A Structural VAR Approach to Test the Present Value Model of the Current Account

Takashi Kano∗
Department of Economics, University of British Columbia

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Abstract

The purpose of this paper is to test the present value model of the current account. Instead of the existing present value test framework, a structural VAR approach is introduced to test the model. The present value model predicts different responses of the current account measure to different structural shocks. By exploiting two kinds of identification schemes based on the small open country assumption and Blanchard and Quah’s(1989) long-run restriction, this paper identifies global permanent, country-specific permanent and country-specific transitory shocks. The theoretical responses of the current account measure to these shocks are tested by the identified structural VMA. A puzzling characteristic of current account movements in the proto-type small open countries, Canada and the U.K., is revealed by the empirical results: current account movements in these countries depend excessively on country-specific transitory shocks.

Key Words: The Current Account; Structural VAR; Identification.
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1 Introduction

The purpose of this paper is to test the present value model of the current account (PVM). This paper introduces a structural VAR (SVAR) approach in order to test the model's predictions on the responses of the current account measure to structural shocks. With identification schemes implied by the model's assumptions, the SVAR approach in this paper not only can directly test the predictions but also provides an inference for an interesting and important question on current account movements: what shocks are main driving forces of the current account?

Sachs (1981, 1982) originally develops the intertemporal current account model by stressing dynamic optimization behavior of forward-looking rational agents. Assuming perfect capital mobility and incomplete international financial markets with the fixed world real interest rate, this model explains current account movements by two motives: the consumption-smoothing motive and the consumption-tilting motive. An important implication of this model is that the consumption-smoothing component has a closed-form solution well-known as the PVM. The PVM characterizes the optimal consumption-smoothing component of the current account as the expected present discounted value of negative changes in net output.

Sheffrin and Woo (1990), Otto (1992) and Ghosh (1995) study the PVM of the current account with different versions of the present value test (PVT). The purpose of the PVT is to test the following predictions of the PVM. Firstly, the consumption-smoothing component of the current account should Granger cause the first difference of net output. Secondly, when one constructs an unrestricted bivariate VAR with the consumption smoothing component of the current account and the first difference of net output, the PVM imposes cross-equation restrictions on the estimated coefficient matrices of the unrestricted VAR. According to the empirical results of the standard PVT literature, the PVM is statistically rejected in many small countries, while the U.S. data provides support. This result is in fact a puzzle because the PVM is based on the small open country assumption.

This paper provides an alternative test for the PVM by investigating the predictions about the responses of the current account measure to structural shocks. As Sachs (1981), Razin (1993) and Glick and Rogoff (1995) argue the PVM predicts that the current account should respond to a shock depending on two characteristics of the shock: whether the shock is global or country-specific and how persistent is the shock? The first characteristic is crucial since the PVM

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1 Obstfeld and Rogoff (1995) provide an excellent survey for the intertemporal current account model. The approach has an explicit precursor in work on the small open optimal growth model by Hamada (1966).

2 Net output is defined as output net of investment and government expenditure. Ghosh (1995) interprets this variable as national cashflow.

3 These authors apply for the current account literature the testing methodology originally developed by Campbell (1987) and Campbell and Shiller (1987) to test the theories of consumption and stock price.

4 To obtain the consumption-smoothing component, for example, Sheffrin and Woo and Otto simply assume the subjective discount rate to be equal to the world real interest rate. In this case, the consumption-smoothing component exactly matches the current account because the optimal consumption follows a perfectly smoothed path (i.e. a random walk). On the other hand, Ghosh derives the consumption-smoothing component by using the model's prediction that the component should be stationary.
predicts that if all countries are homogeneous a global shock does not matter for the current account. No global shock yields an opportunity for the representative consumer in a small open country to change her consumption profile by adjusting international net asset position. This is simply because all countries react to a global shock symmetrically under the homogeneity assumption. Therefore the current account should not respond to a global shock, while country-specific shocks should be significant to current account movements.

The reason the second characteristic is important stems from the PVM’s theoretical aspect that consumption is determined on the basis of the permanent income hypothesis (PIH). On one hand, if a country-specific shock raises net output temporarily, consumption rises by less than an increase in current net output. Hence by the current account identity, the current account responds positively to a country-specific transitory shock. On the other hand, if a country-specific shock raises net output permanently, consumption rises by greater than an increase in current net output. Therefore the current account responds negatively to a country-specific permanent shock.

The PVT framework is not a useful approach to test these predictions, because it provides no scheme to identify the structural shocks. By contrast, the SVAR approach in this paper relies on identification schemes based on the maintained assumptions of the PVM. This allows me to recover the impulse response functions (IRFs) and the forecast error variance decompositions (FEVDs) of the current account with respect to the structural shocks. The recovered IRFs and FEVDs can be then used to check the PVM’s predictions and observe relative importance across the structural shocks on current account movements.

To test the PVM’s predictions, this paper focuses on three structural shocks: global permanent, country-specific permanent and country-specific transitory shocks to net output. In section 2, I introduce stochastic variations in world real interest rates into the basic PVM. The extension of the basic PVM is crucial for this paper’s identification strategy as discussed below. This section derives the linear-approximated present value relation among the current account-net output ratio, the first difference of log net output and the world real interest rate.

In section 3, I show the theoretical responses of the current account measure to three structural shocks. To do so, I characterize the data generating processes (DGP) of log net output and the world real interest rate. In this paper I assume that log of net output is decom-

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5More precisely, if net output follows a unit root process, a country-specific shock raises net output permanently by the same amount. In this case, as shown by Sachs(1981), the current account does not respond to the shock since both consumption and current net output rise by the same amount. Therefore when net output follows a more persistent process than a unit root, like an ARIMA process, the PVM predicts a negative response of the current account to a positive, country-specific shock.

6As a preceding paper, Nason and Rogers(2001) investigate the theoretical restrictions implied by the intertemporal current account model on the joint dynamic behavior of investment and the current account by exploiting the SVAR. On the other hand, this paper originally applies the SVAR approach to test the PVM.

7Bergin and Sheffrin extend the basic PVM by introducing stochastic variations not only in world real interest rates but also terms of trade. The authors find that the PVT cannot reject their extended PVM even in the small open countries like Australia and Canada.
posed into three components; the first is attributed to global permanent shocks, the second to country-specific permanent shocks and the third to country-specific transitory shocks. This decomposition makes it possible to separate permanent and transitory shocks since by construction a country-specific transitory shock has no long-run effect on the log of net output. This is consistent with Blanchard and Quah’s(1989) approach to SVAR identification.

The small open assumption specifies the DGP of the world real interest rate. This assumption implies that country-specific shocks do not matter for the world real interest rate\(^8\). Hence the underlying shocks of the world real interest rate should be global, while a country-specific shock must be orthogonal to the world real interest rate at any forecast horizons. This requirement in turn can be used to identify global and country-specific shocks. Therefore inclusion of stochastic variations in the world real interest rates into the basic PVM is crucial in the sense that it provides identification of global and country-specific shocks.

The specified DGPs and the linear-approximated present value relation yield the structural MA representation of the current account-net output ratio, which characterizes the theoretical responses of the current account measure to three structural shocks. The model’s predictions this paper tests are given as follows. Firstly, if all countries are homogeneous a global shock does not matter for the current account at any forecast horizons. Secondly, the response of the current account-net output ratio to a country-specific shock should depend on the persistency of the shock and be given as the difference between the impact responses of log of net output and consumption to the shock.

Section 4 shows that the PVM in this paper has a structural VMA (SVMA) representation with three endogenous variables, the world real interest rate, the first difference of log of net output and the current account-net output ratio. A novelty of this paper is to show that the theoretical predictions can be restated as cross-equation restrictions on the SVMA. Furthermore it is shown that the PVM provides two schemes to just-identify the SVMA system. The first scheme is based on the lower-triangular long-run matrix as in Blanchard and Quah(1989). The second scheme is Gali’s(1992) approach that uses the impact and long-run restrictions in concert\(^9\). By investigating two different identification schemes, I examine the robustness of the empirical results for identification.

Section 5 describes the data, SVAR estimation and testing method. In particular, this paper studies quarterly data of four G-7 countries: Canada, Japan, the U.K. and the U.S. After estimating the SVMA, it is shown that the cross-equation restrictions given in section 4 can be

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\(^8\)The orthogonality of country-specific shocks to the world real interest rate is a frequently used assumption in the small open RBC literature. For example, see Mendoza(1991), Correia, Neves and Rebelo(1995) and Nason and Rogers(2000). On the other hand, this assumption is recently criticized by Baxter and Crucini(1993) and Blankenau, Kose and Yi(2001). In particular, Baxter and Crucini find that the assumption is empirically indefensible. Since the small open assumption is one of the maintained hypotheses of the PVM, I follow this assumption in order to test the predictions of the model.

\(^9\)It is worth while noting that these identification schemes depend on the assumptions of the PVM, but not on the hypotheses themselves. This fact makes the testing framework of this paper valid.
tested by the asymptotic Wald test.

Section 6 reports the empirical results. This paper emphasizes the robustness of the SVAR estimates for identification. The identification-robust observations of this paper are summarized as follows. First, the IRFs of the current account show that the predictions of the PVM are supported qualitatively: the estimated directions of the IRFs are consistent with the theoretical predictions. Second, the predictions of the PVM are rejected quantitatively in Canada, the U.K. and the U.S.: the cross-equation restrictions are jointly rejected in these countries. In particular, the rejections in the proto-type small open countries, Canada and the U.K., are mainly caused by the failures of the predicted responses of their current account measures to country-specific transitory shocks: in these countries, the responses of the current account-net output ratio to country-specific transitory shocks are too large to support the PVM’s prediction. Finally, the FEVDs reveal that in the small open countries the variations in the current account are dominated by the country-specific transitory shock in not only the short-run but also the long-run. In their SVAR approach, Nason and Rogers (2001) also reports the persistent dependence of the current account on country-specific shocks across the G-7 countries. As they argue, at present there is no consensus intertemporal model that generates persistence in the current account to country-specific transitory shocks. In summary, I uncover a new puzzle of the current account: current account movements in the small open economies depend excessively on country-specific transitory shocks.

The discussion of this paper is closed by conclusive remarks in section 7.

2 The Model

This paper considers a world that consists of many small open countries. Following Glick and Rogoff (1995), I assume that all the countries are homogeneous with respect to preference, endowment and technology. Furthermore, the international financial market is assumed to be incomplete in the sense that no household can buy state-contingent claims to diversify away country-specific shocks. Only riskless bonds, which are denominated in terms of the single consumption good, are traded internationally.\(^\text{10}\)

Consider an infinitely lived representative consumer in a representative small country. The assumption of the small open country implies that this country faces the world real interest rate \(r_t\) that is determined by the international bond market. The standard PVMs, for example Sheffrin and Woo (1990), Otto (1992) and Ghosh (1995), assume the world real interest rate to be fixed. Instead, this paper specifies the world real interest rate to be stochastic, as Bergin and Sheffrin (2000) do. The reason for the extension is that this paper exploits stochastic vari-

\(^{10}\)Incompleteness in the international financial market is one of the maintained assumptions in the intertemporal current account models [see, for example, Obstfeld and Rogoff (1995) and Glick and Rogoff (1995)] and the small open real business cycle models [see, for example, Mendoza (1991) and Cardia (1991)]. By contrast, the two-country real business cycle models [see, for example, Backus, Kehoe, and Kydland (1992) and Baxter and Crucini (1993)] assume the complete financial market. In this literature, the agents in two countries can pool all idiosyncratic risks by trading any contingent claims.
ations in world real interest rates to identify global and country-specific shocks. In addition, this paper assumes the world real interest rate is covariance stationary.

Let $C_t$ be consumption at period $t$, $u(C)$ be the instantaneous utility function of the consumer, and $\beta$ be the subjective discount factor taking a value between 0 and 1, respectively. The consumer’s expected lifetime utility function at period $t$ is then given as

$$
E_t \sum_{i=0}^{\infty} \beta^i u(C_{t+i})
$$

(1)

where $E_t$ is the conditional expectation operator upon the information set at period $t$. Further defining $B_t$, $Q_t$, $I_t$ and $G_t$ to be the international bond holding, output, investment and government expenditure at period $t$, respectively, I give the consumer’s instantaneous budget constraint as follows:

$$
B_{t+1} = (1 + r_t)B_t + Q_t - I_t - G_t - C_t.
$$

(2)

The optimization problem of the representative consumer is then to maximize eq.(1) subject to eq.(2). The first order conditions of this problem are given by the budget constraint (2), the Euler equation

$$
u'(C_t) = \beta E_t (1 + r_{t+1}) u'(C_{t+1}),
$$

(3)

and the transversality condition

$$
\lim_{i \to \infty} E_t R_{t,i} B_{t+i} = 0
$$

(4)

where $R_{t,i}$ is the ex post market discount factor at period $t$ for period $t+i$ consumption, which is defined as

$$
R_{t,i} \equiv \begin{cases} 
1/(\prod_{j=t+1}^{t+i} (1 + r_j)) & \text{if } i \geq 1, \\
1 & \text{if } i = 0.
\end{cases}
$$

(5)

For simplicity, let $NO_t$ denote output net of investment and government expenditure at period $t$: $NO_t \equiv Q_t - I_t - G_t$. Taking the infinite sum of the consumer’s budget constraint (2), and using the transversality condition (4), I can obtain the ex ante intertemporal budget constraint of the consumer as follows:

$$
\sum_{i=0}^{\infty} E_t R_{t,i} C_{t+i} = (1 + r_t)B_t + \sum_{i=0}^{\infty} E_t R_{t,i} NO_{t+i}.
$$

(6)

To obtain the present value representation of the current account measure, I take a log-linear approximation of the Euler equation (3) and a linear approximation of the intertemporal budget constraint (6)\textsuperscript{11}. The approximation begins by dividing the intertemporal budget constraint

\textsuperscript{11}Bergin and Sheffrin(2000) also consider linear approximation of the intertemporal current account model in order to involve stochastic variations of world real interest rates and terms of trade into the present value test literature. While they follow Huang and Lin’s (1993) log-linear approximation, this paper develops an alternative linear approximation to derive a closed form solution of the optimal current account-net output ratio.
After several steps of simple calculation, \( \text{eq.(6)} \) can be rewritten as

\[
\frac{C_t}{NO_t} \left[ 1 + \sum_{i=1}^{\infty} E_t \exp \left\{ \sum_{j=t+1}^{t+i} (\Delta \ln C_j - \ln(1 + r_j)) \right\} \right] = \exp\{\ln(1 + r_t) - \Delta \ln NO_t\} \frac{B_t}{NO_{t-1}} \left[ 1 + \sum_{i=1}^{\infty} E_t \exp \left\{ \sum_{j=t+1}^{t+i} (\Delta \ln NO_j - \ln(1 + r_j)) \right\} \right].
\]

Let \( c, b, \gamma^c, \gamma \) and \( \mu \) denote the means of the consumption-net output ratio \( C_t/NO_t \), the net foreign asset-net output ratio \( B_t/NO_{t-1} \), the first log difference of consumption \( \Delta \ln C_t \), the first log difference of net output \( \Delta \ln NO_t \) and log of gross world real interest rate \( \ln(1 + r_t) \), respectively. \( \text{Eq.(6)} \) is then linearly approximated by taking a first-order Taylor expansion around these means. Appendix A shows the steps of the linear approximation of the intertemporal budget constraint in detail. For any variable \( X_t \), let \( \tilde{X}_t \) denote deviation from its mean value. The linear-approximated intertemporal budget constraint is given as

\[
\tilde{C}_t \approx \frac{1 - \alpha}{\kappa} \tilde{B}_t NO_{t-1} + \frac{1 - \alpha}{\kappa} \ln(1 + r_t) - \frac{1 - \alpha}{\kappa} b \Delta \ln NO_t - c \sum_{i=1}^{\infty} \alpha^i E_t \left\{ \Delta \ln C_{t+i} - \ln(1 + r_{t+i}) \right\} + \frac{1 - \alpha}{1 - \kappa} \sum_{i=1}^{\infty} \kappa^i E_t \left\{ \Delta \ln NO_{t+i} - \ln(1 + r_{t+i}) \right\}. \tag{7}
\]

where \( \alpha = \exp(\gamma^c - \mu) < 1 \) and \( \kappa = \exp(\gamma - \mu) < 1 \). \(^{12}\)

Notice that \( \text{eq.(7)} \) is a linear function of the process of consumption growth rate \( \{\Delta \ln C_{t+i}\}_{i=1}^{\infty} \). To characterize the process of consumption growth, the Euler equation (3) is approximated log-linearly. Suppose that the instantaneous utility function is given as a power function

\[
u(C) = \frac{C^{1-1/\sigma}}{1-1/\sigma}
\]

where \( \sigma \) is the elasticity of intertemporal substitution. This specification of the utility function yields the Euler equation

\[
1 = \beta E_t \left\{ \left( \frac{C_{t+1}}{C_t} \right)^{-\frac{1}{\sigma}} (1 + r_{t+1}) \right\}.
\]

\(^{12}\)The conditions \( \alpha < 1 \) and \( \kappa < 1 \) are required to satisfy boundedness of expected PDV terms of eq.(7). In the following analysis, I impose these conditions: the mean growth rates of consumption and net output are lower than the mean of world real interest rates, respectively. Intuitively, these conditions imply that on the balanced growth path the economy is \textit{dynamically efficient}.
As shown by Campbell and Mankiw (1989) and Campbell (1993), when the world real interest rate and consumption are jointly conditionally homoscedastic and log-normally distributed, the above Euler equation can be rewritten as

\[ E_t \Delta \ln C_{t+1} = \delta + \sigma \ln \beta + \sigma E_t \ln(1 + r_{t+1}) \]

\[ = \delta + \sigma (\ln \beta + \mu) + \sigma E_t \ln(1 + r_{t+1}) - \mu \]  \hspace{1cm} (8)

where \( \delta \) is a constant term including the variances of \( \Delta \ln C_{t+1} \) and \( \ln(1 + r_{t+1}) \) and the covariance between the two terms.\(^{13}\)

Finally to derive an approximated solution of the current account-net output ratio, recall the current account identity

\[ CA_t \equiv r_t B_t + NO_t - C_t \] \hspace{1cm} (9)

By assuming that the economy possesses a balanced growth path, \( \alpha = \kappa \), and using the approximation \( \ln(1 + r_t) \approx r_t \), Appendix B shows that eqs.(7), (8) and (9) together give the present value representation of the current account-net output ratio:

\[ \frac{CA_t}{NO_t} = b_r t + [(\sigma - 1)c + 1] \sum_{i=1}^{\infty} \kappa^i E_t \tau_{t+i} - \sum_{i=1}^{\infty} \kappa^i E_t \Delta \ln NO_{t+i}. \] \hspace{1cm} (10)

Eq.(10) is the desired schedule of the current account-net output ratio, which is represented as a linear present value relation among the current account ratio, the first log difference of net output and the world real interest rate.

Eq.(10) says that the optimal current account ratio is determined by three factors. The third term of the RHS is the consumption smoothing factor, which implies that the representative consumer changes the current account ratio to smooth consumption for expected changes in future growth rates of net output. The second term shows a consumption tilting factor due to expected variations of world real interest rates. If world real interest rates are expected to change in future, the representative consumer desires to alter consumption beyond its smoothed level. The coefficient \( (\sigma - 1)c + 1 \) on the second term implies three different effects to which the consumption tilting factor is attributed: the intertemporal substitution effect (shown by \( \sigma c \)), the income effect (shown by \(-c\)) and the wealth effect (shown by 1).\(^{14}\) Since the mean of

\(^{13}\)It is important to note from the log-linearized Euler equation (8) that perfect consumption smoothing as in previous studies is not the case in this model. First of all, unless \( \delta + \sigma (\ln \beta + \mu) = 0 \), log of consumption has a deterministic trend, as shown by the first two constant terms in the RHS of (8). Secondly, the last term shows that the substitution effect of variations of world real interest rates on the consumption profile. A rise in the world real interest rate makes current consumption more expensive in terms of future consumption. Hence the representative consumer is induced to shift consumption toward the future with elasticity \( \sigma \). These two effects together produce consumption profile that deviates from perfectly smoothed one.

Furthermore, a caveat of the log-linearized Euler equation (8) is that it only cares about first moments of logs of consumption and the world real interest rate. Higher moments of two series are assumed to be fixed.

\(^{14}\)Obstfeld and Rogoff (1996, chapter 1) gives a clear explanation about three effects of the real interest rate change on the current account by two period setting. In particular, the wealth effect comes from the change in permanent net output caused by an world real interest rate change.
the consumption ratio, $c$, is expected to take a value close to one, the intertemporal substitution effect might be the dominant component of the consumption tilting factor. In this case, if world real interest rates are expected to rise, the representative consumer wants to lower consumption below its smoothed level and raise the current account ratio to substitute consumption toward the future. Finally, the first term of the RHS exhibits another consumption tilting factor, which is due to the change in net interest payment from abroad caused by the world real interest rate change. For example, a rise in the world real interest rate increases net interest payment from (to) abroad if the country is a net creditor (debtor). This change in net interest payment prompts the consumer to alter the current account ratio beyond its consumption smoothing level.

What should be noted is that eq.(10) just mentions the country’s desired schedule of the current account ratio. In particular, with the homogeneity assumption across countries, if changes in the world real interest rate and the first log difference of net output are caused by global shocks, there is no opportunity for the representative consumer to intertemporally alter her consumption profile by changing her international asset position. This is because all countries react to a global shock symmetrically and there is no gain of trade across countries. Therefore, as argued by Razin(1993) and Glick and Rogoff(1995), the current account should not respond to a global shock and all that occurs is that the world real interest rate adjusts. On the other hand, a small open country adjusts the current account ratio to a country-specific shock as predicted by eq.(10)\(^{15}\). However because of the small country assumption the world real interest rate should not respond to a country-specific shock.

3 Derivation of the Predicted Responses

This paper focuses on three orthogonal shocks to net output, global permanent, country-specific permanent and country-specific transitory shocks, and tests the predicted responses of the current account ratio to these shocks by the SVAR approach. As discussed in the last section, the distinction between global and country-specific shocks is crucial for testing the PVM because of the model’s prediction that a global shock does not matter for the current account in a small open country. Furthermore decomposing country-specific shocks into permanent and transitory components is important for the discussion of Glick and Rogoff’s(1995) puzzle. Glick and Rogoff find that the estimated impact of a country-specific, permanent productivity shock on the current account is too small to support a fundamental cross-equation restriction of the intertemporal current account model\(^{16}\). More recently, Hoffmann(1999) provides an excellent resolution of the puzzle by applying the permanent - transitory decomposition of Quah(1990). As Quah shows in his resolution for the excess smoothness puzzle of consumption, Hoffmann argues that if country-specific shocks have both important permanent and transitory components, and no identification exists to separate them empirically, the observed responses of the

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\(^{15}\)The small open assumption makes it possible for the representative consumer in a country to borrow and lend as much as she wants given the world real interest rate.

\(^{16}\)More precisely, they observe that the estimated impact of a country-specific, permanent productivity shock on investment tends to be two or three times larger than on the current account, while their model predicts the inverse relationship between the current account and investment.
current account to country-specific shocks are likely to reflect an amalgam of responses to permanent and transitory shocks. In this paper I take into account Hoffmann’s insight on the Glick and Rogoff puzzle in order to avoid the mixed response problem and estimate precise responses of the current account ratio to both permanent and transitory country-specific shocks.

To characterize these shocks, I need to specify the data generating processes (the DGPs) of net output and the world real interest rate. As regards the process of net output, it is well known that the unit root null in net output cannot be rejected by the standard unit root tests. If net output follows a nonstationary process, as shown by Beveridge and Nelson (1981), the process of net output can be decomposed into a stochastic trend term and a transitory term. In this paper I assume that log of net output is decomposed into three components, $\ln NO_t$, $\ln NO_{gp}^t$, $\ln NO_{cp}^t$ and $\ln NO_{cs}^t$, which satisfy

$$\ln NO_t = \ln NO_{gp}^t + \ln NO_{cp}^t + \ln NO_{cs}^t.$$  

The components $\ln NO_{gp}^t$ and $\ln NO_{cp}^t$ are nonstationary and attributed to global permanent shock $\epsilon_{gp}^t$ and country-specific permanent shock $\epsilon_{cp}^t$, respectively. On the other hand, the component $\ln NO_{cs}^t$ is covariance stationary and attributed to a country-specific transitory shock $\epsilon_{cs}^t$. More precisely, it is assumed that for an index $i = \{gp, cp, cs\}$ the first difference of $\ln NO_i^t$ follows an infinite order MA process given as

$$\Delta \ln NO_i^t = c_i + \Gamma_i^{\ln_NO}(L)\epsilon_i^t$$  

where $c_i$ is a constant and $\Gamma_i^{\ln_NO}(L)$ is an invertible, infinite-order polynomial with respect to the lag operator $L$, in which the impact coefficient $\Gamma_i^{\ln_NO}(0)$ does not necessarily equal to one.

Notice that since $\ln NO_{cs}^t$ is covariance stationary no shock has a permanent effect on $\ln NO_{cs}^t$. This implies that the accumulated long-run response of $\Delta \ln NO_i^t$ to $\epsilon_i^t$ should be zero. In other words, it should be the case that

$$\Gamma_{cs}^{\ln_NO}(1) = 0.$$  

Eq.(12) is one of the identification restrictions in this paper. As originally investigated by Blanchard and Quah (1989), the long-run restriction (12) makes it possible to separate permanent and transitory shocks. That is to say, this paper considers a transitory shock as a shock that does not affect net output permanently. The assumed process (11) then implies that $\Delta \ln NO_t$ follows a covariance stationary process such as

$$\Delta \ln NO_t = \Gamma_{gp}^{\ln_NO}(L)\epsilon_{gp}^t + \Gamma_{cp}^{\ln_NO}(L)\epsilon_{cp}^t + \Gamma_{cs}^{\ln_NO}(L)\epsilon_{cs}^t.$$  

where again $\Gamma_{cs}^{\ln_NO}(1) = 0$ from eq.(12).

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18Note that eq.(11) is a maintained structural MA representation of the process $\Delta \ln NO_i^t$, rather than the Wold representation with the impact coefficient equal to one. Instead of being restricted to one, the impact coefficient is estimated as shown below.

19In their seminar paper, Blanchard and Quah define a demand shock as a shock that does not affect output permanently and impose the long-run restriction on the accumulated long-run response of the first difference of output to a demand shock.
The next task is to characterize the process of the world real interest rate. To do so, this paper emphasizes the small open assumption. This assumption requires that a small open country have no influence on the world real interest rate. In other words, a country-specific shock does not matter for the world real interest rate at any forecast horizon. In this paper I exploit this maintained assumption to separate global and country-specific shocks. That is, I assume that the global permanent shock to net output, $\epsilon_{gp}^p$, is a unique shock that can affect the world real interest rate. Hence, the world real interest rate is assumed to follow the infinite-order MA process

$$r_t = c_r + \Gamma^r(L)\epsilon_{gp}^p$$

where $c_r$ is a constant and $\Gamma^r(L)$ is an invertible, infinite-order polynomial, in which the impact coefficient $\Gamma^r(0)$ does not necessarily equal to one. From this specification, the demeaned series of the world real interest rate, $\tilde{r}_t$, is simply given as

$$\tilde{r}_t = \Gamma^r(L)\epsilon_{gp}^p.$$ (14)

Notice that the world real interest rate is determined at the market-clearing level where aggregate excess demand for the riskless bond is zero. Therefore a global shock affects the world real interest rate through its effect on aggregate excess demand in the international bond market. Under the homogeneity assumption across countries, every country has the same excess demand for international riskless bonds. In this case, no country can alter its net foreign asset position to a global shock because all the other countries react to the shock symmetrically. Therefore a global shock has no effect on the current account at any forecast horizon. All that occurs is that the world real interest rate adjusts. From this inference I can construct the first hypothesis with respect to the response of the current account $CA_t$ to a global permanent shock $\epsilon_{gp}^p$:

$$H_0: \frac{\partial CA_t}{\partial \epsilon_{gp}^p} = 0 \text{ for any } i \geq 0.$$ (Hypothesis 1)

Next consider the responses of the current account ratio to permanent and transitory country-specific shocks $\epsilon_{cp}^p$ and $\epsilon_{cs}^s$. Let $\Gamma_{cp}^{ca}(0)$ and $\Gamma_{cs}^{ca}(0)$ denote the impact responses of $CA_t/NO_t$ to $\epsilon_{cp}^p$ and $\epsilon_{cs}^s$, respectively. As shown in Appendix C in detail, eqs. (10) and (13) imply that $\Gamma_{cp}^{ca}(0)$ and $\Gamma_{cs}^{ca}(0)$ should satisfy the following cross-equation restrictions:

$$\Gamma_{cp}^{ca}(0) = \Gamma_{cp}^{no}(0) - \Gamma_{cp}^{no}(\kappa). \quad (R_{cp})$$

and

$$\Gamma_{cs}^{ca}(0) = \Gamma_{cs}^{no}(0) - \Gamma_{cs}^{no}(\kappa). \quad (R_{cs})$$

On the other hand, the response of the current account ratio to a global shock is ambiguous. For example, if a global shock has a positive impact on $\ln NO_t$ and the mean value of $CA_t/NO_t$ is positive, then the current account ratio should respond negatively to the shock.
where for an index \( i \in \{cp, cs\} \) \( \Gamma^{io}(\kappa) \) is \( \Gamma^{io} (L) \) evaluated at \( L = \kappa \).  

The cross-equation restrictions \( R_{cp} \) and \( R_{cs} \) state that the response of the current account-net output ratio to a country-specific shock should be given as the difference between the impact and the discounted long-run responses of the first difference of log net output to the shock. The current account identity (9) restricts the current account-net output ratio to be negatively related to the consumption-net output ratio. Therefore, if a country-specific shock raises net output above (below) consumption, the current account ratio rises (falls). \( \Gamma^{nc}_{cp}(0) \) in \( R_{cp} \) captures the impact effect of shock \( \epsilon^{cp}_t \) on net output. On the other hand, \( \Gamma^{nc}_{cp}(\kappa) \) shows the impact effect of the shock on consumption. Hence the impact effect of the shock on the current account ratio, \( \Gamma^{ca}_{cp}(0) \), is given as the difference \( \Gamma^{nc}_{cp}(0) - \Gamma^{nc}_{cp}(\kappa) \). The same explanation is applicable for \( R_{cs} \).

Define the statistics \( H_{cp} \) and \( H_{cs} \) as
\[
H_{cp} = \Gamma^{ca}_{wp}(L)\epsilon^{gp}_t + \Gamma^{ca}_{cp}(L)\epsilon^{cp}_t + \Gamma^{ca}_{cs}(L)\epsilon^{cs}_t.
\]
Eqs. (13), (14) and (15) imply that the vector \( X_t \) has the following structural VMA (SVMA) representation:
\[
\begin{bmatrix}
\tilde{r}_t \\
\Delta \ln NO_{it} \\
\tilde{C}_A_t/\tilde{NO}_{it}
\end{bmatrix} =
\begin{bmatrix}
\Gamma^r(L) & 0 & 0 \\
\Gamma^{no}_{wp} (L) & \Gamma^{no}_{cp} (L) & \Gamma^{no}_{cs} (L) \\
\Gamma^{ca}_{wp} (L) & \Gamma^{ca}_{cp} (L) & \Gamma^{ca}_{cs} (L)
\end{bmatrix}
\begin{bmatrix}
\epsilon^{gp}_t \\
\epsilon^{cp}_t \\
\epsilon^{cs}_t
\end{bmatrix},
\]
(16)

The cross-equation restrictions \( R_{cp} \) and \( R_{cs} \), I exploit the Wiener-Kolmogorov formula, which is also known as the Hansen and Sargent’s(1980) distributed predicted leads formula.

The underlying fact that consumption is determined by the level of permanent net output makes the impact response of consumption be given as the discounted long-run response of the first difference of log net output. See for example Quah(1990).
or simply
\[ X_t = \Gamma(L)\epsilon_t \] (17)
where \( \epsilon_t \) is the structural shock vector given as \( \epsilon_t = [\epsilon_t^{sp} \; \epsilon_t^{cp} \; \epsilon_t^{cs}]' \). In particular, following the standard exercise in the SVAR literature, I impose that the variance-covariance matrix of the structural shock vector is given as an identity matrix: \( \Sigma \epsilon_t \epsilon_t' = I^{23} \). Notice that the infinite-order lag polynomial \( \Gamma(L) \) has the impact and long-run matrices, \( \Gamma(0) \) and \( \Gamma(1) \), which satisfy
\[
\Gamma(0) = \begin{bmatrix}
\Gamma_r(0) & 0 & 0 \\
\Gamma_{no}^{gp}(0) & \Gamma_{no}^{cp}(0) & \Gamma_{no}^{cs}(0) \\
\Gamma_{ca}^{gp}(0) & \Gamma_{ca}^{cp}(0) & \Gamma_{ca}^{cs}(0)
\end{bmatrix},
\]
(18)
and
\[
\Gamma(1) = \begin{bmatrix}
\Gamma_r(1) & 0 & 0 \\
\Gamma_{no}^{gp}(1) & \Gamma_{no}^{cp}(1) & 0 \\
\Gamma_{ca}^{gp}(1) & \Gamma_{ca}^{cp}(1) & \Gamma_{ca}^{cs}(1)
\end{bmatrix}.
\]
(19)
The lower triangular long-run matrix comes from the small open assumption and the long-run restriction (12), while two exclusion restrictions on the impact matrix are based on the small open assumption.

Testing the maintained hypotheses depends on derivation of the IRFs of the endogenous variables to the structural shocks. To do so, it is necessary to identify the impact matrix \( \Gamma(0) \). This paper follows two standard methods to identify the impact matrix in the SVAR literature. The first is Blanchard and Quah’s(1989) long-run restriction. The second is Gali’s(1992) method that uses both the impact and long-run restrictions in concert. Using two different identification schemes allows me to check whether the empirical results are robust to alternative identifications.

To explain the identification strategy of this paper, notice that since the vector \( X_t \) is stationary, \( X_t \) has the following reduced form VMA representation:
\[ X_t = C(L)\upsilon_t \] (20)
where \( C(L) \) is an infinite order matrix polynomial with respect to the lag operator \( L \) and in particular the coefficient matrix of \( L^0 \) is an identity matrix. \( \upsilon_t \) is the reduced-form disturbance vector with a symmetric positive definite variance-covariance matrix \( \Sigma \). Comparing eq.(17) with eq.(20) immediately provides the following relationships:
\[ \Sigma = \Gamma(0)\Gamma(0)' \] (21)
and
\[ C(L)\Gamma(0) = \Gamma(L). \] (22)
\footnote{That is, the structural shocks are orthogonal at all leads and lags and each shock has a unit variance. Therefore in this paper the impulse response function of a variable is interpreted as the response to a unit standard error shock.}

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Furthermore eq.(22) can rewrite eq.(21) as follows:

$$\Sigma = C(1)^{-1}\Gamma(1)\Gamma(1)'C(1)'^{-1}. \quad (23)$$

Given the estimates of $\Sigma$ and $C(1)$, there are six linear independent equations and nine unknowns of $\Gamma(1)$ in eq.(23). Therefore it needs three additional restrictions to become just identified. Notice that eqs.(16), (18) and (19) imply that the assumptions of the model provide an infinite number of restrictions on the structural parameters in (16): two impact restrictions, three long-run restrictions, and an infinite number of impulse response restrictions. Since three restrictions are needed to just-identify the structural parameters, eq.(16) is an overidentified system. By following the methodology developed by King and Watson(1997) and Nason and Rogers(2001), this paper chooses two sets consisting of three restrictions from the whole overidentifying restrictions in order to just-identify the system.

The long-run matrix (19) implies that the maintained assumptions in this paper provide three long-run restrictions and makes the long-run matrix lower-triangular. Therefore, as exercised by Blanchard and Quah, the long-run matrix (19) is just-identified and the impact matrix can be recovered by eq.(22). Hereafter I call this identification method identification scheme I.

Another method of this paper to identify the impact matrix $\Gamma(0)$ is to exploit together two impact restrictions shown in eq.(18) and the long-run restriction provided by a zero restriction on the (2,3)th element in (19). For exposition, let $A_{i,j}$ denote the $(i,j)$th element in a matrix $A$. Notice that the zero restriction on the (2,3)th element in (19) together with the zero restriction on the (1,3)th element in (18) implies a cross-element restriction on eq.(22) such as

$$C(1)_{2,2}\Gamma(0)_{2,3} + C(1)_{2,3}\Gamma(0)_{3,3} = 0. \quad (24)$$

Since the estimates of $C(1)_{2,2}$ and $C(1)_{2,3}$ are obtained through the reduced-form VAR (RF-VAR) estimation, (24) can be considered as the third impact restriction. The obtained set of three impact restrictions makes it possible to just-identify the impact matrix $\Gamma(0)$ in eq.(23). Hence, the second identification of this paper follows Gali’s method that exploits the impact and long-run restrictions in concert. Again hereafter I call this identification method identification scheme II.

Tables 1 and 2 summarize the two identification schemes and the hypotheses of this paper. In Appendix D, I discuss that all restrictions provided by the identification schemes and the hypotheses can be rewritten as linear restrictions on the impact matrix $\Gamma(0)$. By exploiting this fact, I estimate the impact matrix by the full information ML (the F.I.M.L.) procedure and test each hypothesis based on the asymptotic Wald test.

Moreover the fact that the long-run matrix (19) is lower triangular makes recovering the long-run matrix very simple. $\Gamma(1)$ is simply obtained as the Choleski factor of the matrix $C(1)\Sigma C(1)'^{-1}$. The reason for choosing this long-run restriction from the others is that the restriction is essential for
<table>
<thead>
<tr>
<th>Identification Scheme I</th>
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<tbody>
<tr>
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<td>( \Gamma_{1,2}(1) = 0 )</td>
</tr>
<tr>
<td>Small Open Assumption</td>
<td>( \Gamma_{1,3}(1) = 0 )</td>
</tr>
<tr>
<td>Long-Run Restriction</td>
<td>( \Gamma_{2,3}(1) = 0 )</td>
</tr>
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</table>

Table 1: Identification Schemes

| Hypothesis 1 | \( \frac{\partial CA_t}{\partial \epsilon_{it}} = 0 \) for all \( i \geq 0 \) |
| Hypothesis 2 | \( H_{cp} = \Gamma_{3,2}(0) - \Gamma_{2,2}(0) + \Gamma_{2,2}(\kappa) = 0 \) |
| Hypothesis 3 | \( H_{cs} = \Gamma_{3,3}(0) - \Gamma_{2,3}(0) + \Gamma_{2,3}(\kappa) = 0 \) |

Table 2: Hypotheses

5 Data, SVAR Estimation and Testing Method

This paper studies four G-7 economies, Canada, Japan, the U.K. and the U.S. All data used in this paper are quarterly, span the period Q1:1960-Q4:1997, and are seasonally adjusted at annual rates. The estimation is based on the Q2:1963-Q4:1997 sample, with data prior to Q2:1963 used in data construction. As the world real interest rate series I use a weighted average of \textit{ex ante} real interest rates across G-7 countries, by following Barro and Sala-i-Martin(1990) and Bergin and Sheffrin(2000). The net output and current account series are constructed by using each country’s national accounting data. Appendix E provides detailed information on the source and construction of the data.

The SVAR estimation begins with checking stationarity of the vector \( X_t \). To do so I perform unit root tests for the elements of \( X_t \), \( \tilde{r}_t \), \( \Delta \ln \tilde{NO}_t \) and \( \tilde{CA}_t/\tilde{NO}_t \), based on the augmented Dickey-Fuller (ADF) test. Appendix F summarizes the method and the results of the unit roots tests. Except for \( \tilde{CA}_t/\tilde{NO}_t \) of the U.S., the ADF tests reject the unit root null in all series at least at the 5 \% significance level. Even in the case of \( \tilde{CA}_t/\tilde{NO}_t \) of the U.S., the ADF tests can reject the unit root null at the 10 \% significance level. From this evidence, I consider the series \( \tilde{r}_t \), \( \Delta \ln \tilde{NO}_t \) and \( \tilde{CA}_t/\tilde{NO}_t \) to be stationary fluctuating around zero in the following analysis. In this case, the vector \( X_t \) follows a stationary process.

The next task is to estimate the VMA (20). Since the VMA is invertible, it has an infinite-order VAR representation. Instead of estimating an infinite order VAR, I estimate a truncated VAR with finite order lags as an approximation. To select an optimal lag length, I calculate both the AIC and BIC criteria with maximum lag length fifteen. For Canada, the U.S. and decomposing country-specific shocks into the permanent and transitory components.

As discussed in footnote 15, because of the lower triangular long-run matrix numerical maximization procedure is not needed to recover the impact matrix in identification scheme I. In identification scheme II, the impact matrix is numerically recovered through the F.I.M.L. procedure. See Amisano and Giannini(1997) for the F.I.M.L. estimations of the SVAR models.
the U.K. both criteria pick a lag length of one. A lag length of three is selected for Japan. Estimates of the RFVAR by OLS provide the estimates $\hat{\Sigma}$ and $C(1)$. This allows me to identify the impact matrix $\Gamma(0)$ with each of the identification schemes.

The impact matrix $\Gamma(0)$ is estimated by the F.I.M.L procedure. Since all the restrictions implied by the hypotheses are given as linear restrictions on $\Gamma(0)$ as shown in Appendix D, it is a simple task to test the hypotheses by the asymptotic Wald test. A difficulty is in that Hypothesis 1 provides an infinite number of restrictions on the SVMA system. Hence in this paper, I test Hypothesis 1 up to period 4. Let $W_1$, $W_2$ and $W_3$ denote the Wald statistics for Hypothesis 1 at impact, Hypothesis 2 and Hypothesis 3, respectively. In addition, let $W_4$, $W_5$ and $W_6$ denote the Wald statistics for the joint nulls. $W_4$ is the Wald statistic for the joint null of Hypothesis 2 and Hypothesis 3. $W_5$ is the Wald statistic for the joint null of Hypothesis 1, Hypothesis 2 and Hypothesis 3. Finally, $W_6$ is the Wald statistic for the null that Hypothesis 1 is satisfied up to period 4. Appendix G provides the detailed description of the Wald test in this paper.

As in the standard exercise of the SVAR literature, I estimate the IRFs and the FEVDs of the endogenous variables on the structural shocks. The empirical standard errors of the IRFs and the FEVDs are calculated by generating 10000 nonparametric bootstrapping replications based on the reduced form disturbances. At the same time, I obtain 10000 replications of the statistics $H_{cp}$ and $H_{cs}$ from the bootstrapping exercise. These replications provide the empirical joint distribution of $H_{cp}$ and $H_{cs}$.

6 The Empirical Results

This section reports the empirical results of this paper. This paper emphasizes the robustness of the empirical observations for identifications. Therefore, I concentrate on reporting the identification-robust results.

6.1 Impulse Response Analysis

Figures 1(a) and (b) show the IRFs of the current account across countries under identification schemes I and II, respectively. In each figure, the dark line represents the point estimate and the dashed lines exhibit 95% confidence intervals constructed by the bootstrapping exercise. The identification-robust observations are summarized as follows:

- The IRFs of the current account to a global permanent shock are not significant across all 40 periods after impact.\(^{27}\)
- The impact responses of the current account to a country-specific permanent shock are positive but insignificant in Canada and the U.K. while those of Japan and the U.S. are negative.\(^{28}\)

\(^{27}\)In the U.S. the IRF becomes boundary-significant at a year after impact under identification scheme II, while under identification I no IRFs are significant.

\(^{28}\)In Japan, the response becomes significant at a year after impact though it returns insignificant at 3 years.
• The impact responses of the current account to a country-specific transitory shock are positive and significant across all the countries. The positive responses remain significant for at least 3 years in Canada, Japan and the U.K. and up to 6 years in the U.S.

These results support the basic predictions of the PVM: no response of the current account to a global shock, negative or no response to a country-specific permanent shock and a positive response to a country-specific transitory shock. Figures 2(a) and (b) show the IRFs of log of net output across the countries under identification schemes I and II, respectively. Notice that the responses of log of net output to a country-specific permanent shock trend up in Japan and the U.S. This observation is consistent with the PVM’s prediction that if a country-specific shock raises net output permanently, the current account negatively responds to the shock.

Therefore, the impulse response analysis of this paper qualitatively supports the basic predictions of the PVM. However it does not necessarily mean that the quantitative requirements of the PVM, i.e. cross-equation restrictions, are supported at the same time. Hypotheses 1-3 provide information about the validity of the cross-equation restrictions.

6.2 Testing the Cross-Equation Restrictions

Tables 3(a) and (b) report the results of the asymptotic Wald tests under identification schemes I and II, respectively. Each table shows the Wald statistics and the corresponding p-values for the null hypotheses. The identification-robust rejections are summarized as follows:

- The Wald statistic $W_1$ shows that Hypothesis 1 at impact is rejected in the U.S.
- The Wald statistic $W_4$ shows that Hypotheses 2 and 3 are jointly rejected at least at 12.3\% significance level in Canada and at 0.4\% significance level in the U.K.
- The Wald statistics $W_5$ shows that Hypotheses 1, 2 and 3 are jointly rejected in Canada, the U.K. and the U.S.
- The Wald statistics $W_6$ shows that Hypothesis 1 up to a year is rejected in Canada and the U.S.

These results exhibit that Hypotheses 1-3 are jointly rejected in Canada, the U.K. and the U.S. by the asymptotic Wald test. The strong rejections of the joint hypothesis in the small open countries are consistent with the standard PVT’s results [Sheffrin and Woo(1990), Otto(1992) and Ghosh(1995)]. In particular, notice that in Canada and the U.K. the p-values for $W_3$ take relatively small values across two identifications. This fact suggests a possibility that the rejections in the prototype small open countries are mainly caused by the failures of the cross-equation restriction $R_{cs}$, while the rejection in the U.S. is due to the failure of the neutrality hypothesis.

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29Here the critical level of the p-value for rejection is chosen at 15 percent. I call a rejection identification-robust when the null is rejected across two identifications.

30This statement does not necessarily imply acceptance of the hypotheses in Japan. In fact, Hypotheses 1, 2 and 3 are jointly rejected under identification scheme II even in Japan.

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Figures 3(a) and (b) provide the scatter plots of 10000 replicated pairs of the statistics $H_{cp}$ and $H_{cs}$ by the bootstrapping exercise. In each window, the dark square represents the point estimate and the null is represented by the origin. Observe that in Canada and the U.K. the scatter plots have strikingly similar shapes and almost all replicated pairs are concentrated on the upper regions of the windows regardless of identification. Therefore small-sample uncertainty of the point estimates also provides strong evidence that the rejections in the proto-type small open countries are mainly caused by the failures of the cross-equation restriction $R_{cs}$.

Furthermore, by construction of the statistic $H_{cs}$, the upper-concentrated scatter plots reveal a surprising empirical fact: *in the proto-type small open countries the impact responses of the current account-net output ratio to a country-specific transitory shock are too large to support the PVM.* Of course, this finding is related to Glick and Rogoff’s puzzle: the impact response of the current account to a country-specific permanent productivity shock is too *small.* As a resolution they consider a significant mean reverting component in productivity. However as argued by Hoffmann(1999), Glick and Rogoff’s estimation fails to avoid Quah’s critique: if the econometrician does not observe the permanent and transitory components, he will incorrectly measure the response of the current account to country-specific shocks 31. By contrast, this paper has clear identification schemes based on Blanchard and Quah’s decomposition. It presents a new puzzle: the impact response of the current account measure is too *large* rather than too small. No existing literature points out the *excess* response of the current account measure to a country-specific transitory shock. This empirical finding is a novel contribution of this paper.

### 6.3 Forecast Error Variance Decomposition Analysis

Tables 4(a) and (b) provide the FEVDs of the current account attributed to three structural shocks across the countries under *identification schemes I and II*, respectively. At impact, regardless of identification, a country-specific transitory shock can significantly explain at least 65% of fluctuations in the current account across all the countries. Even at a year after impact, the shock can significantly explain at least 60% of fluctuations in the current account in Canada, the U.K. and the U.S. Therefore the country-specific transitory shock can be considered as the dominant driving force of the current account in the short run.

A striking fact revealed by tables 4(a) and (b) is that *even in the long-run the country-specific transitory shock can significantly and dominantly explain fluctuations in the current account of the small open countries.* For example, at 40 quarters after impact(i.e.10 years after impact), about 80% of fluctuations in the Canadian current account is attributed to the country-specific transitory shock. Similarly, at the same forecast horizon, the shock explains 72% of fluctuations in the U.K. current account. On the other hand, the contribution of the country-specific transitory shock in the large countries becomes small and insignificant.

31 More recently Işcan(2000) introduces nontradable goods into Glick and Rogoff’s analytical framework. His extension makes it possible to decompose country-specific permanent productivity shocks into components originated from tradable and nontradable goods sectors. Although his decomposition is insightful, he still faces on Quah’s critique due to no identification of permanent and transitory shocks.
in the long-run. Especially under *identification schemes I*, at 40 quarters after impact, the country-specific permanent shock turns out to be the main driving force of the Japanese current account, while the global permanent shock explains a dominant part of the U.S. current account fluctuations.

The strong long-run dependence of the current account on the country-specific transitory shock in the small open countries is in fact a puzzling observation of this paper: *in spite of the fact that the effect of a country-specific transitory shock on net output dies out quickly, why can the shock have such a strong influence on small open country’s current account movements in the long-run?* This result is consistent with the finding of Nason and Rogers (2001). In their SVAR approach to study the joint-dynamics of investment and the current account, they reports the persistent dependence of the current account on country-specific shocks across the G-7 countries. As they argue, at present there is no consensus intertemporal model that generates persistence in the current account to country-specific transitory shocks. Together with the excess response of the current account measure to a country-specific transitory shock, this puzzling result provides new goal of the current account research.

### 6.4 Is Identification Successful?

The role of this subsection is to check whether the identification of this paper is successful or not. If the identification is relevant, regardless of the chosen identification scheme, the IRFs of the world real interest rate to a global permanent shock should be estimated *identically* across the countries. Moreover by definition the identified global shock should be perfectly correlated across the countries, while the identified country-specifics permanent and transitory shocks should be uncorrelated.

Figures 4(a) and (b) show the IRFs of $\tilde{r}_t$ to $\epsilon_{gp}^t$ under *identification schemes I and II*, respectively. Regardless of the chosen identification scheme, the estimated IRFs are strikingly similar across the countries with respect to the size and shape. In particular, at impact the world real interest rate rises by about 0.008 to one standard error global shock and the effect of the shock on the world real interest rate becomes insignificant around 4 years after impact. This observation is commonly found in all the countries under two identification schemes.

On the other hand, tables 5(a) and (b) report the cross-correlation matrices of the identified structural shocks across the countries under *identification schemes I and II*, respectively. The tables clearly show that the identified global permanent shocks are highly correlated across the countries while the cross-country correlations of the identified country-specific shocks are negligibly small in almost all the cases. Especially, the minimum correlation of global permanent shocks is 0.7548 between the U.K. and the U.S. under *identification schemes I*. The cross-country correlations of the identified country-specific shocks are less than 0.2 (in absolute

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32This is simply because $\epsilon_{gp}^t$ should be estimated as a global shock. Since the same world real interest rate measure is used in all the countries’ estimations, the response of the world real interest rate to $\epsilon_{gp}^t$ should be estimated identically across all countries if $\epsilon_{gp}^t$ is identified correctly.
value) in almost all the cases with four exceptions\textsuperscript{33}. 

The observations in this subsection, therefore, provide an evidence that the identification of this paper is fairly successful.

6.5 Is the Homogeneity Assumption Relevant?

The estimation of this paper shows that Hypothesis 1 is strictly rejected in the U.S. That is, the identified global permanent shock has a significant effect on the current account in this country. A caveat against this result is that Hypothesis 1 depends on the homogeneity assumption across countries. Since under this assumption all the countries react symmetrically to a global shock, there arise no opportunity to trade across countries. If this assumption does not hold, even a global shock can generate the current account flows across countries that alter their net foreign asset positions. As a result, the current account should adjust to a global shock and Hypothesis 1 becomes irrelevant in this case.

To check the relevance of the homogeneity assumption, it helps to examine the response of net output to a global shock. If the countries are homogeneous with respect to technology, preference and endowment, it is a reasonable inference that the response of net output to a global shock should be identical across the countries.

Observe again figures 2(a) and (b) that show the IRFs of log of net output to a global permanent shock. These figures show that there are variations in the point estimate of the IRFs across the countries, although the IRFs are insignificant in almost all the cases. In particular, Canada shares the similar shape with Japan, while the U.K. shares the similar shape with the U.S. However two country pairs does not have the similar shape of the IRFs.

Therefore, the result that there exist cross-country variations in the estimated IRFs of net output to a global permanent shock casts doubt on the relevance of the homogeneity assumption and in turn Hypothesis 1 itself. Glick and Rogoff implicitly assume the homogeneity across the G-7 countries to construct the neutrality hypothesis of the current account to a global shock. The observation of this paper implies the possibility that even across the G-7 countries the heterogeneity in economic structure is important enough that the neutrality hypothesis becomes irrelevant.

7 Conclusion

This paper tests the PVM’s predictions on the responses of the current account measure to structural shocks and approaches the question: what shocks are main driving forces of the current account? Instead of following the present value test framework, this paper introduces

\textsuperscript{33}The exceptions are the correlation of country-specific permanent shocks between Canada and the U.S. under identification schemes I and II (0.3020 and 0.3104, respectively), the correlation of country-specific transitory shocks between Japan and the U.K. under identification schemes I and II (0.4356 and 0.2409, respectively) and that between Canada and Japan under identification schemes I (-0.2198).
a structural VAR approach for testing the model. The PVM predicts different responses of the current account measure to different structural shocks. By exploiting two kinds of identification schemes based on the small open country assumption and Blanchard and Quah’s (1989) long-run restriction, this paper identifies global permanent, country-specific permanent and country-specific transitory shocks and test the predicted responses of the current account measure to these shocks.

In summary, the identification-robust results from the quarterly data of Canada, Japan, the U.K. and the U.S. are observed as follows: (i) the predictions of the PVM are qualitatively supported by the impulse response analysis, (ii) the predictions of the PVM are, however, quantitatively rejected in Canada, the U.K. and the U.S.: the cross-equation restrictions are jointly rejected in these countries, (iii) the estimated responses of the current account measure to a country-specific transitory shock in the small open countries Canada and the U.K. are too large to support the PVM and finally (iv) in the small open countries the variations in the current account are dominated by the country-specific transitory shock in not only the short-run but also the long-run.

Therefore this paper reveals a never-mentioned puzzling fact: current account movements in the small open countries depend excessively on country-specific transitory shocks. It is a future goal of the current account research to provide a clear resolution for this puzzle.
Appendices

Appendix A: Derivation of the Linear-Approximated Intertemporal Budget Constraint (7)

Dividing the intertemporal budget constraint (6) by \( NO_t \) gives

\[
\frac{C_t}{NO_t} \sum_{i=0}^{\infty} E_t R_{t,i} \frac{C_{t+i}}{C_t} = (1 + r_t) \frac{B_t}{NO_t} + \sum_{i=0}^{\infty} E_t R_{t,i} \frac{NO_{t+i}}{NO_t}.
\]

(A1)

Since for any variable \( X_t \) the relation \( X_{t+i} \approx X_t X_{t+i} \ldots X_{t+i} X_{t+i+1} \) holds, (A1) can be rewritten as

\[
\frac{C_t}{NO_t} \left[ 1 + \sum_{i=1}^{\infty} E_t R_{t,i} \prod_{j=t+1}^{t+i} \left( \frac{C_j}{C_{j-1}} \right) \right] = (1 + r_t) \frac{B_t}{NO_t} + \left[ 1 + \sum_{i=1}^{\infty} E_t R_{t,i} \prod_{j=t+1}^{t+i} \left( \frac{NO_j}{NO_{j-1}} \right) \right].
\]

(A2)

Notice that for any variable \( X_t \), the relation \( \prod_{j=t+1}^{t+i} X_j = \exp\{\sum_{j=t+1}^{t+i} \ln(X_j)\} \) holds. From this relation and the definition of \( R_{t,i} \), (A2) can be further rearranged as

\[
\frac{C_t}{NO_t} \left[ 1 + \sum_{i=1}^{\infty} E_t \exp \left\{ \sum_{j=t+1}^{t+i} (\Delta \ln C_j - \ln(1 + r_j)) \right\} \right] = \exp\{\ln(1 + r_t) - \Delta \ln NO_t\} \frac{B_t}{NO_{t-1}}
\]

\[
+ \left[ 1 + \sum_{i=1}^{\infty} E_t \exp \left\{ \sum_{j=t+1}^{t+i} (\Delta \ln NO_j - \ln(1 + r_j)) \right\} \right].
\]

(A3)

Taking a first-order Taylor expansion of the LHS of (A3) around the mean values gives

\[
\text{The LHS} \approx \frac{1}{1 - \alpha NO_t} \frac{\widetilde{C_t}}{NO_t} + \frac{\alpha}{1 - \alpha} \sum_{i=1}^{\infty} \alpha^i E_t \left\{ \Delta \ln \widetilde{C_{t+i}} - \ln(1 + \widetilde{r_{t+i}}) \right\}.
\]

(A4)

where \( \alpha = \exp(\gamma - \mu) < 1 \). The RHS of (A3) is also approximated as

\[
\text{The RHS} \approx \frac{1}{\kappa NO_{t-1}} \frac{\widetilde{B_t}}{NO_t} + \frac{\kappa}{\kappa} \ln(1 + r_t) - \frac{\kappa \Delta \ln NO_t}{1 - \kappa} \sum_{i=1}^{\infty} \kappa^i E_t \left\{ \Delta \ln \widetilde{NO_{t+i}} - \ln(1 + \widetilde{r_{t+i}}) \right\}.
\]

(A5)

where \( \kappa = \exp(\gamma - \mu) < 1 \). From the results (A4) and (A5), the linear-approximated intertemporal budget constraint (7) is finally given as

\[
\frac{\widetilde{C_t}}{NO_t} \approx \frac{1 - \alpha}{\kappa} \frac{\widetilde{B_t}}{NO_{t-1}} + \frac{1 - \alpha}{\kappa} \ln(1 + r_t) - \frac{1 - \alpha}{\kappa} \Delta \ln \widetilde{NO_t}
\]

\[
- \frac{\alpha}{1 - \kappa} \sum_{i=1}^{\infty} \alpha^i E_t \left\{ \Delta \ln \widetilde{C_{t+i}} - \ln(1 + \widetilde{r_{t+i}}) \right\}
\]

\[
+ \frac{1 - \alpha}{1 - \kappa} \sum_{i=1}^{\infty} \kappa^i E_t \left\{ \Delta \ln \widetilde{NO_{t+i}} - \ln(1 + \widetilde{r_{t+i}}) \right\}.
\]
Appendix B: Derivation of the Approximated Solution of the Optimal Current Account Ratio

Substitute the log-linearized Euler equation (8) into the linear-approximated intertemporal budget constraint (6). For simplicity, assuming that the economy is around the balanced growth path; $\alpha = \kappa$ and using the approximation $\ln(1 + r_t) \approx r_t$, I can obtain the consumption ratio equation as

$$\frac{\widetilde{C}_t}{\widetilde{NO}_t} = 1 - \frac{1 - \kappa}{\kappa} \frac{\widetilde{B}_t}{\widetilde{NO}_{t-1}} + \frac{1 - \kappa}{\kappa} b \widetilde{r}_t - \frac{1 - \kappa}{\kappa} b \Delta \ln \widetilde{NO}_t$$

$$- (\sigma - 1)c \sum_{i=1}^{\infty} \kappa^i E_t \widetilde{r}_{t+i} - \sum_{i=1}^{\infty} \kappa^i E_t \Delta \ln \widetilde{NO}_{t+i} - \sum_{i=1}^{\infty} \kappa^i E_t \Delta \ln \widetilde{NO}_{t+i}. \quad (B1)$$

To derive the optimal current account ratio equation, consider the current account identity (9). Dividing (9) by $NO_t$ rewrites (9) as

$$\frac{CA_t}{NO_t} = 1 + \frac{\exp[\ln(1 + r_t)] - 1}{\exp(\Delta \ln NO_t)} \frac{B_t}{NO_{t-1}} - \frac{C_t}{NO_t}.$$  
Taking a first-order Taylor expansion of the above equation gives

$$\widetilde{CA}_t \approx \frac{1}{\kappa} - \frac{1}{\exp(\gamma)} \frac{\widetilde{B}_t}{\widetilde{NO}_{t-1}} + \frac{b}{\kappa} \widetilde{r}_t - \frac{1}{\kappa} - \frac{1}{\exp(\gamma)} \Delta \ln \widetilde{NO}_t - \frac{\widetilde{C}_t}{\widetilde{NO}_t}. \quad (B2)$$

Substituting the consumption equation (B1) into (B2), I can obtain the equation of optimal current account ratio;

$$\frac{\widetilde{CA}_t}{\widetilde{NO}_t} = \left[1 - \frac{1}{\exp(\gamma)}\right] \frac{\widetilde{B}_t}{\widetilde{NO}_{t-1}} + b \widetilde{r}_t - \left[1 - \frac{1}{\exp(\gamma)}\right] \Delta \ln \widetilde{NO}_t$$

$$+ [(\sigma - 1)c + 1] \sum_{i=1}^{\infty} \kappa^i E_t \widetilde{r}_{t+i} - \sum_{i=1}^{\infty} \kappa^i E_t \Delta \ln \widetilde{NO}_{t+i}.$$  

Since $\exp(\gamma)$ takes a close value to one, it might be a reasonable approximation to set the coefficient $[1 - 1/\exp(\gamma)]$ to zero. Then the optimal current account equation (10) is constructed as follows;

$$\frac{\widetilde{CA}_t}{\widetilde{NO}_t} = b \widetilde{r}_t + [(\sigma - 1)c + 1] \sum_{i=1}^{\infty} \kappa^i E_t \widetilde{r}_{t+i} - \sum_{i=1}^{\infty} \kappa^i E_t \Delta \ln \widetilde{NO}_{t+i}.$$  

Appendix C: Derivation of Cross-Equation Restrictions $H_{cp}$ and $H_{cs}$

To derive the cross-equation restrictions $H_{cp}$ and $H_{cs}$, I exploit the Wiener-Kolmogorov formula, which is well-known as Hansen and Sargent’s(1980) distributed predicted leads formula. For exposition, I give this formula as the following lemma without proof;

**Lemma (Hansen and Sargent(1980)).** For a covariance-stationary process $X_t$ with a Wold MA representation $X_t = A(L)\nu_t$ and $\beta \in (0, 1)$, it is the case that

$$\sum_{i=1}^{\infty} \beta^i E_t X_{t+i} = \beta \left[ A(L) - A(\beta) \right] \frac{L}{L - \beta} \nu_t.$$
By using the present value relation (10), the maintained DGPs of the first log difference of net output and the world real interest rate, (13) and (14), and the above lemma, I derive a structural MA representation of the current account ratio such as

$$\frac{CA_t}{NO_t} = \Gamma_{ca}^{wp}(L)e_t^{wp} + \Gamma_{cp}^{ca}(L)e_t^{cp} + \Gamma_{cs}^{ca}(L)\epsilon_t^{cs}. \quad (C1)$$

where $\Gamma_{ca}^{wp}(L)$, $\Gamma_{cp}^{ca}(L)$ and $\Gamma_{cs}^{ca}(L)$ are infinite-order polynomials, respectively, which satisfy;

$$\Gamma_{ca}^{wp}(L) = b\Gamma^{\nu}(L) + c(\sigma - 1 + 1)\kappa \left[ \frac{\Gamma^{\nu}(L) - \Gamma^{\nu}(\kappa)}{L - \kappa} \right] - \kappa \left[ \frac{\Gamma_{c\nu}^{wp}(L) - \Gamma_{c\nu}^{wp}(\kappa)}{L - \kappa} \right], \quad (C2)$$

$$\Gamma_{cp}^{ca}(L) = -\kappa \left[ \frac{\Gamma_{c\nu}^{wp}(L) - \Gamma_{c\nu}^{wp}(\kappa)}{L - \kappa} \right], \quad (C3)$$

and

$$\Gamma_{cs}^{ca}(L) = -\kappa \left[ \frac{\Gamma_{c\nu}^{cs}(L) - \Gamma_{c\nu}^{cs}(\kappa)}{L - \kappa} \right]. \quad (C4)$$

Since the impact responses of the current account ratio to $\epsilon_t^{cp}$ and $\epsilon_t^{cs}$ are given as $\Gamma_{cp}^{ca}(0)$ and $\Gamma_{cs}^{ca}(0)$, respectively, $H_{cp}$ and $H_{cs}$ are obvious from (C3) and (C4).

**Appendix D: Implied Linear Restrictions on the Impact Matrix**

In this appendix I show that all the restrictions shown in Tables 1 and 2 can be rewritten as linear restrictions on the impact matrix $\Gamma(0)$. For exposition, let $[A]_i^r$ and $[A]_i^c$ denote the $i$ th row and column vectors of a matrix $A$, respectively. First of all, two of three restrictions in Identification Scheme II are zero restrictions on the impact matrix, which mean that they are linear restrictions on the impact matrix in nature. More precisely, let $e_i$ denote a $1 \times 3$ row vector which has zeros as the $j \neq i$ th elements and one as the $i$ th element. Then these exclusion restrictions are rewritten as $\Gamma_{1,2}(0) = e_1[\Gamma(0)]_2^r = 0$ and $\Gamma_{1,3}(0) = e_1[\Gamma(0)]_3^c = 0$, respectively. Secondly, it is the case from eq.(22) that $C(1)[\Gamma(0) = \Gamma(1)$. This relation implies that $\Gamma_{i,j}(1) = [C(1)]_i^r[\Gamma(0)]_j^c$ for any $i, j = 1, 2, 3$. Therefore a long-run restriction $\Gamma_{i,j}(1) = 0$ should be equal to an orthogonality condition between the $i$ th row vector of $C(1)$ and the $j$ th column vector of $\Gamma(0)$ and can be rewritten as a linear restriction on the impact matrix. These results revise Table 1 as follows:

<table>
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<tr>
<th>Identification Scheme I</th>
<th>Identification Scheme II</th>
</tr>
</thead>
<tbody>
<tr>
<td>Small Open Assumption</td>
<td>$[C(1)]_1^r[\Gamma(0)]_2^r = 0$</td>
</tr>
<tr>
<td>Small Open Assumption</td>
<td>$[C(1)]_1^c[\Gamma(0)]_3^c = 0$</td>
</tr>
<tr>
<td>Long-Run Restriction</td>
<td>$[C(1)]_2^r[\Gamma(0)]_3^r = 0$</td>
</tr>
</tbody>
</table>

Table D1

Note that in Table D1 all the restrictions are given as linear restrictions on the impact matrix $\Gamma(0)$.

To find the impulse response functions (the IRFs) of $CA_t$ to the structural shock $\epsilon_{t-i}^{cp}$, for any $i \geq 0$, I take a derivative of the identity $CA_t \equiv (CA_t/NO_t)NO_t$ and obtain the following relation

$$\frac{\partial CA_t}{\partial \epsilon_{t-i}^{cp}} = \frac{CA}{CA/NO} \frac{\partial A_t/NO_t}{\partial \epsilon_{t-i}^{cp}} + \frac{CA}{\partial \ln NO_t} \frac{\partial \ln NO_t}{\partial \epsilon_{t-i}^{cp}}$$

23
where \( CA/NO \) and \( CA \) are means of \( CA_t/NO_t \) and the \( CA_t \) respectively. In particular, the last term in the RHS of the above relation can be given by the accumulated impulse response of \( \Delta \ln NO_t \) to \( \epsilon_{t-1}^{gp} \). Hence the IRF of \( CA_t \) to the structural shock \( \epsilon_{t-1}^{gp} \) is given by

\[
\frac{\partial CA_t}{\partial \epsilon_{t-1}^{gp}} = \frac{CA}{CA/NO} \Gamma_{gp,i}^{ca} + CA \sum_{s=0}^{i} \Gamma_{gp,s}^{no}
\]

where \( \Gamma_{gp,i}^{ca} \) and \( \Gamma_{gp,i}^{no} \) are the impulse responses of \( \Delta \ln NO_t \) and \( CA_t/NO_t \) to \( \epsilon_{t-1}^{gp} \), respectively.

Let \( C_i \) denote the coefficient matrix of \( L^i \) in the VMA (20). Since \( \gamma_i = C_i \Gamma(0) \) for any \( i \geq 0 \), the IRFs, \( \Gamma_{gp,i}^{no} \) and \( \Gamma_{gp,i}^{ca} \), can be written as follows:

\[
\Gamma_{gp,i}^{no} = (C_i \Gamma(0))_{2,1} = [C_i]_{2,1}[\Gamma(0)]_i^n
\]

\[
\Gamma_{gp,i}^{ca} = (C_i \Gamma(0))_{3,1} = [C_i]_{3,1}[\Gamma(0)]_i^n
\]

These equations and eq.(D1) rewrite Hypothesis 1 as

\[
\left\{ [C_i]_1^n + \frac{CA}{NO} \sum_{s=0}^{i} [C_s]_2^n \right\}[\Gamma(0)]_i^n = R_i[\Gamma(0)]_i^n = 0 \quad \forall i \geq 0.
\]

where \( R_i \) is a \( 1 \times 3 \) row vector such that \( R_i = \{(C_i)_{2,1} + (CA/NO) \sum_{s=0}^{i} [C_s]_2^n\} \). Therefore Hypothesis 1 is also rewritten as the linear restriction on the impact matrix.

Next notice from eq.(20) that \( \Gamma(\kappa) = C(\kappa) \Gamma(0) \) and thus \( \Gamma_{i,j}(\kappa) = [C(\kappa)]_{i,j}^n[\Gamma(0)]_j^n \) for any \( i, j = 1, 2, 3 \). By using this fact, I can rewrite Hypothesis 1 for \( i = 2,3 \) as

\[
\Gamma_{3,i}(0) = \Gamma_{2,i}(0) - [C(\kappa)]_{i,j}^n[\Gamma(0)]_j^n
\]

or more compactly, with a \( 1 \times 3 \) row vector \( R = [C_{2,1}(\kappa) \quad C_{2,2}(\kappa) - 1 \quad C_{2,3}(\kappa) + 1] \),

\[
R[\Gamma(0)]_i^n = 0.
\]

Eq.(D3) shows that Hypothesis 2 and 3 are also given as the linear restrictions on the impact matrix. From eqs.(D2) and (D3), I can revise Table 2 as Table D2:

| Hypothesis I | \( R_i[\Gamma(0)]_i^n = 0 \quad \forall i \geq 0 \) |
| Hypothesis 2 | \( R[\Gamma(0)]_i^n = 0 \) |
| Hypothesis 3 | \( R[\Gamma(0)]_i^n = 0 \) |

Table D2

where \( R_i = \{(C_i)_{2,1} + (CA/NO) \sum_{s=0}^{i} [C_s]_2^n\} \) and \( R = [C_{2,1}(\kappa) \quad C_{2,2}(\kappa) - 1 \quad C_{2,3}(\kappa) + 1] \).

A striking fact revealed by Tables D1 and D2 is that under the null hypothesis the impact matrix \( \Gamma(0) \) should be singular. To show this, first consider Identification Scheme I. Notice that there are three linear restrictions on \( \Gamma(0)_{3,3}^{no} \) under the null: \( \left[C(1)\right]_{2,1}[\Gamma(0)]_{2,1}^{no} = 0, \left[C(1)\right]_{2,2}[\Gamma(0)]_{2,2}^{no} = 0 \) and \( R[\Gamma(0)]_{3,3}^{no} = 0 \). Since these restrictions are linearly independent and \( \left[\Gamma(0)\right]_{3,3}^{no} \) is a \( 3 \times 1 \) vector, a unique solution for \( \left[\Gamma(0)\right]_{3,3}^{no} \) exists and should be equal to zero. This implies then that the impact matrix \( \left[\Gamma(0)\right]^{no} \) should be singular under the null. The same result is obtained even with Identification Scheme II. In this case, three linearly independent restrictions on \( \left[\Gamma(0)\right]_{3,3}^{no} \) under the null are given as \( e_1 \left[\Gamma(0)\right]_{2,1}^{no} = 0, \left[C(1)\right]_{2,2}[\Gamma(0)]_{2,2}^{no} = 0 \)
and $R[\Gamma(0)]^c_3 = 0$. Therefore a unique solution for $[\Gamma(0)]^c_3$ exists and equals to zero. The impact matrix should be singular under the null.

The singularity of the impact matrix makes it impossible to examine the LR and LM tests for the null since these asymptotic tests depend on the restricted ML estimates of the test statistics. On the other hand, the asymptotic Wald test that exploits only the unrestricted ML estimates is applicable for this situation.

**Appendix E: Data Description and Construction**

This paper uses quarterly data of four G-7 economies, Canada, Japan, the U.K. and the U.S., which spans the sample period Q1:1960-Q4:1997. All data are seasonally adjusted at annual rates and provided by Datastream and IFS.

To construct a measure of the world real interest rate, $r_t$, I follow the method of Barro and Sala-i-Martin (1990) and Bergin and Sheffrin (2000). I collect short-term nominal interest rates, three-month Treasury bill rates or money market rates, on the G-7 economies from IFS. The inflation rate in each country is calculated by using that country’s CPI and the expected inflation rate is constructed by regressing the inflation rate on its own eight lags. The nominal interest rate is then subtracted by the expected inflation rate to compute an ex-ante real interest rate. The world real interest rate is computed by taking a weighted average of ex-ante real interest rates across the G-7 economies, with time-varying weights for each country based on its share of real GDP in the G-7 total.

To construct the net output and current account series, I use each country’s national accounting data distributed by Datastream. All nominal series are converted to real series by using the GDP price deflators. Following definition, I construct the net output series, $NO_t$, by subtracting gross fixed capital formation, change in stocks and government consumption expenditure from GDP. Taking a log of the net output series and a first difference of the resulting logarithmic series provide the first difference of log net output $\Delta \ln NO_t$. The current account series, $CA_t$, is constructed by subtracting gross fixed capital formation, change in stocks, government consumption expenditure and private consumption expenditure from GNP. Dividing $CA_t$ by $NO_t$ provides the series of the current account-net output ratio, $CA_t/NO_t$.

Finally the three series, $r_t$, $\Delta \ln NO_t$ and $CA_t/NO_t$, are demeaned to construct the series, $\tilde{r}_t$, $\Delta \ln \tilde{NO}_t$ and $\tilde{CA}_t/\tilde{NO}_t$.

**Appendix F: Unit Root Tests**

To check the maintained assumption that $\tilde{r}_t$, $\Delta \ln \tilde{NO}_t$ and $\tilde{CA}_t/\tilde{NO}_t$ are stationary, I apply the augmented Dickey-Fuller test (the ADF test) for the three series. The ADF $\tau$-statistics for a time series $y_t$ is given as a t-statistics of the coefficient $\lambda$ in the following OLS regression

$$\Delta y_t = \lambda y_{t-1} + \sum_{i=1}^{n} \Delta y_{t-i} + \eta_t. \quad (F1)$$

where the lag length $n$ is chosen to render $\eta$ white noise. Since the demeaned series $\tilde{r}_t$, $\Delta \ln \tilde{NO}_t$ and $\tilde{CA}_t/\tilde{NO}_t$ fluctuate around zero and have no clear time trend, I do not include either constant or a time trend in the ADF regression (F1). Davidson and MacKinnon (1993) provide asymptotic 10%, 5% and 1% critical values for the Dickey-Fuller $\tau$-statistics equal to -1.62, -1.94 and -2.56, respectively. I perform this test for three choices of the lag length, one, three and five.

Table F1 summarizes the results of the unit root tests. Except for $\tilde{CA}_t/\tilde{NO}_t$ of the U.S., the ADF tests reject the unit root null in all series at least at the 5% significance level for all cases of the lag length. In the case of $\tilde{CA}_t/\tilde{NO}_t$ of the U.S., the ADF tests reject the unit root null at the 10% significance level.
for three and five lags, while the unit root null cannot be rejected even at 10% significance level for the case of one lag.

Appendix G: The Asymptotic Wald Test

Suppose that the RFVAR parameter matrices with an optimal lag \( p \) are estimated as \( \hat{B} = (\hat{B}_1, \hat{B}_2, \ldots, \hat{B}_p) \) with the variance-covariance matrix \( \hat{V} \). Notice that since the SVMA system in this paper is just-identified by one of the identification scheme, the estimate of the impact matrix \( \hat{\Gamma}(0) \) is given as a nonlinear function of \( \hat{B} \). Let \( \mathcal{H}_1 = R_i[\Gamma(0)]_{11} \), \( \mathcal{H}_{cp} = R[\Gamma(0)]_{22} \) and \( \mathcal{H}_{cs} = R[\Gamma(0)]_{33} \), and \( \mathcal{H}_1, \mathcal{H}_{cp} \) and \( \mathcal{H}_{cs} \) denote the corresponding estimates, respectively. The Delta method yields the variance \( \sigma_i \) of \( \mathcal{H}_i \) for \( i = \{1, cp, cs\} \) as \( \sigma_i = \frac{\partial \mathcal{H}_i}{\partial \hat{B}} \hat{V} \frac{\partial \mathcal{H}_i}{\partial \hat{B}}' \). Then the Wald statistic for \( \mathcal{H}_i \) is constructed as

\[
W_i = (\mathcal{H}_i)^2/\sigma_i.
\]

The asymptotic theory says that \( W_i \) asymptotically follows the chi-squared distribution with degree of freedom 1. By using the chi-squared distribution, I can derive the p-value for the null.

To derive the Wald statistic \( W_4 \) for the joint null \( H_0 : \mathcal{H}_{cp} = \mathcal{H}_{cs} = 0 \), construct a row vector \( \lambda = [\hat{\mathcal{H}}_{cp} \hat{\mathcal{H}}_{cs}] \). Let \( \Omega \) denote the variance-covariance matrix of \( \lambda \), which is given as \( \Omega = \frac{\partial \lambda}{\partial \hat{B}} \hat{V} \frac{\partial \lambda}{\partial \hat{B}}' \). Then the Wald statistic for the joint null is constructed as

\[
W_4 = \lambda \Omega^{-1} \lambda'.
\]

According to the asymptotic theory, \( W_i \) asymptotically follows the chi-squared distribution with degree of freedom 2 \(^{34}\). The same argument is applicable for deriving the Wald statistics \( W_5 \) and \( W_6 \).

References


\(^{34}\)For Wald tests of nonlinear restrictions on the reduced form VAR parameters, see Lütkepohl(1991). Gregory and Veall(1985) point out the problem that in finite samples, changing the form of a nonlinear restriction to a form which is algebraically equivalent under the null hypothesis will change the numerical value of the Wald test statistic. Other test statistics, for example the LR and LM statistics, need the restricted estimators, which it is impossible to derive in this paper’s case. Hence although there exists the sample size problem potentially, I apply the Wald statistics to test the maintained hypotheses.


Nason, J. M. and Rogers, J. H. (2000). The present-value model of the current account has been rejected: round up the usual suspects. *manuscript, Department of Economics, University of British Columbia*.


Figure 1(a): Impulse Responses of the CA under Identification Scheme I

Responses to a Global Permanent Shock

Responses to a Country-Specific Permanent Shock

Responses to a Country-Specific Transitory Shock

Note: The dashed lines represent 95% intervals based on 10000 nonparametric bootstrapping resamples.
Figure 1(b): Impulse Responses of the CA under Identification Scheme II

Responses to a Global Permanent Shock

Responses to a Country-Specific Permanent Shock

Responses to a Country-Specific Transitory Shock

Note: The dashed lines represent 95% intervals based on 10000 nonparametric bootstrapping resamples.
Figure 2(a): Impulse Responses of lnNO under Identification Scheme I

Responses to a Global Permanent Shock

Responses to a Country-Specific Permanent Shock

Responses to a Country-Specific Transitory Shock

Note: The dashed lines represent 95% intervals based on 10000 nonparametric bootstrapping resamples.
Figure 2(b): Impulse Responses of lnNO under Identification Scheme II

Responses to a Global Permanent Shock

Responses to a Country-Specific Permanent Shock

Responses to a Country-Specific Transitory Shock

Note: The dashed lines represent 95% intervals based on 10000 nonparametric bootstrapping resamples.
Figure 3(a)  Bootstrapping Scatter Plots of Hcp and Hcs: Identification Scheme I

Note1: The scatter plots are based on 10000 nonparametric bootstrapping resamples of the RFVAR residuals.
Note2: The dark squares are the ML estimates.
Figure 3(b) Bootstrapping Scatter Plots of $H_{cp}$ and $H_{cs}$: Identification Scheme II

Note 1: The scatter plots are based on 10,000 nonparametric bootstrapping resamples of the RFVAR residuals.

Note 2: The dark squares are the ML estimates.
Figure 4(a): Impulse Responses of the World Real Interest Rate to a Global Permanent Shock under Identification Scheme I

Note: The dashed lines represent 95% intervals based on 10000 nonparametric bootstrapping resamples.
Figure 4(b): Impulse Responses of the World Real Interest Rate to a Global Permanent Shock under Identification Scheme II

Note: The dashed lines represent 95% intervals based on 10000 nonparametric bootstrapping resamples.
Table 3: Asymptotic Wald Tests

<table>
<thead>
<tr>
<th></th>
<th>Canada</th>
<th>U.K.</th>
<th>Japan</th>
<th>U.S.</th>
</tr>
</thead>
<tbody>
<tr>
<td>(W_1)</td>
<td>0.190</td>
<td>0.758</td>
<td>0.014</td>
<td>3.445</td>
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<tr>
<td>p-value</td>
<td>0.663</td>
<td>0.384</td>
<td>0.905</td>
<td>0.063</td>
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<td>(W_2)</td>
<td>0.069</td>
<td>0.001</td>
<td>0.003</td>
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<tr>
<td>p-value</td>
<td>0.793</td>
<td>0.983</td>
<td>0.953</td>
<td>0.519</td>
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<tr>
<td>(W_3)</td>
<td>1.562</td>
<td>1.589</td>
<td>0.015</td>
<td>0.000</td>
</tr>
<tr>
<td>p-value</td>
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<td>0.208</td>
<td>0.903</td>
<td>0.999</td>
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<tr>
<td>(W_4)</td>
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<td>p-value</td>
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<td>0.000</td>
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<table>
<thead>
<tr>
<th></th>
<th>Canada</th>
<th>U.K.</th>
<th>Japan</th>
<th>U.S.</th>
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</thead>
<tbody>
<tr>
<td>(W_1)</td>
<td>10.416</td>
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<td>251.775</td>
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<td>p-value</td>
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<td>0.000</td>
<td>0.000</td>
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<tr>
<td>(W_2)</td>
<td>0.782</td>
<td>1.297</td>
<td>0.009</td>
<td>0.063</td>
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<td>p-value</td>
<td>0.376</td>
<td>0.255</td>
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<tr>
<td>(W_3)</td>
<td>1.827</td>
<td>3.212</td>
<td>0.006</td>
<td>0.001</td>
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<tr>
<td>p-value</td>
<td>0.176</td>
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<tr>
<td>(W_4)</td>
<td>4.186</td>
<td>10.844</td>
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<td>p-value</td>
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</table>

Note 1: The null of \(W_1\) is that Hypothesis 1 is satisfied at impact.
Note 2: The null of \(W_2\) is that Hypothesis 2 is satisfied.
Note 3: The null of \(W_3\) is that Hypothesis 3 is satisfied.
Note 4: The null of \(W_4\) is that Hypotheses 2 and 3 are jointly satisfied.
Note 5: The null of \(W_5\) is that Hypotheses 1, 2, 3 are jointly satisfied.
Note 6: The null of \(W_6\) is that Hypothesis 1 is satisfied up to a year.
Table 4(a): FEVDs of the CA under Identification Scheme I

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|                 |          |          |          |          |          |          |          |          |
|                 | Periods  | GP       | CP       | CS       |          |          |          |          |
|                 |          |          |          |          |          |          |          |          |
|                 | 0        | 0.2958   | 0.0511   | 0.6531   | 0        | 0.1646   | 0.0282   | 0.8072   |
|                 |          | (0.2066) | (0.1062) | (0.1935) |          | (0.2343) | (0.1260) | (0.2353) |
|                 | 1        | 0.3098   | 0.1471   | 0.5431   | 1        | 0.2276   | 0.0396   | 0.7128   |
|                 |          | (0.2286) | (0.1454) | (0.1926) |          | (0.2505) | (0.1327) | (0.2426) |
|                 | 2        | 0.2917   | 0.2531   | 0.4552   | 2        | 0.2663   | 0.0691   | 0.6646   |
|                 |          | (0.2321) | (0.1675) | (0.1846) |          | (0.2584) | (0.1369) | (0.2468) |
|                 | 3        | 0.2745   | 0.3200   | 0.4055   | 3        | 0.2972   | 0.0726   | 0.6302   |
|                 |          | (0.2293) | (0.1764) | (0.1796) |          | (0.2636) | (0.1389) | (0.2489) |
|                 | 4        | 0.2646   | 0.3588   | 0.3766   | 4        | 0.3238   | 0.0738   | 0.6024   |
|                 |          | (0.2276) | (0.1820) | (0.1773) |          | (0.2673) | (0.1397) | (0.2500) |
|                 | 12       | 0.2656   | 0.4334   | 0.3010   | 12       | 0.4616   | 0.0675   | 0.4709   |
|                 |          | (0.2281) | (0.1981) | (0.1735) |          | (0.2811) | (0.1369) | (0.2516) |
|                 | 20       | 0.2784   | 0.4374   | 0.2842   | 20       | 0.5280   | 0.0615   | 0.4104   |
|                 |          | (0.2310) | (0.2013) | (0.1734) |          | (0.2868) | (0.1337) | (0.2529) |
|                 | 40       | 0.2887   | 0.4347   | 0.2766   | 40       | 0.5851   | 0.0558   | 0.3591   |
|                 |          | (0.2354) | (0.2024) | (0.1748) |          | (0.2941) | (0.1306) | (0.2570) |

Note 1: "GP", "CP" and "CS" represent global permanent, country-specific permanent and country-specific transitory shocks, respectively.
Note 2: The numbers in parentheses represent the standard errors based on 10000 nonparametric bootstrapping resamples.
Table 4(b): FEVDs of the CA under Identification Scheme II

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<table>
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Note 1: “GP”, “CP” and “CS” represent global permanent, country-specific permanent and country-specific transitory shocks, respectively.

Note 2: The numbers in parentheses represent the standard errors based on 10000 nonparametric bootstrapping resamples.
Table 5: Cross-Correlation Matrices of the Structural Shocks

(a) Identification Scheme I

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(b) Identification Scheme II

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Table A1  Unit Root Tests

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<td>-1.605</td>
<td>-1.816 *</td>
<td>-1.911 *</td>
</tr>
</tbody>
</table>

Note 1: The unit root tests are based on the ADF t-test. Since each variable is demeaned, the ADF regression does not include both constant and trend.

Note 2: ***, ** and * denote that the unit root null is rejected at 1%, 5% and 10% significance levels, respectively.

Note 3: Asymptotic 1%, 5% and 10% critical values are provided by Davidson and MacKinnon(1993) and equal to -2.56, -1.94 and -1.62, respectively.