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Terms of Trade Shocks and the Current Account

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Abstract

This paper examines the relationship between terms of trade shocks, private saving, and the current account position. The relationship between these variables is theoretically ambiguous: an adverse transitory terms of trade shock can either induce a deterioration or an improvement in the current account, depending on whether the resulting income effects are greater or less than the resulting substitution effects. The substitution effects involve both intertemporally substituting consumption and intratemporally substituting consumption between importables and nontradables. The relative strength of these substitution effects is estimated using data for five OECD countries during 1970/95; both are found to exert large and significant effects on the current account balance.

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Summary

What happens to an economy's current account position if there is an adverse shock to its terms of trade? According to the Harberger-Laursen-Metzler (HLM) effect it deteriorates, because a deterioration in the terms of trade will decrease real income and savings. The HLM (income-based) effect assumes that the home country and the rest of the world produce the same good, and that this good is tradable across countries.

In a three-good (importable, exportable, and nontradable) model of intertemporal consumption, countries will consume both importables and nontradables, and substitution effects will also influence saving decisions. These substitution effects concern the ease with which, in response to a terms of trade shock, countries can switch between importables and nontradables (intratemporal substitution), and between current and future consumption in response to a shift in the relative price of current consumption (intertemporal substitution). Using data from five OECD countries between 1970/95, the paper undertakes generalized method of moments estimation of Euler equations derived from an intertemporal optimizing model of consumption. The results indicate that terms of trade shocks induce large and significant intratemporal and intertemporal substitution effects, which operate to offset any associated income effects on private saving decisions and the current account position.

I. INTRODUCTION

What happens to an economy's current account position if there is an adverse shock to its terms of trade? The answer to this question has been a matter of some debate since the early 1950s, when Harberger (1950) and Laursen and Metzler (1950) conjectured that a deterioration in the terms of trade lowered real income and savings, and consequently worsened the current account position. Moreover, this issue is important to policy makers, given that real non-oil commodity prices have steadily fallen since the mid-1970s, while the volatility of real commodity prices has risen (Reinhart and Wickham (1994)). While this commodity-based channel for the effects of changes in the terms of trade has been an important one, large disturbances in many countries' terms of trade have also occurred over the past two decades through changes in the relative price of intermediate inputs, and in the relative price of manufactured goods. These developments have had important effects on national consumption patterns, and accordingly on the pattern of flows of international capital.

The dominant explanation in the international finance literature of the main channel by which shocks to the terms of trade affect private saving decisions (and current account imbalances) is the Harberger-Laursen-Metzler effect. According to this income-based effect, an adverse transitory movement in the terms of trade results in a decrease in a country's current level of income which is larger than the decrease in its permanent income, causing a fall in aggregate saving and a deterioration of the current account position. This consumption-smoothing response of aggregate saving to transitory movements in the terms of trade assumes that the home country and the rest of the world produce the same good, and that this good is tradable across countries.

However, in a three-good (importable, exportable and nontradable) model of intertemporal consumption in which countries specialize in the production of different goods and transportation costs are non-zero, countries will consume both importables and nontradables, and relative price (substitution effects) will also influence saving decisions. Using empirical estimation of the Euler equations derived from an intertemporal optimizing model of consumption where agents seek to maximize utility through the consumption of importables and nontradables, this paper estimates the key parameters which influence the strength of the substitution effects on private saving decisions arising from movement in the terms of trade. These parameters include the ease with which countries can switch between importables and nontradables in response to a terms of trade shock (intratemporal substitution), and the ease with which countries can switch between current and future consumption in response to a shift in the relative price of current consumption induced by a terms of trade shock (intertemporal substitution). Large values of these parameters imply that in response to an adverse transitory terms of trade shock which raises the intertemporal relative price of consumption and induces a temporary real appreciation, aggregate private saving should rise and the current account position improve. This rise in saving will offset the decline in private saving (and consequent deterioration of the current account position) induced by the traditional consumption-smoothing effect.

A fall in the terms of trade lowers real national income, as a country must then either increase its net external liabilities to finance any consequent adverse movement in the current account, or allocate more of its production to exportables (to finance an unchanged level of importables). While terms of trade shocks are typically a key driver of fluctuations in real incomes in developing countries dependent on trade in primary commodities (Khan and Knight (1983)), such shocks are also important in several developed countries which are important participants in world commodity markets. This paper will examine the effect of terms of trade shocks on private consumption and the current account, using data for the period 1970-95 from three commodity-exporting developed countries (Australia, Canada and New Zealand), as well as two major industrial countries that are also important commodity exporters (the United Kingdom and the United States).

The contributions of this paper are threefold. First, we estimate the persistence of terms of trade shocks using a technique which enables us to determine quantitatively how long shocks to a particular series will last. Second, and in contrast to the nonoptimizing, static Harberger-Laursen-Metzler (HLM) approach to examining the relationship between transitory changes in the terms of trade and the current account, this paper uses a three-good (importables, exportables and nontradables) intertemporal model which provides a way to measure the substitution effects arising from shocks to the terms of trade. For example, an adverse, temporary change in the terms of trade will have both income and substitution effects. The income effect (which lowers current national income relative to future national income) induces consumption-smoothing responses as agents try and maintain their level of consumption, and forms the basis for the HLM effect. The substitution effects involve: changes in the temporal pattern of consumption due to a rise in the price of current consumption relative to the price of future consumption (holding constant the path of the real exchange rate); and allowing for intratemporal effects flowing through a rise in the relative price of nontradables (that is, an appreciation of the real exchange rate), which in turn raises the relative price of current consumption. We use a two-step cointegration-Euler equation approach to derive estimates for individual countries of the strength of both of these substitution effects. Third, we undertake an analysis which demonstrates that using the standard single-good, permanent-income model of consumption results in estimates of the elasticity of intertemporal substitution in consumption which are biased downward, in comparison with estimates based on models (such as that used in this paper) which allow for traded and nontraded goods.

We find that movements in the terms of trade are an important cause of variation in the current account position for most of the developed countries in our sample. In addition, shocks to the terms of trade are very persistent for all countries except the United States. We decompose movements in the terms of trade (which is a series stationary in first differences) into its permanent (random walk) and temporary (stationary) components, and find that the temporary component of movements in the terms of trade is large for all countries, accounting for about half of the variance of quarter-to-quarter changes in the terms of trade. Cointegration and generalized method of moments estimation indicate that terms of trade shocks induce large and significant intratemporal and intertemporal substitution effects, which

operate to offset any associated income effects on saving decisions and the current account position. In contrast to previous results derived from models using a single consumption good, we find estimates of the intertemporal elasticity of substitution which are significantly different from zero, and typically exceed unity. This implies that, for example, in the face of a transitory fall in the terms of trade, consumers in developed countries are willing to defer consumption in response to this increase in the real rate of interest. Moreover, our estimates of the intratemporal elasticity of substitution reveal that adverse terms of trade shocks result in a substitution away from tradables (importables) toward nontradables, with the consequent appreciation of the real exchange rate (higher relative price of nontradables) raising the consumption rate of interest and current saving.

Finally, these results highlight the need for caution in using standard consumption-smoothing models of the current account to draw implications for the magnitude of international capital flows in the face of transitory terms of trade shocks. Certainly, transitory terms of trade shocks give rise to income effects and associated HLM-like, consumption-smoothing behavior (given a positive correlation of private saving and the direction of such shocks). However, our results indicate that transitory terms of trade shocks also give rise to important substitution effects. These substitution effects will tend to ameliorate any income-based decline in private saving and deterioration of the current account, and so reduce the magnitude of any consumption-smoothing-based need for international capital flows. While at first glance the relationship between movements in the terms of trade and the current account balance appears weak, underpinning this relationship are terms-of-trade induced substitution and income effects, which produce offsetting changes in private saving decisions and the current account position.

The remainder of the paper is as follows. Section II explains the several channels though which temporary shocks to the terms of trade can affect private saving and the current account. Section III analyzes the importance of terms of trade shocks for our five developed economies. Section IV describes a stochastic, infinite-horizon representative agent model of consumption, where agents maximize utility by consuming both traded and nontraded goods. The data and econometric modeling methodology are described in Section V, while the econometric results are set out in Section VI. Some concluding comments are contained in Section VII.

II. How Does the Terms of Trade Affect the Current Account?

The effects of terms of trade shocks on external balances was first studied by Harberger (1950) and Laursen and Metzler (1950), using single-good, Keynesian open

economy models which assumed an absence of international capital mobility. Given that both consumption and income are measured in terms of exportables and that the marginal propensity to consume is less than unity, a fall in current income arising from an adverse terms of trade shock would lower private saving.[2] This work was extended into the forward-looking framework by Sachs (1981), Obstfeld (1982), Greenwood (1983) and Svensson and Razin (1983), who argued that the HLM effect is only true for temporary terms of trade shocks. While Obstfeld (1982) and Svensson and Razin (1983) used two-good models (importables and exportables) to look at the influence of shocks to the terms of trade on intertemporal decisions, later work by (among others) Edwards (1989), Gavin (1990), Mendoza (1992) and Ostry and Reinhart (1992) incorporated nontradables, and examined the additional effects terms of trade shocks can exert through changes in the real exchange rate (the relative price of tradables and nontradables).[3]

In theory, the impact of the terms of trade on the current account is ambiguous. An adverse, transitory terms of trade shock will have three effects: it will lower current national income relative to future national income (the consumption-smoothing or HLM effect); it will increase the current price of imports relative to the future price of imports (the consumption-tilting effect); and it will increase the price of tradables (importables) relative to the price of nontradables (the real exchange rate effect). In response to an adverse terms of trade shock, private savings will fall (rise) if the consumption-smoothing effect dominates (is weaker than) the saving-enhancing implications of the consumption-tilting and real exchange rate effects; that is, the current account position (for given national investment and government deficit) will worsen if the income effect associated with the terms of trade shock is stronger than the two substitution effects.

Under the HLM effect, the temporary deterioration of a country's terms of trade produces a transitory fall in current real national income relative to the country's permanent real national income—the fall in permanent income is smaller than the fall in current income. In turn, this would induce consumption-smoothing behavior, with agents in the economy seeking to spread their temporarily low income over reduced consumption in many periods. The result is a fall in aggregate saving for the economy because the fall in consumption

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[2] The most commonly-used measure of the terms of trade is the ratio of export prices to import prices, both expressed in a common currency—usually referred to as the commodity or net barter terms of trade. This is also the concept used in this paper. In addition, in the examples given below it is assumed that in the initial equilibrium prior to any terms of trade shock, the current account is balanced.

[3] Sachs (1981) and Sen and Turnovsky (1989) also emphasize the role of capital accumulation in the relationship between the terms of trade and the current account.

(determined by the marginal propensity to spend out of wealth) is less than the temporary fall in income, which deteriorates the current account position.[4]

Apart from the abovementioned income effect on private savings arising from a deterioration in the terms of trade, there will also be relative price (substitution) effects. This paper concentrates on two relative prices in particular: the ratio of national price levels (real exchange rate) and the relative price of imports in terms of exports (the terms of trade). Changes in the relative cost of living in different countries, and in the relative price of countries' exports and imports, both have key roles to play in determining the extent of intertemporal trade (Obstfeld and Rogoff (1996)).

As to the first of the two substitution effects, assuming that agents borrow and lend internationally in terms of the exportable good, an adverse transitory terms of trade shock makes current imports and current consumption more expensive relative to future imports and future consumption, and so should induce agents to tilt (transfer) their consumption into the future—a rise in current aggregate saving. This rise in the consumption rate of interest (higher cost of current consumption relative to future consumption) occurs through the temporarily higher price of imports raising the general level of prices in the current period (Dornbusch (1983)). The intertemporal elasticity of substitution in consumption (σ) determines the extent to which agents switch consumption from high-price to low-price periods, in response to any given change in the consumption rate of interest. The larger σ, the greater is the response of current-period private saving (and the current account) to any temporary shock to the terms of trade. Conversely, when σ is low there is little tendency to readjust consumption in response to movements in intertemporal relative prices.

The second substitution effect occurs when nontraded goods are introduced into the model, as then a transitory adverse terms of trade shock will make consumption of importables more expensive than the consumption of nontradables, causing agents to substitute into nontradables (assuming the two goods are net substitutes) and so raise the relative price of nontradables. This temporary real appreciation also induces a rise in the consumption rate of interest, as the higher relative price of nontradables temporarily raises the general level of prices—making current consumption relatively more expensive, and inducing a rise in current aggregate saving. The intratemporal elasticity of substitution between tradables (importables) and nontradables (ε) determines the extent to which agents switch from importables to home-produced goods, alter the real exchange rate (that is, the relative price of nontradables) and thereby affect the consumption rate of interest. The larger ε, the greater is the response of current-period private saving (and the current account) to any

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[4] For a constant rate of time preference (invariant to the level of welfare) and given level of investment, if an adverse terms of trade shock is expected to be permanent then agents would revise downward their level of consumption *pari passu* with the fall in expected permanent income (assuming that the marginal propensity to save out of permanent income is zero), yielding no change in saving or the current account position (Svensson and Razin (1983)).

temporary shock to the terms of trade (Ostry and Reinhart (1992)). A value of ε above (below) unity implies that importables and nontradables are gross substitutes (complements).

Following an adverse terms of trade shock: the HLM effect implies that private saving should fall (current account deteriorate) as real income declines; conversely, the consumption tilting and real exchange rate effects imply that private saving should rise (current account improve) as real income declines. On balance, the effect of terms of trade shocks on private saving and the current account will be determined by which of these has the greater relative strength. The strength of the substitution effects, in turn, is a function of the intertemporal elasticity of substitution (σ) and the intratemporal elasticity of substitution between importables and nontradables (ε). Following a temporary fall in the terms of trade, the larger are the intertemporal and intratemporal elasticities of substitution, then the smaller will be any consequent consumption-smoothing-induced fall in private saving, and the smaller will be any deterioration in the current account position.[5]

III. How Important are Terms of Trade Shocks?—The Stylized Facts

We first examine the empirical regularities of international relative prices, focussing here on three key variables. The terms of trade (*TOT*), the ratio of the current account balance (in current prices) to output (in current prices), (*CAGDP*); and the rate of growth of real output (*RY*). All three variables have been constructed using data on our sample of five OECD countries (Australia, Canada, New Zealand, the United Kingdom and the United States), with all data taken from the IMF's *International Financial Statistics* database.

There have been sizeable shifts in the net barter terms of trade in the five developed countries considered in this paper over the previous two decades, in tandem with quite volatile movement in their current account positions and in the growth of real output (Figures 1 and 2).[6] This volatility has been particularly evident for the commodity-exporting countries of Australia, Canada and New Zealand.

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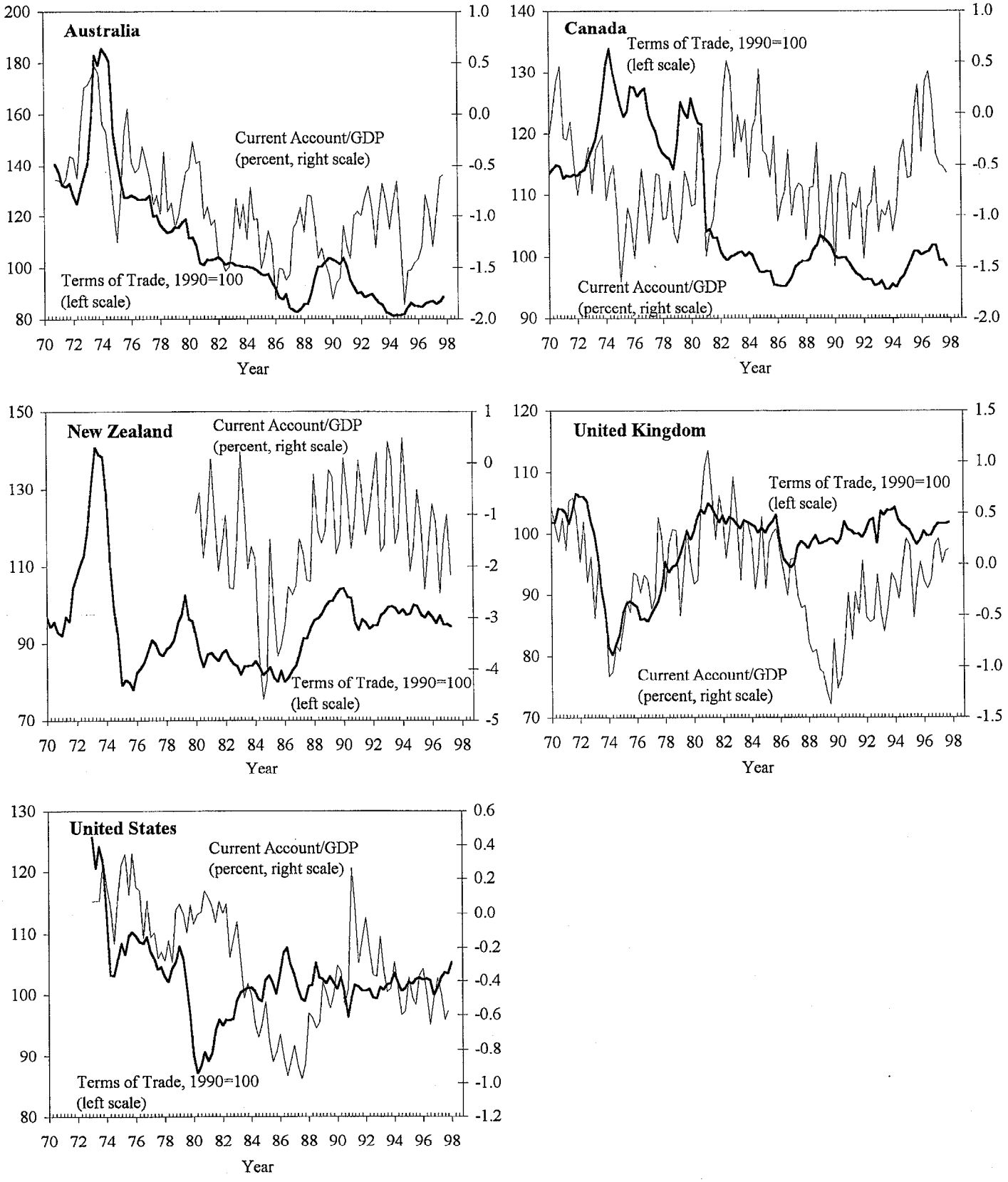
[5] It should be kept in mind that a terms of trade shock is a particular type of shock to an economy's real income. Supply shocks (such as those generated by drought in agricultural-based economies) are another type of shock which lowers an economy's real income. Even in the absence of terms of trade shocks, a country may wish to smooth consumption intertemporally in the face of such temporary supply shocks. However, unlike a negative supply shock in a small open economy, an adverse terms of trade shock has substitution (or relative price) effects. In this case, an adverse terms of trade shock would raise the consumption rate of interest, which reduces real expenditure (raises real saving) in the current period in favor of expenditure in future periods.

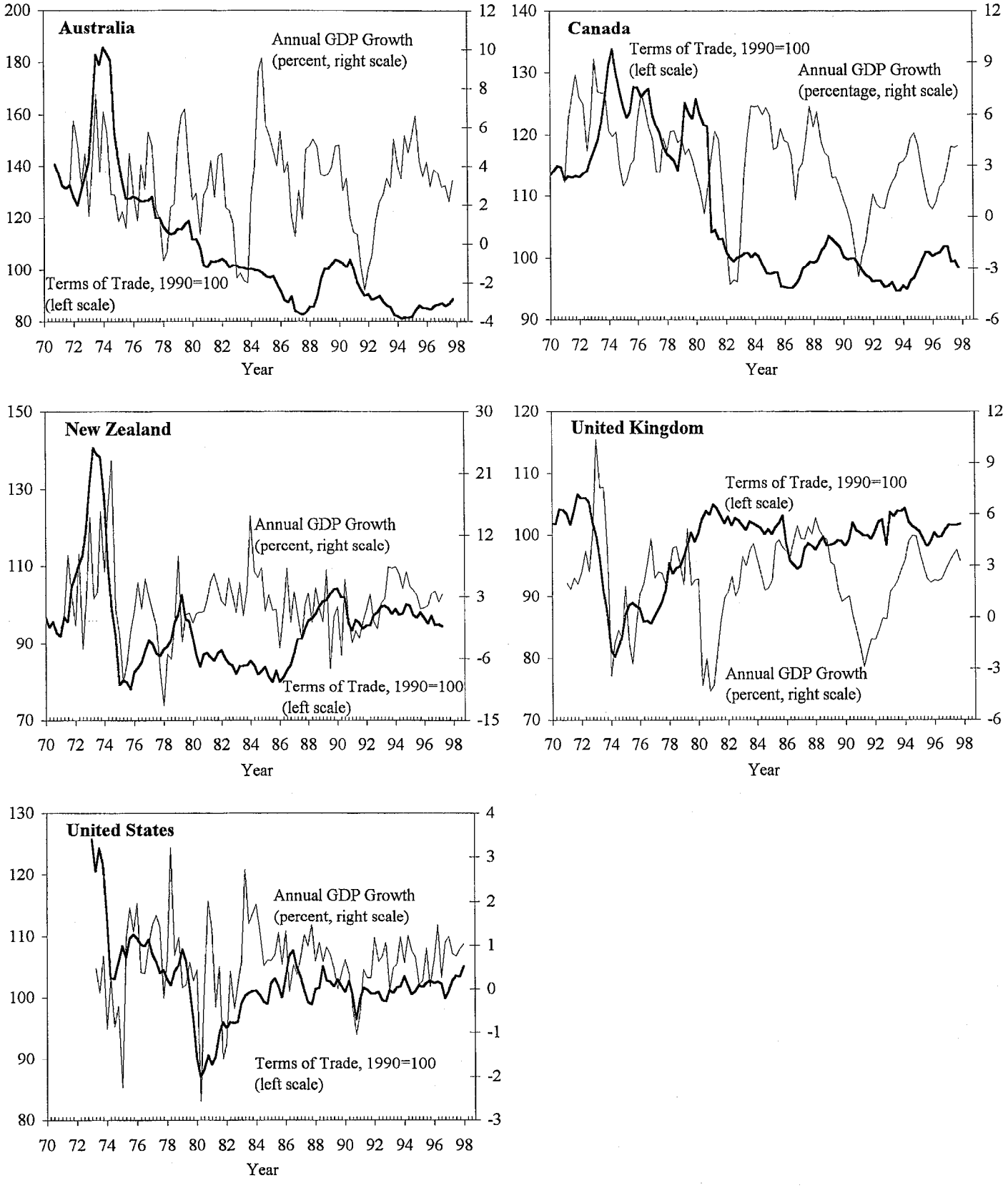
[6] The stylized facts of the terms of trade, including its correlation with net exports, have been previously examined for developed countries by Backus et al. (1992), for developing countries by Mendoza (1992), and for both developed and developing countries by Razin (1995).

Figure 1. Terms of Trade and the Current Account, 1970:1- 1998:1



Source: IMF, IFS. See Appendix for data sources and definition.

Figure 2. Terms of Trade and Real GDP Growth, 1970:1- 1998:1



Source: IMF, IFS. See Appendix for data sources and definition.

Table 1 presents the standard deviation and contemporaneous correlation of the terms of trade and the current account balance. Shocks to the terms of trade are most volatile in the case of Australia, more than double the volatility of the other four countries. In contrast, shocks to the current account position of New Zealand have been far more volatile than in any other country—this is consistent with New Zealand's relatively less diversified export base, in that it specializes in exporting commodities experiencing large relative price changes. The ratio of the two standard errors indicates that the current account is typically far less variable than the terms of trade, yet there is a wide range for the ratio of the variability of the two series, with Australia having the largest (ratio of 16) and New Zealand the smallest (ratio of 1.5) measures of the ratio of the standard deviation of the *TOT* to the standard deviation of *CAGDP*.

The degree of correlation between changes in *TOT* and changes in *CAGDP* is weakly positive for all five countries.[7] This positive (yet small) correlation between changes in *TOT* and changes in *CAGDP* is in line with the HLM effect, implying that income effects (operating through channels such as terms of trade shocks) are stronger than any substitution (consumption-tilting and real exchange rate) effects induced by fluctuations in the terms of trade. However, in all cases except that of the United States, the data cannot reject the null hypothesis that there is zero correlation between changes in *TOT* and changes in *CAGDP*.[8]

A. Contribution of Terms of Trade Shocks to Fluctuations in Current Account Positions

In an attempt to identify the impact of terms of trade shocks on output and the external balance of the five countries, we follow Ahmed and Park (1994) and Otto (1995) in extending the Blanchard and Quah (1989) framework to an open economy setting, using a structural vector autoregression (SVAR) model to identify major sources of economic shocks. Three different shocks are identified: an external shock which is measured by innovations to the terms of trade (*TOT*); a permanent 'supply' shock, which is identified by innovations to the growth of real output (*RY*); and a temporary 'demand' shock, which is identified by innovations to the current account balance as a share of GDP (*CAGDP*). To separately identify these shocks, identifying restrictions are employed to restrict certain long-run multipliers in the structural model. The key restrictions are: first, that the terms of trade is exogenously given; and second, as with Blanchard and Quah (1989), 'demand' shocks are identified using the assumption that such shocks have no long-run effect on the level of real output.

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[7] By weakly positive we mean greater than one standard deviation.

[8] The approximate critical value of the correlation coefficients, computed under the null hypothesis that the true correlation coefficient is zero, is about 0.196 $((1/T^{½})*1.96)$ for T=100 observations.

Table 1. Terms of Trade and the Current Account, 1970-97: Statistical Summary

	Standard Deviation		*Correlation*
	Terms of Trade	Current Account as Share of GDP	
Australia	0.049	0.003	0.110
Canada	0.023	0.004	0.166
New Zealand	0.019	0.011	0.124
United Kingdom	0.018	0.003	0.086
United States	0.024	0.002	0.210*

Notes: For the terms of trade (*TOT*) and the ratio of current account to gross domestic product (*CAGDP*), quarterly data for Australia are from 1970:2-1997:2; Canada (1970:2-1997:4); New Zealand (1980:2-1997:2); the United Kingdom (1970:2-1997:4); and the United States (1973:2-1997:4). The measures of *TOT* and *CAGDP* are in first differences, as they were found to be nonstationary in levels (see Table 2). All data are from IMF, *International Financial Statistics*; see Appendix I for details and definitions. An asterisk (*) indicates that the correlation coefficient is significantly different from zero, given that the approximate critical value of the correlation coefficients, computed under the null hypothesis that the true correlation coefficient is zero, is about 0.196 ($(1/T^{½})*1.96$) for T=100 observations.

The SVAR model (using quarterly data over the period 1970-97) is based on the assumption that *TOT*, *RY* and *CAGDP* are I(1) variables in levels. Results from the Phillips-Perron (1988) unit root test, computed using the Bartlett kernel and with lag lengths determined by the data-dependent method of Andrews (1991) suggest that: for all countries *CAGDP* is I(1) in levels; for all countries *RY* is I(1) in levels; and for all countries *TOT* is also I(1) in levels (Table 2). Accordingly, and consistent with results obtained for developed countries by Backus et al. (1994), there appears to be a good deal of persistence in the terms of trade shocks affecting most of these countries. Our results from the SVAR are also conditioned on the maintained hypothesis of no cointegration among the levels of the variables used in the analysis. If there was cointegration between *RY*, *TOT* and *CAGDP*, then a VAR in first differences would be misspecified. The results of the Phillips-Ouliaris (1990) $Z(t)$ residual-based test for cointegration among the three I(1) variables reveal that for all countries, the null hypothesis of no cointegration cannot be rejected at the five percent level of significance. [9]

Variance decompositions from the SVAR model are presented in Table 3.[10] These reveal that for most countries (except New Zealand) the major source of fluctuations in the current account position is due to demand shocks, with terms of trade shocks also being an important cause of variation. For example, in the case of New Zealand, demand and terms of trade shocks explain, respectively, 29 and 46 percent of the variation in *CAGDP* after 20 quarters. Terms of trade shocks also appear to be important causes of movements in the current account position in the United States and Australia (in both cases explaining about 20 percent of the variation in *CAGDP* after 20 quarters), yet have less influence in the cases of Canada (7 percent) and the United Kingdom (3 percent). For completeness, the results for the variance decomposition of changes in real output (*RY*) are also presented in Table 3. These reveal that after 20 quarters, terms of trade shocks explain little (United Kingdom and New Zealand, less than 10 percent) or some (Australia and the United States, about 20 percent) of

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[9] The critical value for the $Z(t)$ residual-based test of cointegration at the five (ten) percent level of significance is -3.77 (-3.45); values more negative than this imply rejection of the null hypothesis of no cointegration. Results for the $Z(t)$ test statistic (computed using the Bartlett kernel and with lag lengths determined by the data-dependent method of Andrews (1991)) were: Australia (-1.34), Canada (-2.00), New Zealand (-1.18), United Kingdom (-0.83), and United States (0.43).

[10] These measure the percentage of the k-step ahead forecast error variance in a variable which is due to terms of trade, supply and demand shocks. The results presented in Table 3 are based on a SVAR model with eight lags of each variable in each of the three equations.

Table 2. Unit Root Tests

Country	TOT_t	RY_t	$CAGDP_t$	(n_t/m_t)	(p_t/q_t)
Australia					
level	-1.728	2.397	-1.876	0.839	0.345
first difference	-6.802*	-4.565*	-5.661*	-6.017*	-2.989*
Canada					
level	-1.233	-0.511	-2.234	2.242	-1.332
first difference	-8.864*	-3.409*	-8.066*	-3.375*	-6.018*
New Zealand					
level	-2.619	0.460	-1.969	2.375	-1.675
first difference	-6.184*	-7.067*	-6.739*	-5.081*	-7.029*
United Kingdom					
level	-2.127	0.838	-2.065	2.386	-1.791
first difference	-8.113*	-3.197*	-7.667*	-6.181*	-8.251*
United States					
level	-4.001*	1.825	-1.473	1.662	-2.160
first difference	-7.014*	-2.843*	-6.377*	-8.244*	-9.098*

Notes: The Phillips-Perron (PP, 1988) test regressions include an intercept term, and an asterisk (*) indicates that the null hypothesis of a unit root can be rejected at (at least) the 5 percent level of significance (for the PP test the 5 percent critical value is -2.89 (100 observations); -2.93 (50 observations); and -3.00 (25 observations)). Results including a trend term were very similar to those reported above. For the terms of trade (*TOT*), real output (*RY*) and the current account as a share of GDP (*CAGDP*), quarterly data for Australia are from 1970:2-1997:2; Canada (1970:2-1997:4); New Zealand (1980:2-1997:2); the United Kingdom (1970:2-1997:4); and the United States (1973:2-1997:4). For *RY* and *CAGDP*, the unit root tests are carried out on deseasonalized data. For the ratio of consumption of nontradables to the consumption of importables (n/m) and the ratio of the price of importables to the price of nontradables (p/q), annual data for Australia are from 1970-95; Canada (1970-92); New Zealand (1972-93); the United Kingdom (1970-94); and the United States (1970-93). All unit root tests were computed using the Bartlett kernel, and the lag lengths were determined by the data-dependent method of Andrews (1991). See Appendix I for data description and derivation.

Table 3. Structural Vector Autoregression Model: Decomposition of Forecast Error Variance

	Current Account as a Share of GDP			*Real Output*		
	TOT	Supply	Demand	TOT	Supply	Demand
Australia						
1	13.39	7.08	79.53	12.36	64.01	23.63
5	22.12	11.68	66.20	18.17	68.17	13.67
10	20.31	17.17	62.52	18.91	63.75	17.34
20	19.70	17.63	62.66	20.19	62.50	17.31
Canada						
1	0.45	4.60	94.95	22.98	72.73	4.29
5	2.10	5.95	91.96	36.43	59.96	3.61
10	4.72	10.09	85.19	42.62	52.95	4.42
20	6.71	10.52	82.77	42.69	51.57	5.73
New Zealand						
1	59.04	0.23	40.73	4.73	83.94	11.34
5	45.88	22.01	32.11	4.83	82.21	12.97
10	47.41	24.40	28.19	7.74	79.44	12.82
20	46.46	24.51	29.04	9.21	77.50	13.30
United Kingdom						
1	0.16	1.38	98.46	0.12	97.22	2.66
5	2.15	1.72	96.13	1.31	96.31	2.38
10	2.68	2.44	94.88	3.42	93.71	2.87
20	2.82	2.76	94.42	4.08	92.21	3.70
United States						
1	16.66	0.77	82.56	0.98	94.32	4.70
5	18.15	3.73	78.13	16.07	77.96	5.97
10	19.02	5.58	75.40	21.13	72.42	6.44
20	19.91	5.72	74.37	23.58	69.50	6.92

Notes: The decomposition of the variance of the current account as a share of GDP (*CAGDP*) and of the variance of real output (*RY*), is composed of innovations to: the terms of trade (ΔTOT); supply, proxied by real output (ΔRY); and demand, proxied by the current account deficit normalized by GDP ($\Delta CAGDP$), where Δ is the first-difference operator. The number of quarters since the shock are 1, 5, 10 and 20. The error variance was calculated from a structural vector autoregressive (SVAR) analysis, with eight lags of each variable in each of the three equations, using quarterly data on *TOT*, *RY* and *CAGDP* for: Australia from 1970:2-1997:2; Canada (1970:2-1997:4); New Zealand (1980:2-1997:2); the United Kingdom (1970:2-1997:4); and the United States (1973:2-1997:4). For *RY* and *CAGDP*, the SVAR analysis uses deseasonalized data. See Appendix I for data description and derivation.

the variation in *RY*, while Canada appears to be an outlier.[11] In all cases, the major source of fluctuation in *RY* occurs through supply shocks.[12]

B. Persistence of Terms of Trade Shocks

Results obtained using the median-unbiased estimator of Andrews (1993), which corrects for the downward bias imparted by standard unit root estimation, confirm the abovementioned finding—shocks to the *TOT* are very persistent in all countries except the United States (Table 4).[13] For example, for the typical *TOT* shock affecting Australia and Canada, median-unbiased estimates of the autoregressive parameter in the unit root regression have a 90 percent confidence interval between 0.951-1.000 and 0.964-1.000, respectively, with median estimates of unity in both countries. This implies that for both countries it takes an infinite number of quarters for the impulse response of a unit shock to the *TOT* to attain half of its original

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[11] The findings for Canada are consistent with those of Ahmed and Park (1994), who find that in comparison with other small open economies, external shocks contribute little to explaining movements in the Canadian trade balance, yet are important in explaining movements in Canadian real output.

[12] Our results for the variance decomposition of Australia are similar to those obtained by Otto (1995) for the Australian trade balance and for real output, where he found that terms of trade shocks explained 44 percent of the variance in the trade balance and 10 percent of the variance in real output, respectively, after 20 quarters. Our results are also consistent with those of Hoffmaister and Roldós (1997), who find that, for groups of Asian and Latin American countries: external shocks (innovations to the world interest rate and the terms of trade) explain between 10-25 percent of output fluctuations and between 25-30 percent of fluctuations in the trade balance; that most (65 percent) of the variation in output can be attributed to supply shocks; and that most (65 percent) of the variation in the trade balance can be attributed to demand shocks.

[13] Standard least squares estimators of unit root models yield significantly downward-biased estimates of the autoregressive parameter, especially when that parameter is large, as the true distribution of the autoregressive parameter in unit root tests is skewed to the left (resulting in the median exceeding the mean). Moreover, standard methods of testing for a unit root have low power (low probability to reject the null, given the value of the autoregressive parameter) in many cases of empirical interest. This implies that a failure to reject the null hypothesis of a unit root should not be taken as evidence in favor of the null. Accordingly, point and interval estimators (especially median-unbiased ones) will be superior measures of persistence—they are also exact procedures, rather than asymptotic ones; and are robust to non-normal innovations in the unit root model (Andrews (1993)). In addition, estimators produced by this method enable us to calculate how long shocks to a particular series will last.

Table 4. Terms of Trade: Median Unbiased Estimation of Autoregressive Coefficient from an Autoregressive/Unit Root Regression, 1970-97

Country	Estimator	α	HLS
Australia			
(108 obs.)	*OLS*	0.975	
	Median-	1.000	∞
	unbiased	[0.951, 1.000]	[13.903, ∞]
Canada			
(111 obs.)	*OLS*	0.981	
	Median-	1.000	∞
	unbiased	[0.964, 1.000]	[18.967, ∞]
New Zealand			
(109 obs.)	*OLS*	0.949	
	Median-	0.988	59.767
	unbiased	[0.900, 1.000]	[6.569, ∞]
United Kingdom			
(111 obs.)	*OLS*	0.955	
	Median-	0.997	271.190
	unbiased	[0.912, 1.000]	[7.563, ∞]
United States			
(100 obs.)	*OLS*	0.857	
	Median-	0.857	4.477
	unbiased	[0.715, 0.993]	[2.068, 97.001]

Notes: The results of this table are based on model (2) of Andrews (1993, p.143). HLS refers to the half life of a unit shock, and for $\alpha \geq 0$ gives the length of time until the impulse response of a unit shock is half its original magnitude; where α is the autoregressive parameter in the unit root regression and the HLS = $|\log(½)/\log(\alpha)|$. The entries in the rows labeled OLS are the least squares estimates of α; the entries in the rows labeled median unbiased are the median-unbiased estimates of α and HLS. The intervals in square brackets under the median-unbiased estimates are the 90 percent central confidence intervals. The median-unbiased estimators and exact confidence intervals given in this table were derived using quarterly data for Australia from 1970:2-1997:2; for Canada (1970:2-1997:4); for New Zealand (1980:2-1997:2); for the United Kingdom (1970:2-1997:4); and for the United States (1973:2-1997:4). See Appendix I for data description and derivation.

magnitude (Table 4).[14] [15] *TOT* shocks are also extremely persistent for New Zealand and the United Kingdom, taking 60 and 271 quarters, respectively, for the impulse response of a unit shock to the *TOT* to attain half of its original magnitude. In contrast, shocks to the *TOT* of the United States have a half-life of only 4 quarters. These results for the persistence of terms of trade shocks are supported by the stylized facts set out in Table 1—countries with more persistent terms of trade shocks (Australia, Canada, New Zealand and the United Kingdom) are the ones which exhibit relatively less correlation between the terms of trade and the current account.

C. How Large Are the Permanent and Temporary Components of Movements in the Terms of Trade?

Any series which is stationary in first differences (has a unit root, I(1)) can be represented as a combination of stationary and random walk components (Beveridge and Nelson (1981)). Importantly, the persistence of shocks to a particular series does not imply anything about the extent to which fluctuations in that series are typically temporary or typically permanent in nature. Fluctuations in any given series will be partly temporary (captured by the stationary component) and partly permanent (captured by the random walk component), and we can ask how important the temporary or stationary component is to the behavior of the series. That is, how large is the variance of shocks to the stationary component of the *TOT* compared with the variance of quarterly changes in the *TOT*? Following Cochrane (1988), we measure the size of the random walk component of *TOT* from the variance of its long differences. Given that *TOT* is I(1), then the plot of $(1/k)\text{var}(TOT_t - TOT_{t-k})$ should approach the variance of the shock to the random walk (permanent) component. We find that $1/k$ times the variance of k-differences settles down (at about 0.138) to 23 percent of the variance of first differences (about 0.6) for the case of the United States after k=50 quarters, suggesting that the innovation variance of the random walk (permanent) component of *TOT* is about one-quarter of the variance of quarter-to-quarter changes in the *TOT*—there is a large temporary component (explaining about three-quarters) of movements in the U.S. *TOT* (Figure 3). The result for the United Kingdom is similar, with the innovation variance of the random walk component of *TOT* accounting for about one-fifth of the variance of quarter-to-quarter changes in the *TOT*. In contrast, the results for Australia and New Zealand imply that the innovation variance of the random walk (permanent) component of *TOT* is about 45 and 60 percent, respectively, of the variance of quarter-to-quarter changes in the *TOT*—there is a smaller temporary component (explaining about 55 and 40 percent,

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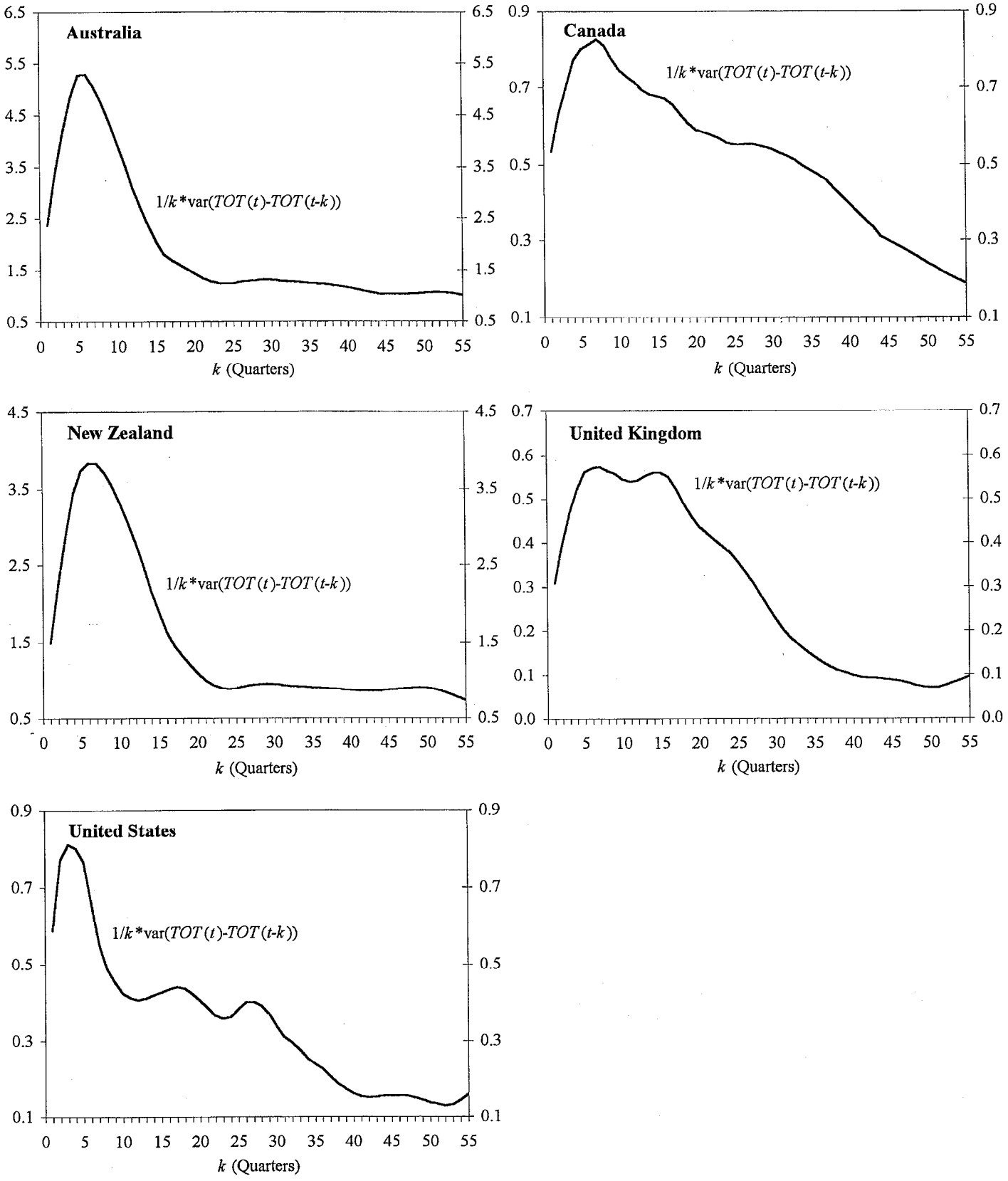
[14] The half life of a unit shock (for $\alpha \geq 0$) is denoted by HLS, and gives the length of time until the impulse response of a unit shock is half its original magnitude; α is the autoregressive parameter in the unit root regression, and HLS = $|\log(½)/\log(\alpha)|$. See Andrews (1993) and the notes to Table 4 for further details.

[15] A finding of infinite persistence of any given time series is equivalent to a finding of a unit root.

Figure 3. Temporary and Permanent Components of Movements in the Terms of Trade, 1970:1- 1998:1



Source: Authors' calculations. See Appendix for data sources and definitions.

respectively) of movements in their *TOT*. This result is consistent with countries having a relatively large traded goods sector experiencing a greater share of permanent shocks to their *TOT*. Finally, the result for Canada also demonstrates the clear dominance of temporary shocks to its terms of trade. Accordingly, we proceed with our empirical analysis of the implications of transitory shocks to the terms of trade, having found that in all five countries changes in the terms of trade have a large temporary component.

D. Implications of Terms of Trade Shocks for Intertemporal Models of the Current Account

In recent years the empirical literature on determinants of the current account has yielded mixed evidence as to the usefulness of the consumption-smoothing approach to the current account. While this approach has been successful in explaining movements in the current account of the United States (Ghosh (1995)) and moderately successful for Australia (Cashin and McDermott (1998a, 1998b)), it has not worked well for the United Kingdom (Obstfeld and Rogoff (1996)) or Canada (Ghosh (1995)). A key element of empirical tests of the consumption-smoothing approach has been the assumption of a single good, which allows for income effects of changes to permanent income arising from movements in the terms of trade, but precludes any substitution effects arising from shocks to the terms of trade. In particular, these studies were based on a consumption-smoothing model which assumed: that the home country and the rest of the world produced goods that were physically identical (so there was no direct role for terms of trade effects); a constant real interest rate (so there was no role for consumption-tilting effects); and assumed there were no transport costs, deeming that all goods were tradable across countries (so excluding nontraded goods and a role for movements in the real exchange rate).

However, the above analysis reveals that terms of trade shocks are an important driver of movements in the current account position of most of the five developed countries. Accordingly, the predictive performance of consumption-smoothing models of the current account could be affected by the existence of important consumption-tilting and real exchange rate effects on private saving induced by temporary terms of trade shocks, both of which are excluded from standard smoothing models. One implication, to which we turn our attention in Section VI of the paper, is that there may be important consumption-tilting and real exchange rate effects influencing the determination of the current account, which we examine using a theoretical and empirical model which allows for these substitution effects.

IV. The Intertemporal Model of Consumption

Following Ostry and Reinhart (1992), in deriving the stochastic intertemporal model of the consumption consider an economy composed of a large number of infinitely-lived consumers, where aggregate consumption is a CES function of the consumption of tradables (importables) and nontradables, and each consumer maximizes a constant intertemporal elasticity of substitution utility function of the form

$$U = [\sigma/(\sigma-1)]E_0\sum_{t=0}^{\infty}\beta^t\left(\omega m_t^{1-1/\varepsilon}+n_t^{1-1/\varepsilon}\right)^{\frac{1-1/\sigma}{1-1/\varepsilon}}, \qquad \omega,\ \beta,\ \varepsilon,\ \sigma > 0, \beta < 1, \qquad (1)$$

where E_0 is the expectations operator conditional on information at time 0, m denotes consumption of importables, n denotes the consumption of nontradables, β is the subjective discount factor, ω is the weight attached to the imported good in the intratemporal utility function, ε is the intratemporal elasticity of substitution between importables and nontradables (the extent to which consumers alter their consumption of importables in response to a change in its price relative to that of nontradables), and σ is the intertemporal elasticity of substitution (the extent to which consumers defer current consumption in response to a higher expected real return). The model assumes perfect capital mobility, and so the country faces a given world real interest rate. The numeraire is assumed to be the exportable good; accordingly, importables, nontradables and interest rates are all measured in units of the exportable.

Let b_t be the economy's stock of net external liabilities at the beginning of period t, r_t the world real interest rate, $\bar{m}_t$ is the endowment of tradable (importable) goods, $\bar{n}_t$ is the endowment of nontradable goods, x_t is the export of tradable goods (exportables), and p_t and q_t denote the relative price of importables and nontradables, respectively. The consumer's budget constraint is then

$$\Delta b_{t+1} = r_t b_t + p_t(m_t-\bar{m}_t) + q_t(n_t-\bar{n}_t) - x_t. \qquad (2)$$

Maximizing (1) subject to (2) yields[16]

$$E_t\left\{\frac{(1+r_t)p_t}{p_{t+1}}\left[\frac{\omega m_{t+1}^{1-1/\varepsilon}+n_{t+1}^{1-1/\varepsilon}}{\omega m_t^{1-1/\varepsilon}+n_t^{1-1/\varepsilon}}\right]^{\frac{\sigma-\varepsilon}{\sigma(\varepsilon-1)}}\left[\frac{m_{t+1}}{m_t}\right]^{-1/\varepsilon}\right\} = \frac{1}{\beta} \qquad (3)$$

$$E_t\left\{\frac{(1+r_t)q_t}{q_{t+1}}\left[\frac{\omega m_{t+1}^{1-1/\varepsilon}+n_{t+1}^{1-1/\varepsilon}}{\omega m_t^{1-1/\varepsilon}+n_t^{1-1/\varepsilon}}\right]^{\frac{\sigma-\varepsilon}{\sigma(\varepsilon-1)}}\left[\frac{n_{t+1}}{n_t}\right]^{-1/\varepsilon}\right\} = \frac{1}{\beta} \qquad (4)$$

and

$$\omega(n_t/m_t)^{1/\varepsilon} = (p_t/q_t). \qquad (5)$$

In equations (3) and (4), the terms $((1+r)p_t/p_{t+1})$ and $((1+r)q_t/q_{t+1})$ take account that the real rate of interest for intertemporal consumption decisions is the world real interest rate, adjusted

[16] It is also assumed that the transversality condition holds, so that at the limit, the discounted value of real net external liabilities equals zero.

for the rate of change of the relative price of importables (equation (3)) or nontradables (equation (4)). That is, the consumption and world real rates of interest will differ from one another when either the terms of trade (relative price of importables) or real exchange rate (inverse of relative price of nontradables) is expected to change over time.[17] This model considers temporary, current changes in the terms of trade.[18] Equations (3) and (4) are stochastic Euler equations which equate the expected gain in utility from the consumption of an additional unit of the importable or nontradable good in the future to the marginal cost of foregoing the consumption of an additional unit today (that is, the intertemporal relative price). Equation (5) states that the intratemporal marginal rate of substitution between importables and nontradables should be equated to their relative prices.

V. Data and Estimation Method

The representative agent model outlined above was estimated using time series data for the five developed countries, and the coverage of the data spans the period 1970-95 (see Appendix I for additional details on the sources and description of the data). In estimating equations (3) to (5), while much of the data required can be derived from various standard sources, data on the consumption of tradables (importables) and nontradables is not typically available, even for most developed countries. Accordingly, we follow the earlier literature (Ostry and Reinhart (1992), Goldstein and Officer (1979)) in assuming that: the consumption of nontradables comprises the consumption of import substitutes plus imports of consumer goods; that all domestic production of import substitutes is consumed locally; that domestic production of import substitutes is domestic production of traded goods (assumed to be the output of the agricultural, mining and manufacturing sectors of the economy) less exports (which are assumed not to be consumed domestically); and that the consumption of nontradables equals the production of nontradables (GDP less the output of the agricultural, mining and manufacturing sectors of the economy). To account for the fact that import barriers have reduced through time in the majority of the five countries, we define the price of traded goods as p, where $p = dm(1+((dut/m)/100))$, dut is the amount of import duties, dm is

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[17] The terms of trade is defined as the relative price of importables in terms of exportables; a rise in the price of importables denotes a deterioration of the terms of trade. The real exchange rate is defined as the inverse of the relative price of nontradables in terms of exportables; a rise in the price of nontradables denotes a real appreciation.

[18] In contrast, a permanent deterioration in the terms of trade would: induce no income-based consumption-smoothing behavior (that is, no HLM effect); and assuming constant expenditure shares between importables and nontradables, there would be no change in the consumption rate of interest and so no intertemporal substitution effect. However, a permanent deterioration in the terms of trade would permanently alter the relative price of nontradables, and so the intratemporal substitution (real exchange rate) effect would operate.

the deflator for imports, and m is imports of goods (national accounts basis)—see Appendix I for details.[19]

In estimating the model we use a two-step procedure similar to Cooley and Ogaki (1996), which combines a cointegration-based estimate of preference parameters with generalized method of moments (GMM)-based estimation of Euler equations.[20] First, we can test whether (n_t/m_t) and (p_t/q_t) are cointegrated to obtain estimates of ω and ε from equation (5). Given that our estimated equation (5) is a cointegrated regression then our estimates of ω and ε will be super-consistent, and so we can impose in the subsequent estimation of equation (3) the values of ω and ε, increasing the efficiency of the estimated parameters. Second, we can estimate equation (3) using Hansen's (1982) generalized method of moments (GMM), to obtain an estimate of the additional parameters of interest (β and σ). The six instrumental variables used in the GMM estimation include a constant, n_{t-1}/n_{t-2}, m_{t-1}/m_{t-2}, $(1+r_{t-2})p_{t-2}/p_{t-1}$, m_{t-1}, and n_{t-1}, where there are two free parameters and accordingly four overidentifying restrictions.[21]

VI. Empirical Results

Using Phillips-Perron (1988) unit root tests, we find that (n_t/m_t) and (p_t/q_t) are integrated of order one, and so the possibility of cointegration exists (Table 2). The results of Phillips-Ouliaris (1990) $Z(t)$ residual-based tests for cointegration reject the null hypothesis of no cointegration between (n_t/m_t) and (p_t/q_t) at the 10 percent significance level for Australia, and at the 5 percent significance level for all other countries (Table 5). In estimating equation (3), the calculated Hansen (1982) J-test statistics for all model runs are small, indicating that

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[19] This adjustment is made as import prices need to take account of customs duties in order to correctly measure the relative price of importables; for most of the countries in our sample (Australia, New Zealand and Canada), the intertemporal variation in customs duties has been large because tariffs have been reduced considerably over time. Between the early 1970s and mid-1990s, import duties as a share of imports have fallen from about 12 to 5 percent for Australia, 6 to 3 percent for Canada, 8 to 4 percent for New Zealand, 3 to near zero percent for the United Kingdom, and 6 to 3 percent for the United States (IMF, *Government Finance Statistics*; see also Appendix I).

[20] It is important to note that the system of equations (3), (4), and (5) are not independent, so we only need to estimate two of these equations.

[21] Given rational expectations, the forecast error which form part of the residual of equation (3) will be orthogonal to any instrument known to economic agents at time t. This allows us to use many instruments to estimate a few parameters, and yields a set of overidentifying restrictions to test the usefulness of the model. The instrument set excludes variables measured at time t, as otherwise the moving average process in the error term would engender correlation between these variables and the residual.

Table 5. Cointegration and GMM Estimates of the Euler Equations, 1970-95

	ε	σ	β	ω	J	$Z(t)$
Australia	1.646	1.146	0.928	3.459	5.779	-3.118
	(0.178)	(0.338)	(0.016)	(0.101)	(0.216)	
Canada	4.102	1.476	0.910	3.353	3.588	-3.461
	(0.525)	(0.589)	(0.022)	(0.101)	(0.465)	
New Zealand	0.699	2.652	0.968	28.697	3.510	-6.031
	(0.082)	(1.768)	(0.035)	(0.309)	(0.476)	
United Kingdom	1.676	2.237	0.938	3.521	0.569	-6.372
	(0.265)	(0.202)	(0.018)	(0.176)	(0.966)	
United States	5.628	0.722	0.884	1.509	0.197	-6.668
	(5.925)	(0.344)	(0.043)	(0.244)	(0.995)	

Notes: The parameter estimates have been obtained from a cointegrating regression of equation (5), which yields estimated values for ω (weight on the imported good in the utility function) and ε (intratemporal elasticity of substitution). These values are then inserted into equation (3), which is estimated by generalized method of moments (GMM) to yield estimated values for σ (intertemporal elasticity of substitution) and β (subjective discount rate). Standard errors are in parenthesis under each of the parameter estimates—in columns 1 and 4 the standard errors are robust to serial correlation and heteroscedasticity, and are computed using the Bartlett kernel.

Data for the cointegrating and GMM regressions are annual, and for Australia are from 1970-95; Canada (1970-92); New Zealand (1972-93); the United Kingdom (1970-94); and the United States (1970-93). See Appendix I and notes to Table 2 for data description and derivation.

J is the Hansen (1982) J-statistic for the test of the validity of the overidentifying restrictions, which is distributed as a $\chi^2(4)$; the associated p-value at which the null hypothesis that the overidentifying restrictions can be rejected is given below it in parentheses. $Z(t)$ is the Phillips-Ouliaris (1990) residual-based cointegration test—the 5 percent (10 percent) critical value for this test is -3.37 (-3.07); values more negative than this imply rejection of the null hypothesis of no cointegration. The Phillips-Ouliaris (1990) cointegration test was computed using the Bartlett kernel, and lag lengths were determined by the data-dependent method of Andrews (1991).

the overidentifying restrictions imposed by the model are not rejected by the data (Table 5). That is, the two parameters estimated satisfy the six orthogonality conditions implied by the instrument set (and upon which the model estimates are assumed to hold), in that the addition of extra instruments does not alter the value of J very much.

A. Estimates of the Subjective Discount Rate

In common with previous work, we find that for all countries the subjective discount rates are robustly estimated, with $\hat{\beta}$ being well identified and falling in the range 0.884-0.968. Interestingly, New Zealand economic agents appear to discount future consumption much more heavily than their counterparts in the other countries in our sample (Table 5).

B. Estimates of the Intertemporal Elasticity of Substitution

In contrast to the result of Hall (1988), we find that the intertemporal elasticity of substitution is statistically different from zero (0.722) for the United States, and is so for all other countries except New Zealand (Table 5). For Australia, Canada and the United Kingdom, the estimate of σ is found to be 1.146, 1.476 and 2.237, respectively.[22] This indicates that in these countries, when faced by a transitory shock to the terms of trade, agents will change their private saving decision by a large amount, offsetting any consumption-smoothing response to movements in the terms of trade. Moreover, these results are larger than those for developing countries obtained by of Ostry and Reinhart (1992), which is consistent with the more open capital accounts of these five developed countries.[23] Our higher developed-country estimates of σ are also consistent with previous empirical findings that poor consumers (and countries) tend to have a lower intertemporal elasticity of substitution than do rich consumers (and countries), because those expenditure-inelastic goods (such as subsistence and necessary (food) goods) which dominate the expenditure of poor consumers and countries are less substitutable through time than are expenditure-elastic goods (Atkeson and Ogaki (1996)).

Interestingly, while the estimate of σ for New Zealand is not significantly different from zero, it is clearly the largest coefficient of the five countries, indicating a willingness to transfer consumption across periods. As a check of the consumption-tilting implications of our estimates of σ, we estimated the consumption-tilting parameter (γ) from a cointegrating

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[22] Such values imply that $\hat{\theta}$, the coefficient of relative risk aversion (which is the inverse of $\hat{\sigma}$), lies in the range 1.385-0.447, which are consistent with previous findings in the literature that economic agents in more developed countries are typically less risk averse (have lower θ) than agents in developing countries (Hu (1993), Hahm (1998)). The smaller is θ (larger is σ), then the more slowly marginal utility falls as consumption rises, indicating the agents are more willing to allow their consumption to vary over time.

[23] However, these results are consistent with Ostry and Reinhart's (1992) findings that more developed regions have higher estimates of σ.

regression of national cash flow (output net of investment and government consumption expenditure) net of payments on external liabilities on private consumption (Ghosh (1995), Cashin and McDermott (1998a, 1998b)).[24] This parameter ($\hat{\gamma}$) also provides an indication of the willingness of a country to tilt its consumption towards the present or future, and is driven by differences between the subjective discount rate and the world real interest rate. The results of the Phillips-Ouliaris (1990) $Z(t)$ residual-based test for cointegration are given in Table 7, and reveal that the null hypothesis of no cointegration between national cash flow net of payments on external liabilities and private consumption is rejected for all countries (except Canada and the United Kingdom) at the 5 percent level of significance. Consistent with our findings for the estimates of σ, these estimates of γ reveal that of the five countries, New Zealand clearly has the most pronounced tendency to tilt consumption across periods.

In addition, these findings are not consistent with a common result in the literature which uses a linearized version of the Euler equations, namely that of no relationship between the real interest rate and changes in consumption. This literature finds that the intertemporal elasticity of substitution in consumption is not significantly different from zero for developing countries (Giovannini (1985)) or for the United States (Hall (1988), Campbell and Mankiw (1989), Patterson and Pesaran (1992)).[25] Ostry and Reinhart (1992) for developing countries, Patterson and Pesaran (1992) for the United Kingdom, and Ogaki and Reinhart (1998) and Hahm (1998) for the United States are notable exceptions in finding that the consumption-savings allocation decision responds to changes in expected real interest rates.[26]

Notwithstanding these findings of an intertemporal elasticity significantly different from zero, most researchers have used a single-equation, linearized version of the Euler equations estimated here (equations (3) and (4)), and assumed that there was a single consumption good—no allowance was made for potential differential effects of changes in

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[24] Phillips-Perron (1988) unit root tests reveal that both national cash flow net of payments on external liabilities and private consumption are integrated of order one for all countries, and so the possibility of cointegration exists (Table 6).

[25] Under the standard expected utility framework, such a value for the elasticity of intertemporal substitution implies an implausibly large (infinite) degree of relative risk aversion. See Hahm (1998) for a summary of previous work on United States consumption.

[26] Ostry and Reinhart (1992) found that for a panel of 13 developing countries, σ' estimated by an equation similar to equation (3) was significantly different from zero, yet less than unity. A similar result was found by Ogaki and Reinhart (1998) in estimating the intertemporal elasticity of substitution between consumer durables and nondurables in the United States.

Table 6. Unit Root Tests

Country	$c_t \equiv (n_t+m_t)$	R_t	pc_t	$(y_t-i_t-gc_t)-rb_t$
Australia				
level	-1.074	-2.702	8.694	4.805
first difference	-5.369*	-14.072*	-5.646*	-5.066*
Canada				
level	-2.230	-5.698*	1.858	2.866
first difference	-9.688*	-34.749*	-4.219*	-4.803*
New Zealand				
level	-3.205*	-8.529*	0.176	-0.547
first difference	-22.189*	-48.089*	-6.795*	-12.163*
United Kingdom				
level	-3.553*	-13.632*	1.957	3.432
first difference	-31.254*	-75.929*	-3.629*	-4.769*
United States				
level	-4.418*	-17.833*	4.115	3.981
first difference	-37.837*	-96.983*	-5.137*	-4.029*

Notes: The Phillips-Perron (PP, 1988) test regressions include an intercept term, and an asterisk (*) indicates that the null hypothesis of a unit root can be rejected at (at least) the 5 percent level of significance (for the PP test the 5 percent critical value is -3.00 (25 observations). Variable *c* denotes aggregate consumption (the consumption of nontradables, *n*, plus tradables, *m*); *R* denotes the real interest rate (yield on three-month national Treasury bills, net of domestic inflation ($\Delta\ln(def)$), where *def* is the GDP deflator, for all countries except New Zealand (where *R* is the government bond yield net of domestic inflation)); *pc* denotes private consumption; and ((*y-i-gc*)-*rb*) is national cash flow (*ncf* = output (*y*) less aggregate investment (*i*) less government consumption (*gc*)), less interest payments on net external liabilities (*rb*). For c ≡ (*n*+*m*) and *R*, data for Australia are from 1970-95; for Canada (1970-92); for New Zealand (1972-93); for the United Kingdom (1970-94); and for the United States (1970-93). For private consumption (*pc*) and national cash flow less interest payments on net external liabilities ((*y-i-gc*)-*rb*), data for all five countries Australia are from 1950-97. All unit root tests were computed using the Bartlett kernel, and the lag lengths were determined by the data-dependent method of Andrews (1991). See Appendix I for data description and derivation.

Table 7. Estimates of the Single-Good, Linearized Euler Equation and Consumption-Tilting Parameter

	σ'	μ	γ	$Z(t)$
Australia	-0.008	0.112	0.910	-3.499
	(0.003)	(0.008)	(0.010)	
Canada	-0.013	0.125	0.959	-2.946
	(0.005)	(0.019)	(0.009)	
New Zealand	-0.004	0.105	0.883	-5.124
	(0.004)	(0.017)	(0.022)	
United Kingdom	-0.009	0.120	0.983	-2.845
	(0.003)	(0.010)	(0.013)	
United States	-0.003	0.072	0.959	-3.624
	(0.004)	(0.011)	(0.007)	

Notes: The parameter estimates for columns 1 and 2 have been obtained from an instrumental variables (IV) regression of equation (6), that is, $\Delta c \equiv \Delta(n+m)$ on R and a constant (the change in aggregate consumption on the real interest rate and a constant), which yields estimated values for μ (constant term) and σ' (intertemporal elasticity of substitution). Consistent with Hall (1988), the instrument set used for all IV regressions was R_{t-2}, Δc_{t-2} and $\Delta(\ln(def_{t-2}))$. $\hat{\gamma}$ is the estimate of the consumption-tilting parameter, derived from the cointegrating regression of national cash flow net of interest payments on external liabilities, $((y\text{-}i\text{-}gc)\text{-}rb)$, on private consumption, (pc). Standard errors are in parentheses under each of the parameter estimates.

For the OLS regression of $\Delta c \equiv \Delta(n+m)$ on R and a constant, annual data for Australia are from 1970-95; Canada (1970-92); New Zealand (1972-93); the United Kingdom (1970-94); and the United States (1970-93). See Appendix I and notes to Table 6 for data description and derivation.

$Z(t)$ is the Phillips-Ouliaris (1990) residual-based test of cointegration between private consumption (pc) and national cash flow net of interest payments on external liabilities $((y\text{-}i\text{-}gc)\text{-}rb)$—the 5 percent (10 percent) critical value for this test is -3.37 (-3.07); values more negative than this imply rejection of the null hypothesis of no cointegration. For the cointegrated regression of pc on $((y\text{-}i\text{-}gc)\text{-}rb)$, data for Australia are from 1951-97; for Canada (1950-97); for New Zealand (1954-96); for the United Kingdom (1950-97); and for the United States (1950-94). The Phillips-Ouliaris (1990) cointegration test was computed using the Bartlett kernel, and lag lengths were determined by the data-dependent method of Andrews (1991) See Appendix I and notes to Table 6 for data description and derivation.

real interest rates on the consumption of traded and nontraded goods.[27] The typical stochastic real interest rate version of the permanent income model considered a single consumption good of the form

$$\Delta c_t = \mu + \sigma' R_t + e_t, \tag{6}$$

where Δc_t is the first-difference of the (natural logarithm) of aggregate consumption (here the sum of consumption of importable and nontradable goods), R_t is the real interest rate; e_t is a random disturbance term; μ is a constant term; and the estimated coefficient on the real interest rate, σ', is the intertemporal elasticity of substitution.[28] Using instrumental variables regression to account for possible contemporaneous correlation between the disturbance term and the regressors, we estimated equation (6) for the five countries in our sample. We found that while in most countries (all except New Zealand and the United States) there was a statistically significant relationship between changes in consumption and the real interest rate, this was very close to zero (Table 7).[29] For New Zealand and the United States, consistent with Hall (1988), Campbell and Mankiw (1989), and Patterson and Pesaran (1992), σ' was estimated to be not statistically different from zero. These results are consistent with early studies based on the aggregate consumption function, which typically found a low interest

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[27] Estimates of σ' derived from the single-consumption-good literature are likely to be inaccurate for countries which are subject to frequent shocks to their terms of trade that alter their real exchange rates and, accordingly, change the relative price of importables and nontradables. Somewhat counterintuitively, the assumption of a single consumption good appears to be reasonably valid for New Zealand (see Table 5). See Carroll (1997) for a discussion of the deficiencies associated with estimating log-linear consumption Euler equations.

[28] Most recent studies of the relation between aggregate consumption and asset returns have used a particular definition of consumption, typically either consumption of nondurable goods or of nondurables plus services. Following Hu (1993) we use aggregate consumption in our estimation of equation (6), as the definition of durables and nondurables is often arbitrary in most official statistics, and it is inappropriate to exclude consumer durables expenditure (which is the component of consumption most likely to be affected by movements in interest rates).

[29] As noted by Patterson and Pesaran (1992), Hall's (1988) model (and our equation (6)) predicts that the sign of the estimate of σ', the coefficient on the real interest rate, will be positive—higher real interest rates will encourage agents to defer consumption into the future.

elasticity of savings (Hall (1988)).[30] Accordingly, our results indicate that in estimating the key influences on consumer preferences, it is important to relax the restrictive assumption that there is only one type of consumer good, and specifically allow for the consumption of both tradables and nontradables. In effect, once the influence on consumption of tradables and nontradables caused by a change in their relative prices is controlled for, then the real rate of interest (the relative price of consumption today versus consumption tomorrow) does become a key determinant of intertemporal consumption patterns—periods of high real interest rates are then associated with periods of rapid growth in consumption. Not allowing for the differential influence of the real interest rate on the consumption of tradables and nontradables biases estimates of the intertemporal elasticity of substitution toward zero.[31]

C. Estimates of the Intratemporal Elasticity of Substitution

The estimated parameter for the intratemporal elasticity of substitution between tradables (importables) and nontradables (ε) is statistically significant, positive and greater than unity for Australia, Canada and the United Kingdom, indicating that the consumption of importables and nontradable are substitutes (Table 5). Agents in these countries, when faced with a transitory adverse shock to their terms of trade, substitute away from relatively expensive importables and consume relatively inexpensive home-produced goods (nontradables), inducing a consequent appreciation of the real exchange rate, an increase in the consumption rate of interest and a rise in private saving to offset the fall in private saving arising from the HLM effect.[32] For the United States, our estimate of ε is not statistically significant. However, our finding for New Zealand that importables and nontradables are complements ($\hat{\varepsilon}<1$) indicates that in response to a transitory adverse shock to the terms of trade, agents will not be able to substitute away from relatively expensive importables and consume relatively inexpensive nontradables, and so will not induce a consequent rise in private saving to offset the fall in private saving arising from the HLM effect. In the New Zealand case, the substitution effects ($\hat{\sigma}$ not statistically different from zero and $\hat{\varepsilon}<1$)

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[30] Using U.S. postwar data on consumption, Hall (1988) contradicts Summers' (1982) finding of an intertemporal elasticity of substitution of about one, finding an estimate of σ' of 0.10 (not significant), which is close to that of this paper. As with this paper and unlike Summers (1982), Hall (1988) used appropriate instruments in his estimation (those uncorrelated with the disturbances): the change in consumption, the rate of inflation and the real interest rate (all lagged two periods).

[31] Ogaki and Reinhart (1998) find a similar result in estimating the intertemporal elasticity of substitution between consumer durables and nondurables in the United States. They find that ignoring the intratemporal substitution between durables and nondurables results in a misspecification bias which yields negative point estimates of the intertemporal elasticity of substitution.

[32] Our results for these OECD countries are in line with the findings of Ostry and Reinhart (1992), particularly that more developed regions exhibit lower intertemporal substitution.

reinforce the saving-inhibiting income effect arising from the terms of trade shock, and the current account unambiguously deteriorates.

In summary, the parameter estimates obtained here indicate that for Australia, Canada and the United Kingdom, consumption-tilting and real exchange rate effects of transitory terms of trade shocks would operate to offset any changes in private saving (and the current account) induced by the consumption-smoothing (HLM) effect. In contrast, and not unexpectedly, shocks to the terms of trade are a less important determinant of private saving decisions for the United States, probably due to its much smaller dependence on external trade. In the case of New Zealand, there is little evidence of an offsetting rise in private saving flowing from the terms-of-trade-induced consumption-tilting or real exchange rate effects. Such findings are consistent with New Zealand's high ranking among our sample countries in its period-average current account deficit and end-period net external liabilities, which for the period 1970-97 and at end-1995 were, respectively, (in percent of GDP): New Zealand (-5.51 and 76), Australia (-3.63 and 59), Canada (-2.42 and 44), the United Kingdom (-0.45 and -5) and the United States (-1.04 and 11).

VII. Conclusion

What happens to an economy's current account position if there is an adverse shock to its terms of trade? According to the Harberger-Laursen-Metzler (income) effect, an adverse transitory movement in the terms of trade results in a decrease in a country's current level of income which is larger than the decrease in its permanent income, causing a fall in aggregate saving and a deterioration of the current account position. The Harberger-Laursen-Metzler effect assumes that the home country and the rest of the world produce the same good, and that this good is tradable across countries.

However, in a three-good (importable, exportable and nontradable) setting in which countries consume both importables and nontradables, then substitution effects will also influence private saving decisions. These substitution effects concern the ease with which, in response to a terms of trade shock, countries can switch: between importables and nontradables (intratemporal substitution), and between current and future consumption in response to a shift in the relative price of current consumption (intertemporal substitution).

Using data for five commodity-exporting developed countries for the period 1970-95, cointegration and generalized method of moments estimation indicate that terms of trade shocks induce large and significant intratemporal and intertemporal substitution effects, which operate to offset any associated income effects on saving decisions and the current account position. In contrast to previous results derived from models using a single consumption good, we find estimates of the intertemporal elasticity of substitution which are significantly different from zero, and typically exceed unity. Moreover, our estimates of the intratemporal elasticity of substitution reveal that adverse terms of trade shocks result in a substitution away from importables toward nontradables, with the consequent appreciation of the real exchange

rate (higher relative price of nontradables) raising the consumption rate of interest and current saving.

These results highlight the need for caution in using traditional consumption-smoothing models of the current account to draw implications for the magnitude of international capital flows in the face of transitory terms of trade shocks. Transitory terms of trade shocks do give rise to income effects and associated consumption-smoothing behavior (given a positive correlation of private saving and the direction of such shocks). However, our results indicate that transitory terms of trade shocks also give rise to important substitution effects. These substitution effects will tend to ameliorate any income-based decline in private saving and consequent deterioration of the current account position, and so reduce the magnitude of any consumption-smoothing-based need for international capital flows. While at first glance the relationship between movements in the terms of trade and the current account balance appears weak, underpinning this relationship are terms-of-trade induced substitution and income effects, which produce offsetting changes in private saving decisions and the current account position.

Sources and Description of the Data

Quarterly data for the structural VAR and Andrews (1993) analyses are taken from the International Monetary Fund's *International Financial Statistics* database, generally for the period 1970-1997. Data on real output in local currency is real GDP (line 99b.r); the terms of trade, formed by the ratio of the index of exports prices (line 76 or line 74) to the index of import prices (line 76.x or line 75); the ratio of the current account balance to nominal GDP, formed by the ratio of the U.S. dollar-denominated current account balance (line 78ald), converted to local currency (for all countries except the U.S.) using the end-of period exchange rate to the U.S. dollar (line ag or line ae), divided by nominal GDP in local currency (line 99b.c). The nominal GDP data are seasonally adjusted at annual rates. New Zealand data on nominal GDP for 1980:2-1988:2 is from the OECD's *Analytical Database*. Data for Australia are from 1970:2-1997:2; for Canada are from 1970:2-1997:4; for New Zealand are from 1980:2-1997:2; for the United Kingdom are from 1970:2-1997:4; and for the United States are from 1973:2-1997:4.

Annual data for the GMM estimation are as follows. *International Financial Statistics* (International Monetary Fund): r (three-month nominal U.S. Treasury bill rate); *def* (GDP deflator); R (real interest rate, formed as rr (the nominal interest rate on local three-month Treasury bills) minus the rate of change of *def*); y (nominal GDP at market prices); *pop* (national population); *Government Finance Statistics* (International Monetary Fund): *dut* (import duties); *World Economic Outlook* (International Monetary Fund): *dm* (deflator for imports); *Analytical Database* (Organization for Economic Cooperation and Development): *amm* (GDP derived from agriculture (including hunting, forestry and fishing), mining (including quarrying) and manufacturing); x (exports of goods, national accounts basis); m (imports of goods, national accounts basis); *dc* (personal consumption deflator); *International Trade Statistics Yearbook* (United Nations): s (share of consumer good imports in total merchandise imports).

Using the above annual series, additional data were constructed as follows. Domestic production of traded goods (national accounts basis) is *amm*; domestic production of import substitutes (national accounts basis) is *dms*, where $dms = amm - x$; imports of consumer goods (converted to a national accounts basis) is *mcg*, where $mcg = s.m$; per capita consumption of importable goods (national accounts basis) is m, where $m = (mcg + dms)/pop$; the price of importable goods is p, where $p = dm(1+((dut/m)/100))$; per capita consumption of nontraded goods (national accounts basis) is n, where $n = (y - amm)/pop$; the price of nontraded goods is q, where $q = dc$. In calculating m, it is assumed that exports are not domestically consumed and that all domestically-produced import substitutes are consumed locally; p is used as the deflator. In calculating n, it is assumed that the consumption of nontraded goods equals the production of nontraded goods; q is used as the deflator. All consumption data are converted to a per capita basis by dividing the aggregates by the national population (*pop*). The world interest rate is given by r. Data are for the following

countries and sample periods: Australia (1970-95), Canada (1970-92), New Zealand (1972-93), United Kingdom (1970-94), and United States (1970-93).

Data for the cointegration analysis of Section VI are taken from the *Analytical Database* (Organization for Economic Cooperation and Development): *n* and *m*; and from *International Financial Statistics* (International Monetary Fund): *y*; *pc* (private consumption spending); *i* (aggregate investment, comprising gross fixed capital formation plus the change in stocks) and *gc* (government consumption spending). Data for Australia are from 1951-97; for Canada (1950-97); for New Zealand (1954-96); for the United Kingdom (1950-97); and for the United States (1950-94).

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