta R^2 - 1}{\beta R^2}\right) \frac{\alpha}{1-\gamma} \left[\frac{R}{R-1} - \frac{\beta R^2 - 1}{\beta R^2} \right] \\ &= \left(1 - \frac{1}{\beta R^2}\right) [A_t + \Delta] - \frac{\alpha}{1-\gamma} \left\{ \frac{\beta R^2 - 1 - \beta R^2 + \beta R}{\beta R(R-1)} \right\} \\ &= \left(1 - \frac{1}{\beta R^2}\right) [A_t + \Delta] - \alpha \left(\frac{\beta R}{\beta R - 1} \right) \left\{ \frac{\beta R - 1}{\beta R(R-1)} \right\} \end{aligned}$$

Substituting back for Δ , we have:

$$c_t = -\frac{\alpha}{R-1} + \left(1 - \frac{1}{\beta R^2}\right) \left(A_t + \sum_{j=0}^{\infty} \left(\frac{1}{R}\right)^j E_t y_{t+j}\right)$$

Assuming zero initial asset holding (implying $A_t = 0$), $\beta R = 1$, and denoting $\delta = (1/R)$, the consumption function may be written as:

$$c_t = (1 - \delta) \sum_{j=0}^{\infty} \delta^j E_t y_{t+j}$$

This is the consumption function equation used to describe the current account balance equation in Chapter 3.

APPENDIX B

Starting from no net indebtedness at the beginning of period t , the current account balance of a country in period t can be modeled with the help of the following four equations:

$$C_t = (1 - \delta) \sum_{j=0}^{\infty} \delta^j E_t Y_{t+j}, \quad [\text{b.1}]$$

$$Y_t = \tau_t + \varepsilon_{T,t}, \quad [\text{b.2}]$$

$$\tau_t = \tau_{t-1} + \varepsilon_{p,t}, \quad [\text{b.3}]$$

$$b_t = Y_t - C_t, \quad [\text{b.4}]$$

where

C_t = consumption in period t ,

Y_t = real income in period t ,

τ_t = permanent component in Y_t ,

$\varepsilon_{p,t}$ = shocks to the permanent component in Y_t ,

$\varepsilon_{T,t}$ = transitory shocks to Y_t ,

b_t = current account balance in period t .

Now the permanent component in Y_t can be written as:

$$\tau_t = \tau_{t-1} + \varepsilon_{p,t} \quad [\text{b.5}]$$

Equivalently, we may write:

$$(1 - L)\tau_t = \varepsilon_{p,t}$$

where L is the lag operator. Solving the above difference equation, we get:

$$\tau_t = \sum_{s=0}^{\infty} \varepsilon_{p,t-s} \quad [\text{b.6}]$$

Combining [b.5] and [b.6], we can express Y_t in terms of $\varepsilon_{p,t}$ and $\varepsilon_{T,t}$ as follows:

$$Y_t = \sum_{s=0}^{\infty} \varepsilon_{p,t-s} + \varepsilon_{T,t}.$$

Current account balance in period can be written as:

$$\begin{aligned}
b_t &= Y_t - (1-\delta) \sum_{j=0}^{\infty} \delta^j E_t Y_{t+j} \\
&= Y_t - (1-\delta)Y_t - \delta(1-\delta)E_t Y_{t+1} - \delta^2(1-\delta)E_t Y_{t+2} - \delta^3(1-\delta)E_t Y_{t+3} - \dots
\end{aligned}$$

So, current account balance in period t is given by:

$$b_t = \delta Y_t - \delta(1-\delta)E_t Y_{t+1} - \delta^2(1-\delta)E_t Y_{t+2} - \delta^3(1-\delta)E_t Y_{t+3} - \dots$$

Substituting $Y_t = \sum_{s=0}^{\infty} \varepsilon_{p,t-s} + \varepsilon_{T,t}$, we can write:

$$b_t = \delta \left[\sum_{s=0}^{\infty} \varepsilon_{p,t-s} + \varepsilon_{T,t} \right] - \delta(1-\delta)E_t \left[\sum_{s=0}^{\infty} \varepsilon_{p,t+1-s} + \varepsilon_{T,t+1} \right] - \delta^2(1-\delta)E_t \left[\sum_{s=0}^{\infty} \varepsilon_{p,t+2-s} + \varepsilon_{T,t+2} \right] - \delta^3(1-\delta)E_t \left[\sum_{s=0}^{\infty} \varepsilon_{p,t+3-s} + \varepsilon_{T,t+3} \right] - \dots$$

Expanding the terms under summation sign, and rearranging, we can write:

$$\begin{aligned}
b_t &= \delta(\varepsilon_{p,t} + \varepsilon_{p,t-1} + \varepsilon_{p,t-2} + \varepsilon_{p,t-3} + \dots + \varepsilon_{T,t}) \\
&\quad - \delta E_t(\varepsilon_{p,t+1} + \varepsilon_{p,t} + \varepsilon_{p,t-1} + \varepsilon_{p,t-2} + \varepsilon_{p,t-3} + \dots + \varepsilon_{T,t+1}) \\
&\quad + \delta^2 E_t(\varepsilon_{p,t+1} + \varepsilon_{p,t} + \varepsilon_{p,t-1} + \varepsilon_{p,t-2} + \varepsilon_{p,t-3} + \dots + \varepsilon_{T,t+1}) \\
&\quad - \delta^2 E_t(\varepsilon_{p,t+2} + \varepsilon_{p,t+1} + \varepsilon_{p,t} + \varepsilon_{p,t-1} + \varepsilon_{p,t-2} + \varepsilon_{p,t-3} + \dots + \varepsilon_{T,t+2}) \\
&\quad + \delta^3 E_t(\varepsilon_{p,t+2} + \varepsilon_{p,t+1} + \varepsilon_{p,t} + \varepsilon_{p,t-1} + \varepsilon_{p,t-2} + \varepsilon_{p,t-3} + \dots + \varepsilon_{T,t+2}) \\
&\quad - \delta^3 E_t(\varepsilon_{p,t+3} + \varepsilon_{p,t+2} + \varepsilon_{p,t+1} + \varepsilon_{p,t} + \varepsilon_{p,t-1} + \varepsilon_{p,t-2} + \varepsilon_{p,t-3} + \dots + \varepsilon_{T,t+3}) + \dots
\end{aligned}$$

Now $E_t \varepsilon_{p,t+i} = 0$ ($\forall i > 0$), and $E_t \varepsilon_{T,t+k} > 0$ ($\forall k > 0$). Hence we can write:

$$\begin{aligned}
b_t &= \delta(\varepsilon_{p,t} + \varepsilon_{p,t-1} + \varepsilon_{p,t-2} + \varepsilon_{p,t-3} + \dots + \varepsilon_{T,t}) - \delta(\varepsilon_{p,t} + \varepsilon_{p,t-1} + \varepsilon_{p,t-2} + \varepsilon_{p,t-3} + \dots) \\
&\quad + \delta^2(\varepsilon_{p,t} + \varepsilon_{p,t-1} + \varepsilon_{p,t-2} + \varepsilon_{p,t-3} + \dots) - \delta^2(\varepsilon_{p,t} + \varepsilon_{p,t-1} + \varepsilon_{p,t-2} + \varepsilon_{p,t-3} + \dots) \\
&\quad + \delta^3(\varepsilon_{p,t} + \varepsilon_{p,t-1} + \varepsilon_{p,t-2} + \varepsilon_{p,t-3} + \dots) - \delta^3(\varepsilon_{p,t} + \varepsilon_{p,t-1} + \varepsilon_{p,t-2} + \varepsilon_{p,t-3} + \dots) + \dots
\end{aligned}$$

It is obviously from the above equation that when the current account balance, b_t , is expressed in terms of transitory shocks ($\varepsilon_{T,t}$) to real income, and shocks to the permanent component in real income ($\varepsilon_{p,t}$), all terms involving $\varepsilon_{p,t}$ cancel each other out, and we get:

$$b_t = \delta \varepsilon_{T,t}.$$

Hence we see that starting from a position of no net indebtedness in period t, permanent shocks to income leave the current account unaffected. On the other hand, a transitory shock in period t leads to a surplus in the current account balance.

APPENDIX C

From standard theory of consumer behavior, given our assumptions about preferences and prices, the following relations may be derived:

$$\eta_{c_0, \alpha} = \gamma(\varpi - 1), \quad [\text{c.1}]$$

$$\eta_{c_1, \alpha} = -[(1 - \gamma)\sigma + \gamma], \quad [\text{c.2}]$$

$$\eta_{n_t, p_{nt}} = -\beta_{mt}\sigma_{nm} - \beta_{xt}\sigma_{nx} - \beta_{nt}, \quad [\text{c.3}]$$

$$\eta_{n_t, p_{mt}} = \beta_{mt}(\sigma_{nm} - 1), \quad [\text{c.4}]$$

where

$\eta_{c_0, \alpha}$ = the elasticity of C_0 with respect to α_{c1} ,

γ = is the average propensity to save, defined as the ratio of future consumption to lifetime wealth (in present value),

$\eta_{c_1, \alpha}$ = the elasticity of C_1 with respect to α_{c1} ,

$\eta_{n_t, p_{nt}}$ = elasticity of demand for domestic nontradable good in period t ($t = 0, 1$) with respect to p_{nt} ($t = 0, 1$),

$\eta_{n_t, p_{mt}}$ = elasticity of demand for domestic nontradable good in period t ($t = 0, 1$) with respect to p_{mt} ($t = 0, 1$),

σ = intertemporal elasticity of substitution, defined as

$$\sigma = \frac{\partial \log(C_1 / C_0)}{\partial \log[(\partial U / \partial C_0) / (\partial U / \partial C_1)]},$$

σ_{ij} = Allen elasticity of substitution between goods i and j .

Note that $\sigma > 0$, $0 < \gamma < 1$, $\eta_{c_1, \alpha}$ is negative. Further, since $-\beta_{mt}\sigma_{nm} - \beta_{xt}\sigma_{nx}$ denotes the compensated effect, it is non-positive because of the negative semidefiniteness of the Slutsky substitution matrix.

In equilibrium the markets for nontradables must clear in each period and in each country. That implies the equilibrium requirements:

$$C_{n0}[p_{n0}, p_{m0}, P_0 C_0(\alpha_{c1}, W_{c0})] = Q_{n0}, \quad [c.5]$$

$$C_{n1}[p_{n1}, p_{m1}, P_1 C_1(\alpha_{c1}, W_{c0})] = Q_{n1}. \quad [c.6]$$

Given initial output, total differentiation of [c.5] and [c.6] gives:

$$(\eta_{n_0 p_{n0}} + \beta_{n0}) \hat{p}_{n0} + (\eta_{n_0 p_{m0}} + \beta_{m0}) \hat{p}_{m0} + \eta_{c_0 \alpha} \hat{\alpha}_{c1} + \hat{W}_{c0} = 0, \quad [c.7]$$

$$(\eta_{n_1 p_{n1}} + \beta_{n1}) \hat{p}_{n1} + (\eta_{n_1 p_{m1}} + \beta_{m1}) \hat{p}_{m1} + \eta_{c_1 \alpha} \hat{\alpha}_{c1} + \hat{W}_{c0} = 0, \quad [c.8]$$

where a circumflex above a variable denotes a proportional change, and we have used the property that the elasticity of the price index P_i with respect to a change in one of the temporal relative prices (p_{m1} , or p_{n1}) is simply equal to the corresponding expenditure share. Under our assumptions, the discount factor relevant for domestic (real) consumption, α_{c1} , evolves according to:

$$\hat{\alpha}_{c1} = \beta_{m1} \hat{p}_{m1} + \beta_{n1} \hat{p}_{n1} - \beta_{m0} \hat{p}_{m0} - \beta_{n0} \hat{p}_{n0}. \quad [c.9]$$

We may recall that (assuming no initial net debt at the beginning of period 0) real wealth is given by:

$$W_{c0} = [\{Q_{x0} + p_{n0} Q_{n0}\} + \alpha_{x1} \{Q_{x1} + p_{n1} Q_{n1}\}] / P_0.$$

Totally differentiating the above equation, we may obtain:

$$\hat{W}_{c0} = -\beta_{m0} \hat{p}_{m0} - \gamma \beta_{n0} \hat{p}_{n0} + \gamma \beta_{n1} \hat{p}_{n1}. \quad [c.10]$$

Substituting [c.9] and [c.10] in [c.7] and [c.8] and rearranging, we can obtain the following system:

$$\begin{bmatrix} -\beta_{m0} \sigma_{nm} - \beta_{x0} \sigma_{nx} - \beta_{n0} \gamma \sigma & \gamma \beta_{n1} \sigma \\ \beta_{n0} (1 - \gamma) \sigma & -\beta_{m1} \sigma_{nm} - \beta_{x1} \sigma_{nx} - \beta_{n1} (1 - \gamma) \sigma \end{bmatrix} \begin{bmatrix} \hat{p}_{n0} \\ \hat{p}_{n1} \end{bmatrix} \\ = \begin{bmatrix} -\beta_{m0} \{ \sigma_{nm} - (1 - \gamma) - \gamma \sigma \} & -\beta_{m1} \gamma (\sigma - 1) \\ -\beta_{m0} (1 - \gamma) (\sigma - 1) & -\beta_{m1} \{ \sigma_{nm} - \gamma - (1 - \gamma) \sigma \} \end{bmatrix} \begin{bmatrix} \hat{p}_{m0} \\ \hat{p}_{m1} \end{bmatrix}$$

From the above equation we can solve for \hat{p}_{n0} and \hat{p}_{n1} in terms of \hat{p}_{m0} and \hat{p}_{m1} . Specifically, we can obtain:

$$\begin{aligned}
\hat{p}_{n0} &= \Delta^{-1} [\beta_{m0} \{ \sigma_{nm} - (1-\gamma) - \gamma\sigma \} \{ \beta_{m1} \sigma_{nm} + \beta_{x1} \sigma_{nx} + \beta_{n1} (1-\gamma)\sigma \} \\
&\quad \{ \beta_{m0} (1-\gamma)(\sigma-1)\gamma\beta_{n1}\sigma \}] \hat{p}_{m0} \\
&\quad + \Delta^{-1} [\gamma\beta_{m1} (\sigma-1) \{ \beta_{m1} \sigma_{nm} + \beta_{x1} \sigma_{nx} + \beta_{n1} (1-\gamma)\sigma \} \\
&\quad \beta_{m1} \{ (\sigma_{nm} - \gamma) + (1-\gamma)\sigma \} \gamma\beta_{n1}\sigma] \hat{p}_{m1} \\
\hat{p}_{n1} &= \Delta^{-1} [\beta_{m0} (1-\gamma)(\sigma-1) \{ \beta_{m0} \sigma_{nm} + \beta_{x0} \sigma_{nx} + \beta_{n0} \gamma\sigma \} \\
&\quad + \beta_{m0} (1-\gamma)\beta_{n0}\sigma \{ \sigma_{nm} - (1-\gamma) - \gamma\sigma \}] \hat{p}_{m0} \\
&\quad + \Delta^{-1} [\beta_{m1} \{ \sigma_{nm} - \gamma - (1-\gamma)\sigma \} \{ \beta_{m0} \sigma_{nm} + \beta_{x0} \sigma_{nx} + \beta_{n0} \gamma\sigma \} \\
&\quad + \beta_{m1} (1-\gamma)\beta_{n0}\gamma\sigma(\sigma-1)] \hat{p}_{m1}
\end{aligned}$$

where ,

$$\Delta = [\beta_{m0}\sigma_{nm} + \beta_{x0}\sigma_{nx}] [\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx} + \beta_{n1}(1-\gamma)\sigma] + \beta_{n0}\gamma\sigma[\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx}] > 0.$$

Now we can obtain the following results:

$$\frac{d \log p_{n0}}{d \log p_{m0}} = [\sigma_{nm} - (1-\gamma) + \gamma\sigma] \Phi_1 + (\sigma_{nm} - 1) \Phi_2, \quad [c.11]$$

where

$$\begin{aligned}
\Phi_1 &= \frac{\beta_{m0}(\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx})}{(\beta_{m0}\sigma_{nm} + \beta_{x0}\sigma_{nx})[\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx} + \beta_{n1}(1-\gamma)\sigma] + \beta_{n0}\gamma\sigma(\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx})}, \\
\Phi_2 &= \frac{\beta_{m0}\beta_{n1}(1-\gamma)\sigma}{(\beta_{m0}\sigma_{nm} + \beta_{x0}\sigma_{nx})[\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx} + \beta_{n1}(1-\gamma)\sigma] + \beta_{n0}\gamma\sigma(\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx})}.
\end{aligned}$$

It may be noted that both $\Phi_1 > 0$, and $\Phi_2 > 0$.

$$\frac{d \log p_{n1}}{d \log p_{m0}} = [\sigma_{nm} - 1] \Phi_3 + (1-\gamma)(\sigma-1) \Phi_4, \quad [c.12]$$

where

$$\begin{aligned}
\Phi_3 &= \frac{\beta_{m0}\beta_{n0}(1-\gamma)\sigma}{(\beta_{m0}\sigma_{nm} + \beta_{x0}\sigma_{nx})[\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx} + \beta_{n1}(1-\gamma)\sigma] + \beta_{n0}\gamma\sigma(\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx})}, \\
\Phi_4 &= \frac{\beta_{m0}(\beta_{m0}\sigma_{nm} + \beta_{x0}\sigma_{nx})}{(\beta_{m0}\sigma_{nm} + \beta_{x0}\sigma_{nx})[\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx} + \beta_{n1}(1-\gamma)\sigma] + \beta_{n0}\gamma\sigma(\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx})}. \\
\frac{d \log p_{n0}}{d \log p_{m1}} &= (\sigma-1)\Omega_1 + (\sigma_{nm} - 1)\Omega_2, \quad [c.13]
\end{aligned}$$

where

$$\Omega_1 = \frac{\gamma\beta_{m1}(\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx})}{(\beta_{m0}\sigma_{nm} + \beta_{x0}\sigma_{nx})[\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx} + \beta_{n1}(1-\gamma)\sigma] + \beta_{n0}\gamma\sigma(\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx})},$$

$$\Omega_2 = \frac{\gamma\beta_{m1}\beta_{n1}\sigma}{(\beta_{m0}\sigma_{nm} + \beta_{x0}\sigma_{nx})[\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx} + \beta_{n1}(1-\gamma)\sigma] + \beta_{n0}\gamma\sigma(\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx})}.$$

$$\frac{d \log p_{n1}}{d \log p_{m1}} = [\sigma_{nm} - \gamma - (1-\gamma)\sigma]\Omega_3 + (\sigma_{nm} - 1)\Omega_4, \quad [c.14]$$

where

$$\Omega_3 = \frac{\beta_{m1}(\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx})}{(\beta_{m0}\sigma_{nm} + \beta_{x0}\sigma_{nx})[\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx} + \beta_{n1}(1-\gamma)\sigma] + \beta_{n0}\gamma\sigma(\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx})},$$

$$\Omega_4 = \frac{\gamma\beta_{m1}\beta_{n0}\sigma}{(\beta_{m0}\sigma_{nm} + \beta_{x0}\sigma_{nx})[\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx} + \beta_{n1}(1-\gamma)\sigma] + \beta_{n0}\gamma\sigma(\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx})}.$$

The effect of permanent changes in terms of trade on the prices of domestic nontradable good in periods 0 and 1 are obtained under the condition that $dp_{m0} = dp_{m1} = dp_m$. Under this condition, we get:

$$\frac{d \log p_{n0}}{d \log p_m} = [\sigma_{nm} - 1]\Delta_1, \quad [c.15]$$

where

$$\Delta_1 = \frac{\beta_m(\beta_m\sigma_{nm} + \beta_x\sigma_{nx} + \beta_n\sigma)}{(\beta_{m0}\sigma_{nm} + \beta_{x0}\sigma_{nx})[\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx} + \beta_{n1}(1-\gamma)\sigma] + \beta_{n0}\gamma\sigma(\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx})}.$$

Similarly, we have

$$\frac{d \log p_{n1}}{d \log p_m} = [\sigma_{nm} - 1]\Delta_2, \quad [c.16]$$

where

$$\Delta_2 = \frac{\beta_m(\beta_m\sigma_{nm} + \beta_x\sigma_{nx} + \beta_n\sigma)}{(\beta_{m0}\sigma_{nm} + \beta_{x0}\sigma_{nx})[\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx} + \beta_{n1}(1-\gamma)\sigma] + \beta_{n0}\gamma\sigma(\beta_{m1}\sigma_{nm} + \beta_{x1}\sigma_{nx})}.$$

It may be noted that the changes in the price of the domestic nontradable good in the two periods are same.

APPENDIX D

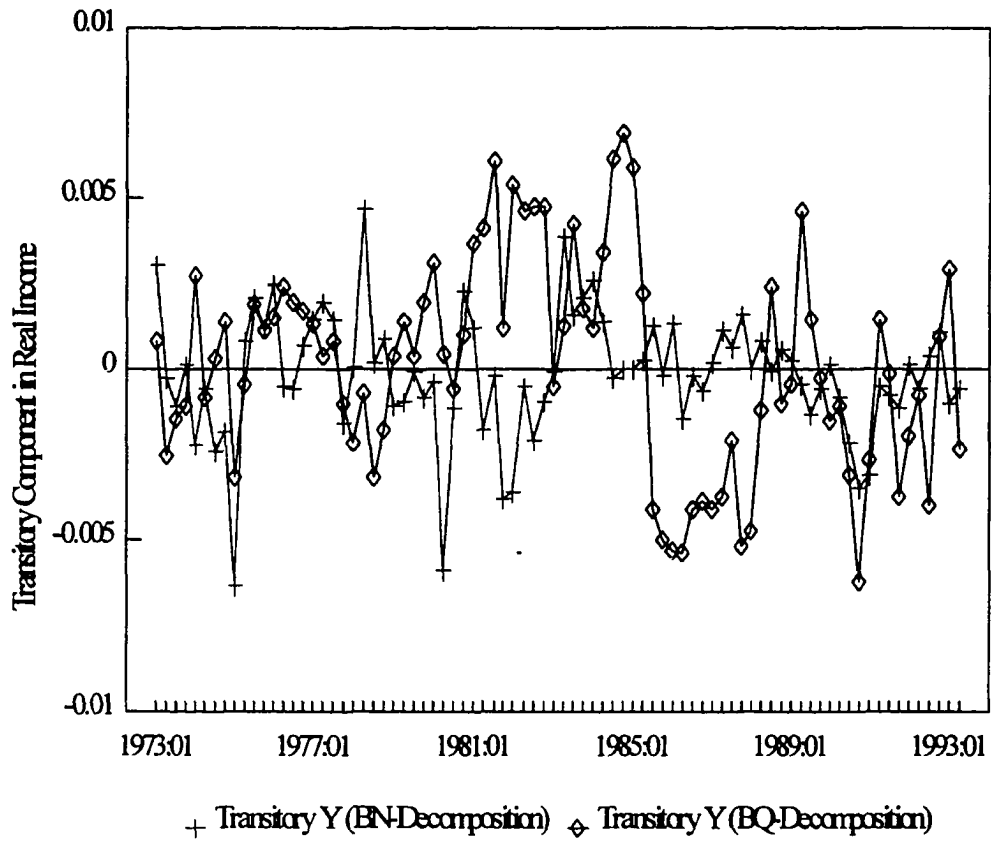


Figure D.1 Transitory Components in Real GNP: The United States

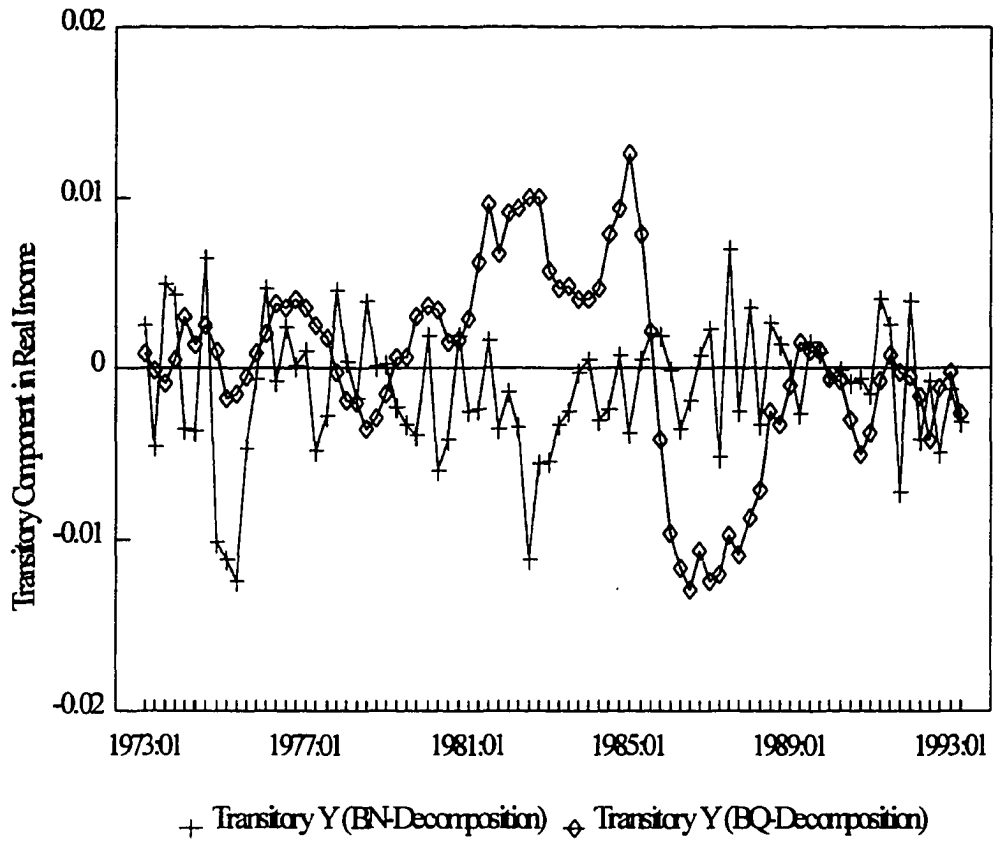


Figure D.2 Transitory Components in Real GNP: Rest of the World vis-à-vis the US

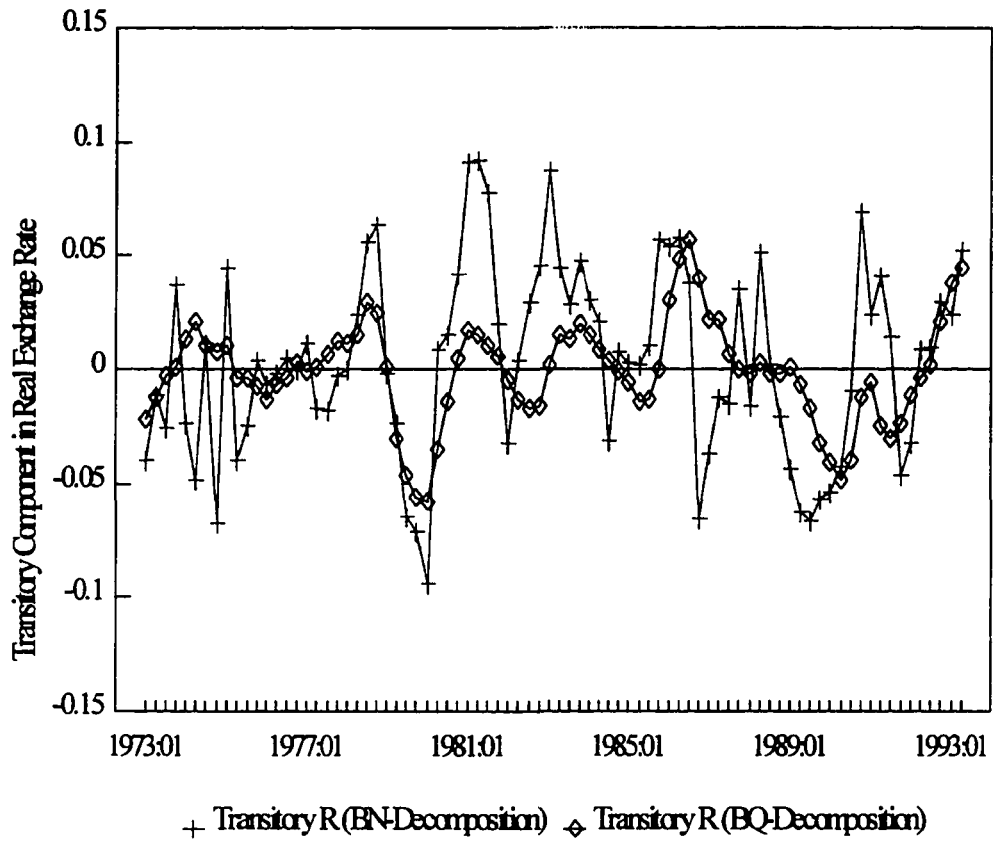


Figure D.3 Transitory Component in Real Exchange Rate: The United States

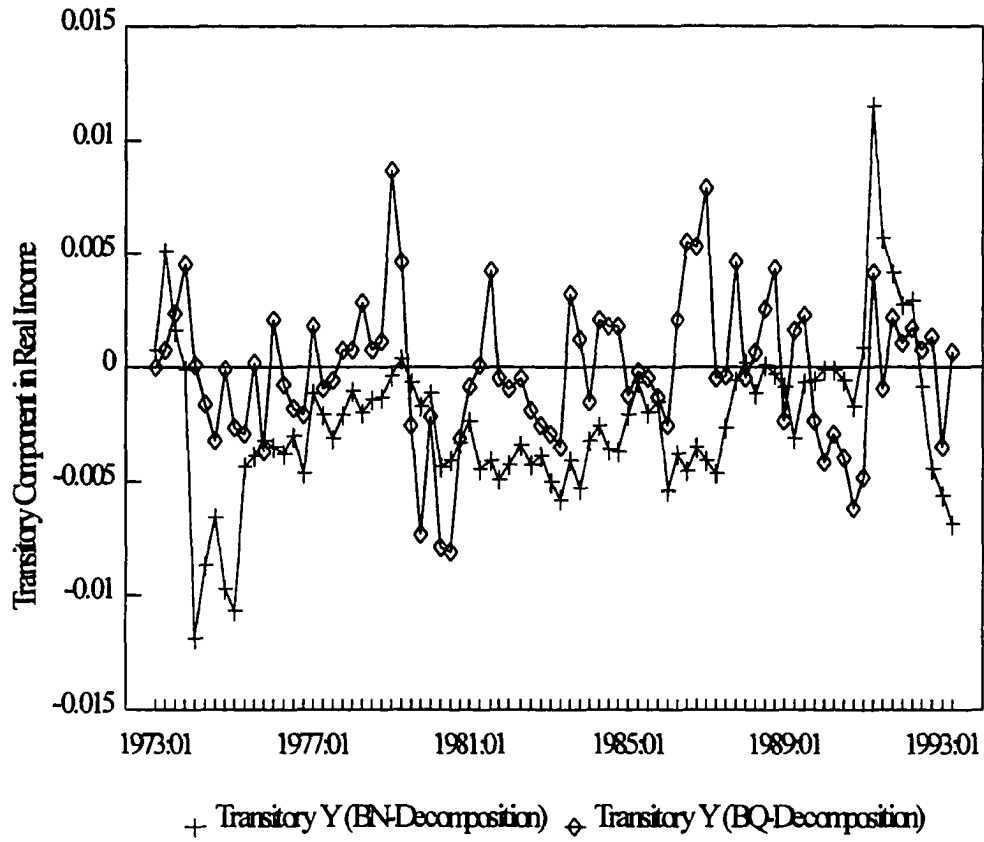


Figure D.4 Transitory Component in Real Income: Japan

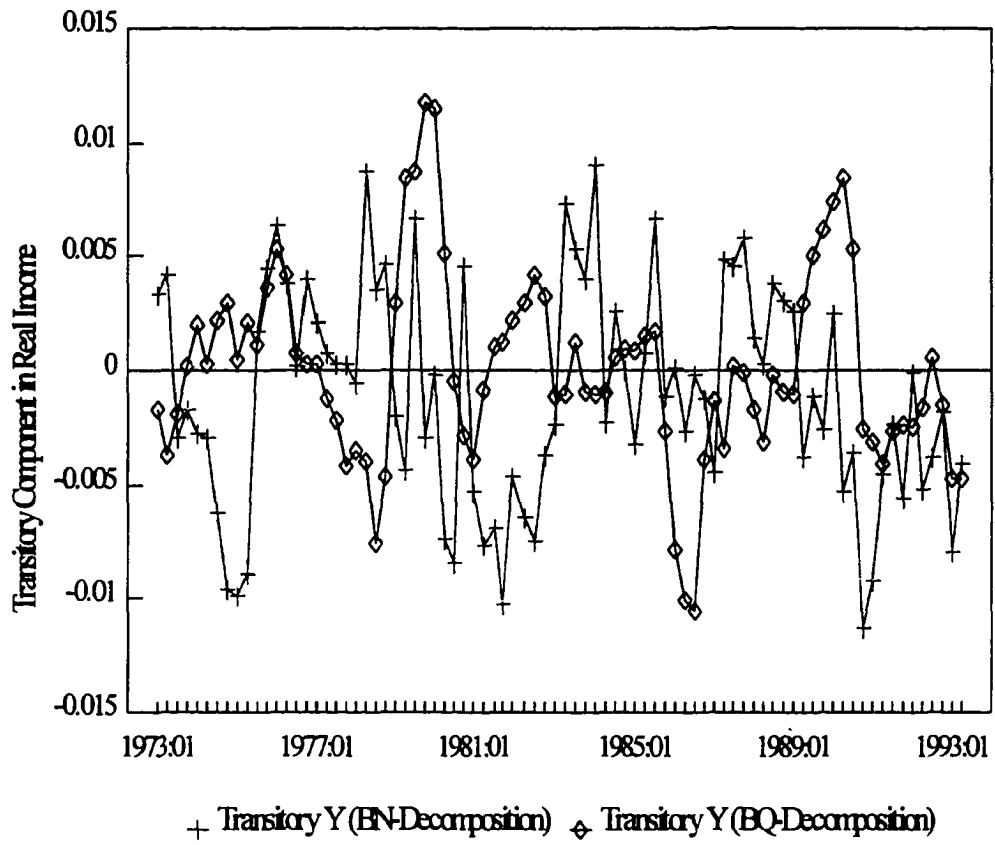


Figure D.5 Transitory Component in Real Income: Rest of the World vis-à-vis Japan

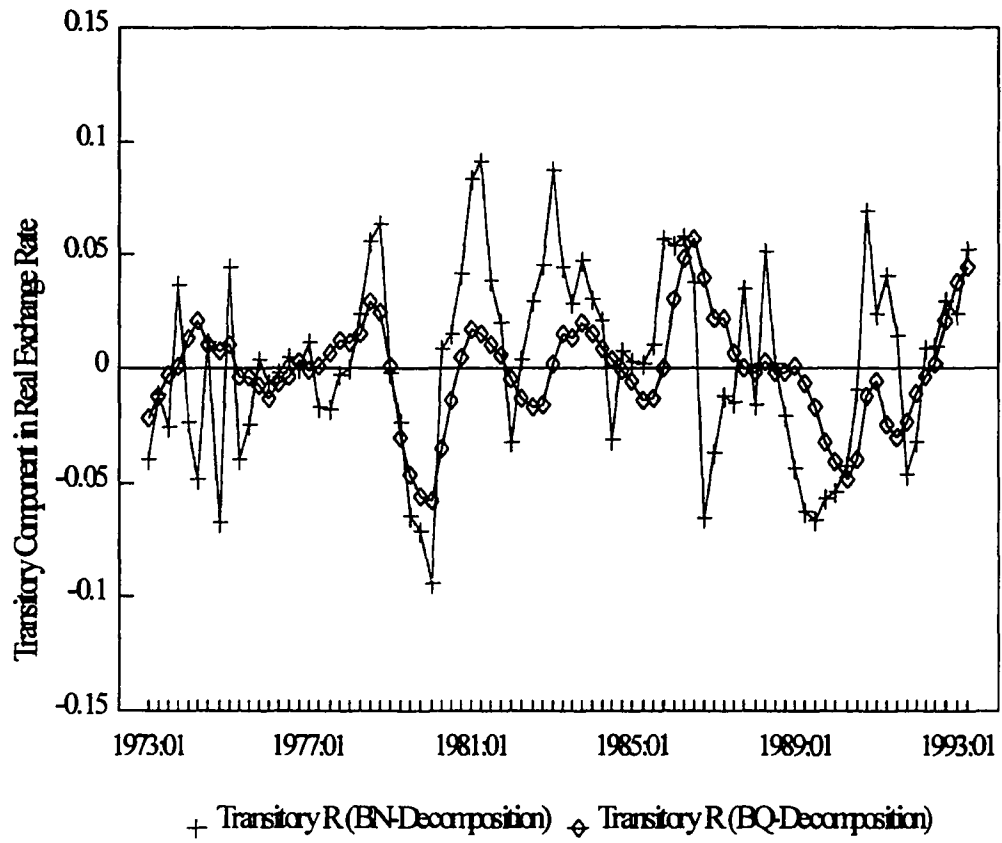


Figure D.6 Transitory Component in Real Exchange Rate: Japan

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