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ASPECTS OF MACROECONOMIC SAVING

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Tillägnad mina föräldrar,
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Abstract

This thesis deals with various aspects of macroeconomic saving. It consists of an introduction and four self-contained papers.

Paper I, "From closed to open door policy: An empirical study of China's international capital mobility, 1958-98," is an empirical application of the open economy permanent income hypothesis (PIH). The purpose of this paper is to use the PIH to test and measure the degree of China's international capital mobility during the period 1958-98. In contrast to all previous known country studies using this framework, the hypothesis that capital has been at least mobile enough to allow for optimal consumption smoothing behavior is rejected. In this model, it appears that a country experiences suboptimally low capital mobility only when severe restrictions are placed on the capital account, as in the case of China. Partial barriers to international capital movements do not necessarily imply inability to smooth consumption optimally.

Paper II, "The open economy excess sensitivity hypothesis: Theory and Swedish evidence," extends the theory of open economy consumption behavior by applying Flavin's (1993) excess sensitivity hypothesis (ESH) to the current account. The ESH can be interpreted as a generalization of the PIH that allows for *any* degree of international capital mobility. As such, the ESH can account for why the PIH fails and for the related puzzle of an "excessively volatile" current account. Furthermore, the ESH suggests an alternative approach for assessing a country's degree of international capital mobility. Using annual Swedish data for the period 1951-99, the empirical results imply that, in contrast to the PIH, the ESH cannot be rejected. The results suggest that Sweden's degree of international capital mobility is higher than the degree that is perfect according to the PIH.

Paper III, "The PIH and the standard deviation ratio: A Monte Carlo Study," evaluates the coverage accuracy of small-sample confidence intervals for the standard deviation ratio summary statistic. This is the statistic that is used to test

and measure the degree of international capital mobility in the PIH framework. Three methods are considered to construct the confidence intervals: the asymptotic delta method, Runkle's (1987) standard bootstrap method, and Kilian's (1998) bias-corrected bootstrap-after-bootstrap method. Monte Carlo simulations suggest that the asymptotic delta method is unreliable and that researchers should rather use bias-corrected bootstrap confidence intervals when making inference from the standard deviation ratio.

Paper IV, "Has Sweden's government budget policy been too discretionary? Evidence from a generalization of the tax smoothing hypothesis," deals with the saving behavior of the government. Barro's (1979) tax smoothing hypothesis (TSH) assumes that the government is always subject to an "optimal" degree of discretion in budget policy, i.e., optimal in the sense that it minimizes the welfare costs from taxation. Paper IV proposes a generalization of the TSH that relaxes this crucial assumption. Postwar evidence for Sweden indicates that in contrast to the TSH, the generalized model provides close to a perfect fit: Tax smoothing behavior in combination with more discretion in budget policy relative to what is optimal, can explain all shifts in the central government's budget balance, including the dramatic shifts during the period 1970-96.

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Introduction and summary

Saving, which essentially can be thought of as future consumption, constitutes one of the most important parts of the economic system. Understanding saving and consumption is crucial for understanding how the economy, including variables such as GDP and unemployment, evolves over time. In essence, saving and consumption decisions are central to the standards of living in the long run.

1. Overview

In this thesis I present four self-contained papers that study and test various aspects of four closely related hypotheses of macroeconomic saving behavior: the open economy permanent income hypothesis (PIH), the open economy excess sensitivity hypothesis (ESH), the tax smoothing hypothesis (TSH), and the generalized tax smoothing hypothesis (GTSH).[1] One aspect of these hypotheses that is used repeatedly throughout the thesis, is to view them as being *rational expectations present value models* of the form

$$S_t = \pm(1 - \varphi) \sum_{s=t+1}^{\infty} \delta^{s-t} E_t \Delta X_s, \tag{1}$$

where δ is a known discount factor, φ is an estimable parameter, E_t is the expectations operator, Δ is the first difference operator, X_s is an observable exogenous stochastic variable, and S_t is the observable endogenous saving variable of interest. In the case of the PIH and ESH, S_t is the current account and X_s is an income measure, and the right hand side of (1) is preceded by a minus sign. In the case of the TSH and GTSH, S_t is the government's budget surplus and X_s is government

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[1] The "open economy" prefix is used only initially to stress that in this thesis, the PIH and the ESH examine changes in a country's net *foreign* assets, i.e., the current account, subject to changes in a broad income measure. Many papers that follow from Friedman's (1957) and Hall's (1978) seminal work, including those by Campbell (1987) and Flavin (1981, 1993), are concerned with a different aspect of the PIH, namely the accumulation of private wealth of all kinds in response to fluctuations in after-tax labor income. Such studies are often referred to as the PIH in a closed economy context.

expenditure, both expressed as fractions of output, and the right hand side of (1) is preceded by a positive sign.[2]

The fundamental assumption of all four hypotheses is the economic agent's "smoothing" behavior, and this is an assumption that always holds to some extent as long as $\varphi < 1$. In the case of the PIH and ESH, the agent is a country or a representative "national" individual that smooths consumption. From (1), this means that when future income is expected to increase, today's consumption is revised upward by means of borrowing abroad, which implies that today's current account goes into a deficit. In this way, instead of waiting until the actual increase in income materializes, today's consumption can be increased. That is, by borrowing abroad, it is possible for the country as a whole to smooth consumption over time. In the case of the TSH and GTSH, the government smooths the tax rate. For instance, when government expenditure is expected to increase, the TSH and GTSH imply that the tax rate is smoothed by means of running a budget surplus.

Campbell (1987) and Campbell and Shiller (1987) develop a method to test the implications of stochastic present value models such as (1). By means of (1) and a vector autoregression (VAR) for the bivariate $(\Delta X_t, S_t)$ process, one can calculate a theoretical or predicted time series of S_t. The *economic* significance of the model can then be evaluated by comparing the predicted saving series with the actual saving series in a time series plot. If the underlying hypothesis is true, i.e., if (1) holds, the predicted and actual series should coincide. The *statistical* significance of the model can be evaluated by testing the cross-equation restrictions of the VAR that are implied by (1). The distinction between economic and statistical significance can be important because, as Campbell and Shiller (1987) point out, even if the hypothesis that we study is statistically rejected, it is entirely possible that the predicted time

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[2] Equation (1) was first derived by Campbell (1987) and Campbell and Shiller (1987). It is a transformation of the standard present value model formulation (cf. Campbell and Shiller, 1987). Present value models of a similar structure as the ones in this thesis include the expectations theory for interest rates, the present value model of stock returns, the monetary model of exchange rates, the Cagan model of hyperinflation, and linear quadratic adjustment cost models. For a recent survey of these models, see Engsted (2002).

series of S_t does pretty well in tracking the actual series. Thus, even though the hypothesis under study is statistically rejected, the model may still contribute to our understanding of the saving behavior that we are trying to explain.

Ghosh (1995) uses the Campbell-Shiller method and the PIH to construct a measure and a test of the degree of international capital mobility. The measure is the ratio of the standard deviation (or variance) of the predicted current account to the standard deviation (or variance) of the actual current account. According to Ghosh, this *standard deviation ratio* captures the extent to which the degree of international capital mobility deviates from the degree that is perfect according to the PIH. If the model is true, i.e., if the PIH including the assumption of perfect capital mobility is true, the standard deviation ratio should equal unity. When the ratio is above unity, the standard deviation of the actual current account is lower than the standard deviation of the predicted current account. This suggests suboptimally low capital mobility and the presence of effective barriers to international capital movements. On the other hand, a higher standard deviation of the actual current account relative to the predicted current account, i.e., a ratio below unity, indicates a higher degree of capital mobility than the degree that is perfect according to the PIH. Ghosh argues that this "excess volatility" of the current account may suggest that speculation is to some extent a driving force behind capital flows.

The excess volatility of the current account and the related failure of the PIH is an empirical puzzle, at least in most industrialized countries (Sheffrin and Woo, 1990; Otto, 1992; Ghosh, 1995; and Obstfeld and Rogoff, 1995, 1996). For developing countries, the model appears to perform better. Ghosh and Ostry (1995) apply the model to forty-five developing countries including many with some form of restrictions placed on their international capital flows. They find that for a majority the PIH cannot be rejected, suggesting that capital mobility is sufficiently high to allow for optimal consumption smoothing. In fact, all known previous studies measuring capital mobility by means of Ghosh's (1995) method have generally been

unable to statistically conclude suboptimally low international capital mobility.

2.   Summary of the thesis

Paper I, "From closed to open door policy: An empirical study of China's international capital mobility, 1958-98," is a straightforward empirical application of the PIH. Using Ghosh's (1995) method, the purpose of the paper is to test and measure the degree of China's international capital mobility during the period 1958-98. The results indicate that in contrast to all previous studies using this framework, the hypothesis that capital has been at least mobile enough to allow for optimal consumption smoothing behavior is rejected. In summary, taken together with previous studies it appears that in the PIH model a country experiences suboptimally low capital mobility only when severe restrictions are placed on the capital account, as in the case of China. Partial barriers to international capital movements do not necessarily imply inability to smooth consumption optimally. In fact, less restrictive controls may even lead to excessive capital flows.

The purpose of Paper II, "The open economy excess sensitivity hypothesis: Theory and Swedish evidence," is to try to account for the failure of the PIH and the related puzzle of an excessively volatile current account by applying Flavin's (1993) excess sensitivity hypothesis (ESH) to the open economy setting. It is argued that the ESH can be interpreted as a generalization of the PIH that allows for *any* degree of international capital mobility. As such, the ESH can account for why the PIH fails and for the related puzzle of an excessively volatile current account. That is, the ESH provides a formalization of Ghosh's (1995) argument that the failure of the PIH is due to the violation of the assumption of perfect international capital mobility. Furthermore, the ESH suggests an alternative approach for assessing a country's degree of international capital mobility.

Using annual Swedish data for the period 1951-99, the empirical results of Paper II imply that, in contrast to the PIH, the ESH cannot be rejected. The results

suggest that Sweden's degree of international capital mobility is higher than the degree that is perfect according to the PIH. Following Ghosh (1995), one possible explanation may be that, in addition to changes in economic fundamentals, the current account is also to some extent driven by speculative capital flows.

Paper III, "The PIH and the standard deviation ratio: A Monte Carlo study," deals with a pure econometric issue that a researcher is confronted with when making statistical inference from on the standard deviation ratio. As noted above, testing the PIH involves the estimation of a VAR for the bivariate process (ΔX_t, S_t). The standard deviation ratio is a summary statistic that is a highly non-linear function of the coefficients of this VAR. As argued by Bekaert et al. (1997) and Kilian (1998), because of this non-linearity and because of bias in the VAR parameters, in small samples, asymptotic theory may be unreliable when making inference. The purpose of this paper is to investigate this matter by evaluating the coverage accuracy of small-sample confidence intervals for the standard deviation ratio. Three methods are used to construct the confidence intervals: the asymptotic delta method, Runkle's (1987) standard bootstrap method, and Kilian's (1998) bias-corrected "bootstrap-after-bootstrap" method. Monte Carlo simulations suggest that the asymptotic delta method is unreliable and that researchers should rather use bias-corrected bootstrap confidence intervals when making inference from the standard deviation ratio.

The purpose of Paper IV, "Has Sweden's government budget policy been too discretionary? Evidence from a generalization of the tax smoothing hypothesis," is to propose a generalization of Barro's (1979) tax smoothing hypothesis (TSH) and then empirically test whether it can explain the shifts in the Swedish central government's budget balance during recent decades. The TSH implicitly assumes that the government is always subject to an "optimal" degree of discretion in budget policy, i.e., optimal in the sense that it minimizes the welfare costs from taxation. By contrast, the *generalized* tax smoothing hypothesis (GTSH) proposed in this

paper takes into account the possibility that a tax smoothing government may be subject to a degree of discretion in budget policy that differs from the degree that is optimal according to the TSH. It is assumed that a specific degree of discretion in budget policy translates into a specific constraint on the government's borrowing and lending capabilities. The constraint can either be stricter or softer than the constraint that corresponds to the degree of borrowing and lending that is optimal according to the TSH.

The empirical results of Paper IV indicate that for the full sample period of 1952-99, neither the TSH nor the GTSH can be rejected. However, in the subperiod 1970-96, the TSH is rejected while the GTSH is not. Visually, the GTSH provides close to a perfect model fit; the predicted and actual budget surplus time series are almost identical throughout the entire sample period. The evidence suggests that besides tax smoothing behavior, there has been *more* discretion in Swedish budget policy *relative* to what is optimal according to the TSH.

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Paper I

From closed to open door policy: An empirical study of China's international capital mobility, 1958-98

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Abstract

This paper employs the intertemporal consumption smoothing approach to the current account to test the degree of China's international capital mobility during the period 1958-98. In contrast to all previous known country studies using this framework, the hypothesis that capital has been at least mobile enough to allow for optimal consumption smoothing behavior is rejected. In this model, it appears that a country experiences suboptimally low capital mobility only when severe restrictions are placed on the capital account, as in the case of China. Partial barriers to international capital movements do not necessarily imply inability to smooth consumption optimally.

Keywords: Capital mobility; China; Consumption smoothing; Current account
JEL classification: F32; F41; F47

1. Introduction

China's economic reforms and their implications for the country's growth have been remarkable. Since the introduction of the reforms in 1978, the country has grown at an annual average rate of over 9 percent, and real GNP per capita has almost quintupled. One of the reforms that has undoubtedly contributed significantly to the fast growth is the open door policy, initiated in 1979. In sharp contrast to the pre-reform policy aimed at self-sufficiency, the open door policy has led to increased international linkages in various forms. Key elements of the policy have included geographical targeting and decentralization of the management system, adoption of a system of import and export licensing and of tariffs and taxes, and liberalization of the exchange rate regime.[1] As a result, the country's foreign trade system has undergone a complete reorientation, transforming China into a country where foreign trade serves as the main engine of growth. At the end of 1996, China announced to the IMF that it met the requirements for current account convertibility. According to Chan et al. (1999), the process of marketization of China's foreign trade system was largely complete by 1997. In 1998, China's exports amounted to US\$184 billion, placing the country among the world's top ten exporters. Hence, from an insignificant pre-reform position, China is today one of the major international traders and also a significant participant in world capital markets.

The purpose of this paper is to test and measure the degree of China's international capital mobility during the period 1958-98 by means of the intertemporal consumption smoothing approach to the current account. According to this approach, when agents are smoothing consumption optimally, the current account is equal to the present discounted value of expected declines in *national cash flow*, where national cash flow is an income measure defined as output less government consumption less investment. As shown by Campbell (1987), by means of the theo-

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[1] For details of the reform of the foreign trade system, see for instance Lardy (1990) and Chan et al. (1999).

retical model and a vector autoregression for the current account and national cash flow, one can construct an optimal (predicted) time series of the current account. Then, following Ghosh (1995) and Ghosh and Ostry (1995), capital mobility can be assessed by visually comparing the derived optimal current account with the actual current account in a time series plot. If the model is true, i.e., if the assumptions of consumption smoothing and perfect international capital mobility hold, the two series should be identical.

More formally, the extent to which the degree of international capital mobility deviates from the degree that is perfect according to the consumption smoothing approach, is measured by studying the variance or standard deviation *ratio* of the optimal and actual current account series. Given the null hypothesis that the model is true, this ratio is equal to unity. If the ratio is above unity, the standard deviation of the actual current account is lower than the standard deviation of the optimal current account. This suggests suboptimally low capital mobility and the presence of effective barriers to international capital movements. On the other hand, a higher standard deviation of the actual current account relative to the optimal current account, i.e., a ratio below unity, indicates a higher degree of capital mobility than the degree that is perfect according to the consumption smoothing approach. Such an outcome may suggest that speculation is the driving force behind capital flows.

Studies using the consumption smoothing approach to test the degree of capital mobility have been done for countries with very different degrees of access to world capital markets. Ghosh (1995) concludes that for Japan, Germany, the United Kingdom, and Canada, capital mobility since World War II has been excessive, perhaps due to the presence of speculative forces.[2] Ghosh and Ostry (1995) find that for the majority of the 45 developing countries in their postwar sample, optimal and actual current account time series move closely together, suggesting that capital is mobile enough to allow for optimal consumption smoothing behavior, and that

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[2] The United States is the only country in Ghosh's study where the ratio of variances of the optimal and actual current account is not significantly different from unity.

effective barriers to mobility are relatively low. Similar results for other countries are obtained by Cashin and McDermott (1998), Agénor et al. (1999), Makrydakis (1999) and Hussein and de Mello Jr. (1999).[3] Of the studies done so far, covering more than 50 countries, the capital mobility null hypothesis of a variance (or standard deviation) ratio equal to unity has never been rejected in a way that suggests that mobility has been suboptimally low. That is, for all countries, the results have indicated an optimal-actual ratio equal to or significantly smaller than unity, which, from the consumption smoothing perspective, can be interpreted as perfect or excessive international capital mobility, respectively.[4]

In addition to the empirical evidence of excessive capital mobility, many countries have recently experienced severe financial crises that have been to some extent worsened by the resulting massive capital flows. Thus, considering both the results of previous empirical research and the recent financial crises, the question whether completely deregulated capital markets are optimal at all times is motivated. Of course, as shown below, the intertemporal consumption smoothing approach is based on many simplifying assumptions, which may not be valid. For a model to produce satisfactory results, it should be able to predict suboptimally low capital mobility when the true mobility is low. Since the model has been applied to countries with very restricted markets as well as to countries with a lower degree of regulation, but in general has indicated that capital has been at least mobile enough to allow for optimal consumption smoothing, it seems possible that the model systematically overstates the degree of capital mobility.[5] From this perspective, it is especially interesting to apply the model to China, since the country was perhaps the most

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[3] See also Sheffrin and Woo (1990) and Otto (1992). For a general survey of the theory and empirical results in this area, the reader should consult Obstfeld and Rogoff (1995, 1996).

[4] In some countries under study, the ratio had an estimated value greater than unity, which suggests relatively low capital mobility. In those cases, however, the null hypothesis of a ratio equal to unity could not be rejected.

[5] It should be noted that older approaches, the most famous being Feldstein and Horioka's (1980), suggest instead that international capital mobility amongst the major industrialized nations may be severely limited. For a discussion and criticism of the Feldstein-Horioka (1980) approach and other methods, see for instance Ghosh (1995) and Ghosh and Ostry (1995).

closed economy in the world during the first half of the sample period (1958-78), and then underwent gradual liberalization in the second half (1979-98). If the same conclusion is reached for China as for the other countries, it suggests that the consumption smoothing approach in its current form is an inadequate tool for measuring a country's degree of international capital mobility.

This paper is organized as follows. The next section outlines the theory for deriving and estimating the optimal current account. Section 3 contains the empirical estimation and results of tests of the model. Section 4 draws conclusions.

2.   Theory

The current account is defined as the change in a country's net foreign assets. In the balance of payments accounts, this is identically equal to capital flows including changes in reserves. Under the assumptions of consumption smoothing and perfect international capital mobility, the current account then serves as a buffer to smooth consumption in the face of expected changes in national cash flow.

In deriving the optimal current account series in line with these assumptions, the framework of Sachs (1982) is followed. Consider an open economy that produces a single tradable good. The economy is small in the sense that it faces a given real world interest rate. Normalizing population to 1, the single infinitely-lived consumer's preferences are given by

$$U_t = E_t \left\{ \sum_{s=t}^{\infty} \beta^{s-t} u(C_s) \right\} \qquad 0 < \beta < 1, \qquad (1)$$

where E_t is the conditional expectations operator; β is the subjective discount factor; $u(.)$ is the instantaneous utility function; and C_s denotes consumption of the single good. Assuming a constant real interest rate, r, the budget constraint is given by

$$B_{t+1} = (1+r)B_t + Q_t - C_t - G_t - I_t, \qquad (2)$$

which, upon rearrangement, gives

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$$CA_t = B_{t+1} - B_t = rB_t + Q_t - C_t - G_t - I_t, \tag{3}$$

where CA_t is the current account; B_{t+1} is net foreign assets at the end of period t; rB_t is interest earned on previously acquired assets; Q_t is the level of output; G_t is the level of government consumption; and I_t is the level of investment.[6] The problem, then, is to maximize (1) subject to (2), with a transversality condition imposed. To obtain a simple and empirically tractable closed form solution for consumption, a quadratic form for the utility function is posited, and it is assumed that $\beta(1+r) = 1$. The solution can be written as

$$C_t = rB_t + \frac{r}{1+r} \sum_{s=t}^{\infty} \left(\frac{1}{1+r} \right)^{s-t} E_t Z_s, \tag{4}$$

where Z_s is national cash flow, defined as

$$Z_s = Q_s - G_s - I_s. \tag{5}$$

The small country assumption implies that output and investment may be treated as exogenous to the consumption decision. With a given interest rate, the agent undertakes investment independently of domestic consumption-preferences up to the point where the marginal product of capital equals the world interest rate. Hence, investment and output are chosen independently of consumption. Government consumption is assumed to be exogenous as well.

Equation (4) is an open economy version of the permanent income hypothesis. It states that consumption is equal to permanent national cash flow, or annuity value of wealth, and as such, it captures the notion of consumption smoothing. As Campbell (1987) points out, another way to see this is to formulate the hypothesis as a statement about saving, i.e., about the current account. Substitute (4) into (3) and rearrange the terms such that

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[6] The model assumes that a riskless bond is the only internationally traded asset.

$$CA_t = -\sum_{s=t+1}^{\infty} \left(\frac{1}{1+r}\right)^{s-t} E_t \Delta Z_s, \tag{6}$$

where $\Delta Z_s = Z_s - Z_{s-1}$. Thus, according to the consumption smoothing approach, the current account is equal to the present discounted value of expected declines in national cash flow. Obviously, when future positive changes in national cash flow are expected, today's current account goes into deficit, since agents borrow in order to smooth consumption over time. Conversely, if agents expect future declines in national cash flow, the current account goes into surplus. This is also consumption smoothing behavior, or equivalently, as Campbell (1987) puts it, people save "for a rainy day."

Campbell (1987) and Campbell and Shiller (1987) develop an empirical method to estimate $E_t \Delta Z_s$ and test the validity of (6), as follows. Note first that the current account can be written as

$$CA_t = rB_t + Z_t - C_t = \begin{bmatrix} 1 & 1 & -1 \end{bmatrix} \begin{bmatrix} rB_t \\ Z_t \\ C_t \end{bmatrix} = \begin{bmatrix} \delta_1 & \delta_2 & \delta_3 \end{bmatrix} X_t = \delta' X_t. \tag{7}$$

Next, assume that Z_t is stationary in first differences. Then it follows from (6) that CA_t is stationary in levels, since it is a linear combination of expected changes in Z_t. According to (7), given that rB_t and C_t are stationary in first differences, the components of the vector X_t are cointegrated, and δ is the cointegrating vector. By the Granger representation theorem (see Engle and Granger, 1987), it follows that an error correction representation exists:

$$\Delta X_t = m + LB(L)\Delta X_t + \gamma \delta' X_{t-1} + \varepsilon_t, \tag{8}$$

where m is a 3×1 vector of intercept terms; $B(L)$ is a matrix polynomial in the lag operator; γ is a 3×1 speed of adjustments vector; and ε_t is a 3×1 vector of white noise

disturbances. Using (8) for forecasting purposes involves the cumbersome recursive updating of X_{t-1}. Fortunately, this problem can be circumvented by transforming (8) into the vector autoregression (VAR)

$$\begin{bmatrix} \Delta C_t \\ \Delta Z_t \\ CA_t \end{bmatrix} = c + LA(L) \begin{bmatrix} \Delta C_t \\ \Delta Z_t \\ CA_t \end{bmatrix} + G\varepsilon_t, \tag{9}$$

where

$$G = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ \delta_1 & \delta_2 & \delta_3 \end{bmatrix}, \tag{10}$$

and

$$c = Gm. \tag{11}$$

Ooms (1994) provides a proof of this transformation. Using Hall's (1978) well-known result that consumption in this setting follows a martingale, it follows that the first equation in (9) is redundant. Thus, (9) reduces to a VAR in $(\Delta Z_t, CA_t)$, and this is the empirical model from which the forecast of changes in national cash flow, $E_t \Delta Z_s$, is obtained. Given that the data is annual and that the sample is relatively small, this forecast is assumed to be based on the first-order VAR

$$\begin{bmatrix} \Delta Z_t \\ CA_t \end{bmatrix} = \begin{bmatrix} \Psi_{11} & \Psi_{12} \\ \Psi_{21} & \Psi_{22} \end{bmatrix} \begin{bmatrix} \Delta Z_{t-1} \\ CA_{t-1} \end{bmatrix} + \begin{bmatrix} u_{1t} \\ u_{2t} \end{bmatrix}, \tag{12}$$

where the means of ΔZ_t and CA_t have been removed.

The main interest in (12) concerns the regression in which ΔZ_t is the dependent variable. It is the present discounted value of all date t forecasts of this variable, conditional on the agent's full information set, that will determine the optimal date t current account. That is, according to (6), future expected changes in national

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cash flow are reflected in today's current account. Then, intuitively, not only will ΔZ_{t-1} be important in determining ΔZ_t, but also CA_{t-1} should help predict ΔZ_t, since it may contain additional information. Hence, an implication of the model is that the current account should Granger cause changes in national cash flow.

The forecast of a one period change in national cash flow is then

$$E_t \Delta Z_s = \begin{bmatrix} 1 & 0 \end{bmatrix} \begin{bmatrix} \Psi_{11} & \Psi_{12} \\ \Psi_{21} & \Psi_{22} \end{bmatrix}^{s-t} \begin{bmatrix} \Delta Z_t \\ CA_t \end{bmatrix} = \begin{bmatrix} 1 & 0 \end{bmatrix} \Psi^{s-t} \begin{bmatrix} \Delta Z_t \\ CA_t \end{bmatrix}, \quad (13)$$

where Ψ is the 2×2 matrix of coefficients, Ψ_{ij} $(i, j = 1, 2)$. Substituting (13) into (6) and simplifying, gives

$$\widehat{CA}_t = -\begin{bmatrix} 1 & 0 \end{bmatrix} \left(\frac{1}{1+r} \Psi \right) \left(I - \frac{1}{1+r} \right) \Psi^{-1} \begin{bmatrix} \Delta Z_t \\ CA_t \end{bmatrix} =$$

$$\begin{bmatrix} \Phi_1 & \Phi_2 \end{bmatrix} \begin{bmatrix} \Delta Z_t \\ CA_t \end{bmatrix} = \Phi' \begin{bmatrix} \Delta Z_t \\ CA_t \end{bmatrix}, \quad (14)$$

where $\widehat{CA}_t$ is the optimal current account, and each element in Φ is a nonlinear function of the underlying VAR parameters. As noted, the optimal current account series, $\widehat{CA}_t$, can be compared to the actual series, CA_t. If the model is true, the two series should be identical. That is, if the model is true, it follows that

$$\widehat{CA}_t = \begin{bmatrix} 0 & 1 \end{bmatrix} \begin{bmatrix} \Delta Z_t \\ CA_t \end{bmatrix} = CA_t. \quad (15)$$

Thus, the parameter restrictions can be written

$$\begin{bmatrix} 0 & 1 \end{bmatrix} = -\begin{bmatrix} 1 & 0 \end{bmatrix} \left(\frac{1}{1+r} \Psi \right) \left(I - \frac{1}{1+r} \Psi \right)^{-1}, \quad (16)$$

or, equivalently, by postmultiplying by $\left(I - \frac{1}{1+r} \Psi \right)$:

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$$\begin{bmatrix} 0 & 1 \end{bmatrix} \left(I - \frac{1}{1+r}\Psi \right) = -\begin{bmatrix} 1 & 0 \end{bmatrix} \left(\frac{1}{1+r}\Psi \right)$$

$$\Longleftrightarrow$$

$$\begin{bmatrix} \Psi_{21} & (\Psi_{22} - (1+r)) \end{bmatrix} = \begin{bmatrix} \Psi_{11} & \Psi_{12} \end{bmatrix}. \tag{17}$$

In summary, the model derived in this section sets the following agenda for the empirical estimation. First, to verify that national cash flow is stationary in first differences and that the current account time series is stationary, unit root and cointegration analyses are performed. Next, the VAR is estimated, and Granger causality of the current account is tested for. By means of the estimated VAR parameters, the optimal current account series according to (14) is calculated and visually compared to the actual current account series. Then, the degree of capital mobility is assessed formally by testing the null hypothesis that the standard deviation ratio of the optimal and actual current account series is equal to unity. Finally, the validity of the overall model is evaluated by testing the linear restrictions in (17).

3.   Empirical estimation and results

Annual national account data from China's National Bureau of Statistics are used, and all variables are expressed in 1978 yuan per capita. Further details concerning the data are provided in the Appendix. The full estimation period runs from 1958 to 1998. The period of liberalization (1979-98) is of course of special interest, and is therefore studied separately, but to be complete and to be able to compare results, the model is also estimated for the pre-reform period (1958-78).

Table 1
Unit root tests

| | 1958-98 | | 1958-78 | | 1979-98 | |
Variable	ADF	LM	ADF	LM	ADF	LM
C_t	3.94	1.04	-1.59	3.43	3.12	1.24
Z_t	7.67	0.60	-1.77	3.33	2.91	0.08
V_t	7.30	0.45	-1.77	3.33	9.01	0.08
CA_t	-3.25^{**}	1.23	-4.08^{**}	0.85	-2.24^{**}	0.72
ΔC_t	-3.74^*	0.27	-6.40^{**}	2.26	-3.69^*	0.88
ΔZ_t	-3.79^*	0.60	-6.38^{**}	2.06	-3.33	0.34
ΔV_t	-4.58^{**}	2.05	-6.38^{**}	2.06	-3.81^*	0.28

Notes: LM is the Lagrange multiplier test for residual correlation from lags 1 to 2. "**" indicates rejection at the 1 percent level of significance. "*" indicates rejection at the 5 percent level of significance.

The first step in the empirical analysis is to verify that all involved variables have the required order of integration. The approach is to write the current account as

$$CA_t = V_t - C_t = \begin{bmatrix} 1 & -1 \end{bmatrix} \begin{bmatrix} V_t & C_t \end{bmatrix}', \tag{18}$$

where V_t is *total* cash flow, defined as $GNP_t - G_t - I_t$.[7] Accordingly, it is necessary to verify that V_t and C_t are both individually I(1) and cointegrated such that CA_t is I(0). In addition, the model also assumes that Z_t is I(1). To verify the order of integration of the series, augmented Dickey-Fuller (ADF) tests are performed (for details, see Dickey and Fuller, 1981). First, models with 4 lags of the first difference of the dependent variable are estimated. Then, if possible, lags are reduced according to model reduction criteria, which are mainly F-tests for lag reduction and Lagrange Multiplier tests for residual autocorrelation. Table 1 displays the test results. For C_t, Z_t, and V_t, the null hypothesis of a unit root cannot be rejected. By contrast, for CA_t, the null is rejected, which suggests that V_t and C_t are cointegrated. The ADF test for a unit root in the first difference of C_t, Z_t, and V_t is also performed. Now the

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[7] Differently put, V_t is equal to the sum of national cash flow, Z_t, and interest earned on previously acquired foreign assets, rB_t.

null is rejected, except for ΔZ_t in the period 1979-98. However, the null can safely be rejected at 10 percent, and given the short sample and that it is well-known that the power of the ADF test is low, it seems reasonable to assume that Z_t is I(1).

The Dickey-Fuller test for CA_t is supplemented by a Johansen (1988) cointegration test. The result of this test is provided in Table 2. The null hypothesis is that the cointegrating vector $(1, -1)$ belongs to the cointegrating space, such that $(1, -1)(V_t, C_t)' = CA_t$ is I(0). As the table shows, it is not possible to reject the null in any of the time periods. In summary, the results from the unit root and cointegration analyses indicate that C_t, Z_t, and V_t are stationary in their first differences, while CA_t is stationary in levels.

Table 2
Restricted cointegration tests

	1958-98	1958-78	1979-98
χ^2	2.37	2.03	0.50
p-value	0.12	0.15	0.48

Notes: χ^2 is the test statistic for the null hypothesis that the restricted cointegrating vector $(1, -1)$ belongs to the cointegrating space, such that $(1, -1)(V_t, C_t)' = CA_t$ is I(0).

Table 3 shows the main results of the VAR analysis. Note that the parameter estimates for the whole period and for the period 1979-98 are very similar. By contrast, the estimates for the period 1958-78 differ, especially for Ψ_{12}. The reason is that, although the current account moves and is different from zero in the pre-reform period, the magnitude of the movements in the liberalization period is much larger. The table also shows the results of the test of the null hypothesis that CA_{t-1} non-Granger causes ΔZ_t. For the full period and for the period 1979-98, it is possible to reject the null at the 7.8 percent level of significance. At first sight, the corresponding test result for the period 1958-78 appears stronger, since it is possible to reject the null at 4.5 percent. However, this result is not in accordance with the model, since according to (6), CA_t predicts future *declines* in ΔZ_t. The sign of

$\widehat{\Psi}_{12}$ for the period 1958-78 is positive, contradicting this prediction. By contrast, during the period 1979-98, the implication of the consumption smoothing model is satisfied, since CA_{t-1} negatively Granger causes ΔZ_t.

Table 3
Estimated VAR parameters

	1958-98	1958-78	1979-1998
$\widehat{\Psi}_{11}$	0.934	0.522	0.917
	(0.089)	(0.183)	(0.105)
$\widehat{\Psi}_{12}$	-0.824	3.923	-0.831
	(0.466)	(1.735)	(0.471)
$\widehat{\Psi}_{21}$	0.116	0.055	0.131
	(0.086)	(0.019)	(0.112)
$\widehat{\Psi}_{22}$	0.324	0.495	0.317
	(0.228)	(0.129)	(0.226)
$\chi^2_{Granger}$	3.12	5.11	3.12
p-value	0.077	0.024	0.077

Notes: The standard errors (within parentheses) are White (1980) heteroscedasticity consistent. $\chi^2_{Granger}$ is the test statistic for the null hypothesis that CA_{t-1} non-Granger causes ΔZ_t.

The VAR parameters are used to derive the optimal current account time series. Following Sheffrin and Woo (1990), results are presented for two different cases; one based on an interest rate of 4 percent per year, and the other based on a rate of 14 percent.[8] By visually comparing the optimal and actual series, it is possible to assess the model's performance and the degree of China's international capital mobility. Figure 1 shows the actual and optimal series for each of the two subperiods. As expected, the model performs miserably during the period 1958-78, and the optimal current account differs substantially from the actual current account. But there is dramatic improvement following the liberalization in 1979. Since then, the optimal current account seems to have done pretty well in capturing the shifts of the actual series. Still though, both of the optimal series, especially

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[8] Sheffrin and Woo's use of a 14 percent rate of interest stems from Bernanke's (1985) study of consumption-income relations.

the one based on the 4 percent interest rate, are more volatile compared to the actual current account. Thus, the visual evidence suggests that, although China has undergone much liberalization since 1979, there are still effective barriers to the country's international capital mobility.
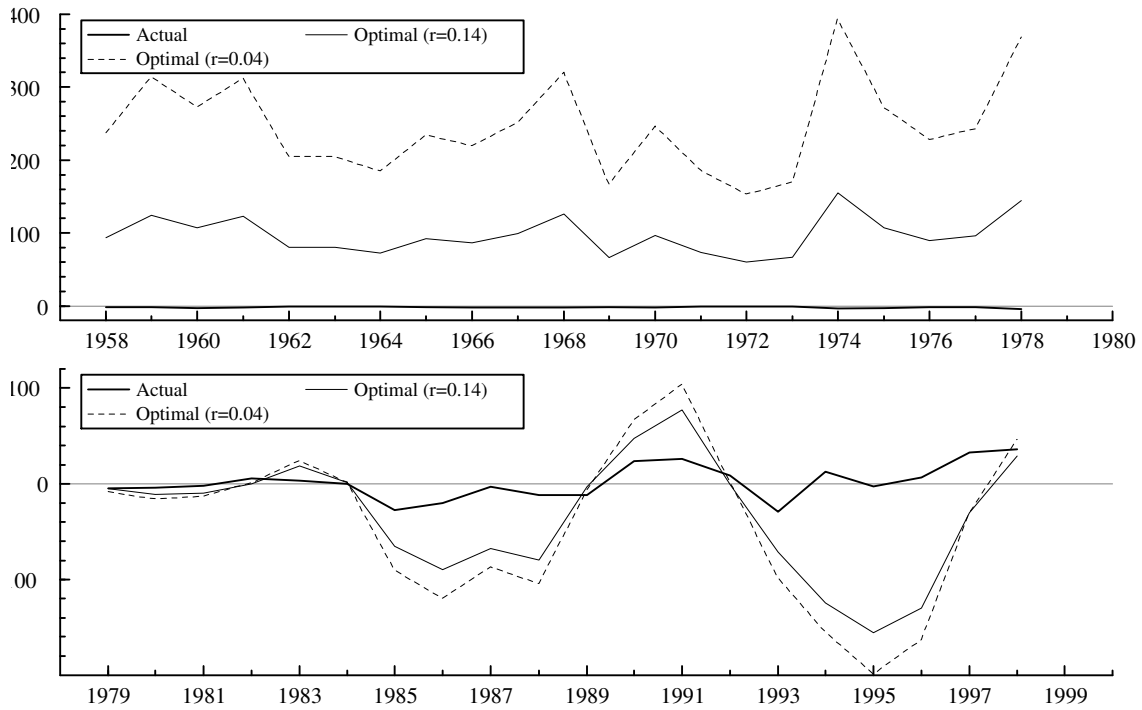
Fig.1. *Optimal and actual current account series. Notes:* Upper panel: 1958-78; lower panel: 1979-98.

The tests of the model confirm the visual analysis. Table 4 reports the estimated standard deviation ratios. The table also displays the estimated correlation between the optimal and actual current series. As Makrydakis (1999) points out, since the optimal current account is based on a forecast and is a highly nonlinear function of the VAR parameters, there might be significant small sample bias inherent in statistics stemming from it. Therefore, to reduce the possibility of erroneous conclusions, bootstrapping is used in order to make inference about capital mobility from the estimated standard deviation ratios. Bootstrapping is also used to make inference

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about the correlation coefficients.

However, while Makrydakis (1999) uses bootstrapping to obtain standard errors, the approach in this paper is different. As pointed out by Efron and Tibshirani (1993), using standard normal theory assumptions to make inference based on the bootstrap estimate of standard error may lead to erroneous conclusions when the sample size is small. In this case, it is better to use the percentiles of the bootstrap histogram to construct confidence limits. Furthermore, as emphasized by Bekaert et al. (1997) and Kilian (1998), in a VAR framework, caution with respect to bias in the ordinary least-squares estimates has to be taken before calculating these percentiles. To take this into account, Kilian (1998) develops a bias-correcting "bootstrap-after-bootstrap" method, which is used here. The method is basically divided into two parts. First, based on 1000 artificial series of ΔZ_t and CA_t, 1000 bootstrap replications of each VAR parameter are obtained in order to estimate and correct for bias in each of the original parameters.[9] Then, by use of these bias-corrected VAR estimates, another 2,000 artificial series of ΔZ_t and CA_t are produced. For each series of ΔZ_t and CA_t, the four VAR parameters are estimated and, again, bias-corrected. Next, the ratio of standard deviations and correlation of the optimal and actual current account series are calculated. Repeating this procedure yields 2000 estimated standard deviation ratios and correlation coefficients. Finally, 95 percent confidence intervals are constructed from the 2.5th and 97.5th percentile interval endpoints of the two resulting empirical distributions.[10] These endpoints are printed within brackets beneath the corresponding statistic in Table 4.

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[9] This is done by means of sampling with replacement from the original VAR residuals. To initialize the procedure, two initial observations are randomly selected using the block method of Stine (1987).

[10] Kilian (1998) applies the bootstrap-after-bootstrap algorithm to obtain confidence intervals for impulse response functions. In addition to the brief description given here, the algorithm also involves a stationarity correction procedure if the modulus of the largest root of the involved VAR companion matrices is greater than or equal to unity. For complete details of the bootstrap-after-bootstrap method and for proof of its asymptotic validity, see Kilian's paper.

<div align="center">

Table 4

Tests of the model

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	1958-98		1958-78		1979-98	
	4%	14%	4%	14%	4%	14%
Ratio	6.45*	4.83*	75.11*	29.44*	4.45*	3.41*
	[2.57 28.86]	[2.08 13.59]	[15.12 174.2]	[11.41 56.21]	[1.42 35.72]	[1.23 14.02]
Corr	0.27*	0.23*	-0.83*	-0.83*	0.49*	0.44*
	[-0.34 0.76]	[-0.37 0.71]	[-0.99 -0.55]	[-0.99 -0.55]	[-0.47 0.89]	[-0.50 0.84]
χ^2_W	77.3**	80.4**	110.1**	112.6**	43.1**	45.4**

Notes: Ratio is the estimated ratio of standard deviations of the optimal and actual current account series. *Corr* is the estimated correlation between the optimal and actual current account series. χ^2_W is the Wald test statistic of the restrictions implied by the model. The numbers within brackets are the 95 percent confidence interval endpoints for each statistic, constructed from the bootstrap simulations. "**" indicates rejection at the 1 percent level of significance. "*" indicates rejection at the 5 percent level of significance.

Under the null hypothesis of optimal consumption smoothing and perfect international capital mobility, the correlation coefficient and standard deviation ratio should both equal unity. From the table it is evident that, irrespective of the interest rate used, the null is not contained in the confidence interval for the standard deviation ratio; not even for the 1979-98 period. Since the calculated confidence intervals only contain ratios that are greater than unity, this suggests that capital mobility has been suboptimally low. The confidence intervals for the correlations also do not contain the null. Note that, for any given time period, the correlation coefficients and their corresponding confidence intervals are quite similar under both interest rates. By contrast, the precision with which the standard deviation ratios are estimated in the different periods greatly improves in the 14 percent interest rate case. In general, the model seems to work a little better with the higher interest rate.

Finally, Table 4 also reports the formal test of the whole model. If the consumption smoothing model is true, the restrictions in (17) hold, and it follows that $\Phi' = (0, 1)$ such that the optimal current account is equal to the actual current

account. The Wald test of these restrictions for each period and interest rate is presented in the bottom row of the table. It is evident that, regardless of which period and interest rate one chooses to look at, the restrictions implied by the model are strongly rejected.

4. Conclusions

The visual and statistical evidence presented in this paper indicate that the consumption smoothing model is rejected for China. Specifically, using interest rates of 4 and 14 percent, estimated standard deviation ratios of the optimal and actual current account series are significantly greater than unity, even during the period of liberalization (1979-98). Among the approximately 50 countries studied so far by means of the intertemporal consumption smoothing approach to the current account, China stands out as the only country where such a result has been obtained. This may make sense, however, since the country was a more or less closed economy in the first half of the sample period, and then underwent gradual liberalization in the second half.

Comparing the standard deviation ratios for the periods 1958-78 and 1979-98 in Table 4, it is obvious that they are considerably smaller in the latter period. Also, from the optimal and actual current account series during the period 1979-98 displayed in Figure 1, it is visually evident that the optimal current account captures most shifts in the actual current account. A possible interpretation is hence that international capital mobility has indeed increased dramatically during the past two decades, but is still persistently suboptimal due to remaining effective barriers.

The results lend support to the intertemporal consumption smoothing approach in the sense that it is now evident that the model is capable of predicting in all directions, either indicating that capital mobility has been excessive, as in Ghosh (1995); or that it has been in line with optimal consumption smoothing, as for instance in Hussein and de Mello Jr. (1999); or that it has been suboptimally

low, as in this paper. Given this model and its assumptions, the message then appears to be that, in order to experience suboptimally low capital mobility, one needs to have severe restrictions, such as currency inconvertibility under the capital account.[11] As suggested by the results of Ghosh and Ostry (1995), partial controls do not necessarily imply inability to smooth consumption optimally. In fact, less restrictive controls may even lead to excessive capital flows.

However, it must be emphasized that there are many simplifying assumptions underlying this model, and hence several joint hypotheses that are subject to the tests, and model rejection can thus not be attributed to any single element. Remember that in addition to perfect capital mobility and consumption smoothing behavior, assumptions include a single tradable good, a constant interest rate, and exogenous government expenditure in a very simple setup. Deeper insights on capital mobility issues in this framework will require relaxation of these assumptions and further research.

Appendix

The annual data used in this paper are taken from *The Historical National Accounts of the People's Republic of China 1952-1995* and *China Statistical Yearbook 1999*. All data are expressed in 100 million yuan and converted into real terms per capita by dividing by the implicit GDP deflator (1978=100) and population. The usual definitions apply, with one small exception. In the data set, the following is an accounting identity:

$$GDP_t - D_t \equiv Q_t \equiv C_t + I_t + G_t + NX_t, \qquad (A.1)$$

where GDP_t is the gross domestic product, Q_t is the gross domestic expenditure

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[11] In 1997, the Chinese government announced the goal of achieving capital account convertibility by the year 2000. Due to the Asian economic crisis, the timetable has changed, and the government now emphasizes a "gradual" liberalization (Groombridge, 2001). In 2001, capital controls in China included, for example, controls for foreign direct investments of residents and portfolio investments of both residents and nonresidents (for an overview of the controls, see Hu, 2001).

(or, in China, GDP by expenditure approach), C_t is private consumption, I_t is investment, G_t is government consumption, and NX_t is net exports of goods and services. D_t is a positive or negative discrepancy defined as $GDP_t - Q_t$, resulting from the transformation of data from the Chinese Material Product System (MPS) into the System of National Accounts (SNA). In practice, D_t is small, in general below 3 percent of GDP, with an average of 0.13 percent of GDP. For complete details of the estimation and sources of the aggregates, and the transformation from MPS to SNA, the reader should consult the first of the publications mentioned above.

Besides the aggregates in (A.1), gross national product, GNP_t, is also given in the data set. The current account is then defined as in (18), i.e.,

$$CA_t = V_t - C_t = GNP_t - G_t - I_t - C_t. \tag{A.2}$$

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Paper II

The open economy excess sensitivity hypothesis: Theory and Swedish evidence

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Abstract

This paper extends the theory of open economy consumption behavior by applying Flavin's (1993) excess sensitivity hypothesis (ESH) to the current account. The ESH can be interpreted as a generalization of the open economy permanent income hypothesis (PIH) that allows for *any* degree of international capital mobility. As such, the ESH can account for why the PIH fails and for the related puzzle of an "excessively volatile" current account. Furthermore, the ESH suggests an alternative approach for assessing a country's degree of international capital mobility. Using annual Swedish data for the period 1951-99, the empirical evidence imply that, in contrast to the PIH, the ESH cannot be rejected.

*Keywords***:** Excess sensitivity; Permanent income; Consumption; Current account; Capital mobility
JEL classification: E21; F32

1. Introduction

The open economy permanent income hypothesis (PIH) implies that when a country expects a change in future income, it uses the current account to smooth consumption intertemporally. Consequently, the current account, which by definition is equal to the change in the country's net foreign assets, measures saving for the economy as a whole. In the case of an expected decline in income, the current account will be in surplus, as the PIH implies that the country will save "for a rainy day" (Campbell, 1987). Conversely, if income is expected to rise, consumption will be smoothed by means of borrowing in world capital markets, and the current account will show a deficit.

By means of a vector autoregression (VAR) for income and the current account, it is possible to test the restrictions that are implied by the PIH. Moreover, it is possible to visually characterize the fit of the model. From the VAR and the underlying theory, one can construct a prediction of the current account. An assessment of the model can then be made by comparing the actual current account with the predicted current account in a time series plot. If the PIH is true, then the two series should be identical.[1]

Empirically, the PIH usually fails, and it fails in a way that makes the actual current account fluctuate more than the predicted current account. This "excess volatility" of the current account is a puzzle that appears to be present in most industrialized countries. For instance, using postwar data for Canada, Germany, Japan, and the United Kingdom, Ghosh (1995) rejects the PIH and finds that the current accounts of these countries are excessively volatile. The United States is the only country in Ghosh's sample where the evidence is in favor of the PIH. Other

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[1] The method is due to Campbell (1987) who originally applied the PIH model to investigate saving decisions of individuals subject to labor income (see also Campbell and Shiller, 1987). Such a setup is often referred to as the PIH in a closed economy context. The extension of the PIH to the open economy, i.e. the extension of the PIH to the current account, using Campbell's method is due to Sheffrin and Woo (1990), Otto (1992), and Ghosh (1995). For a survey, see Obstfeld and Rogoff (1995).

studies, including Sheffrin and Woo (1990), Otto (1992), and Obstfeld and Rogoff (1995, 1996) reach similar conclusions. For developing countries, the model appears to perform better. Ghosh and Ostry (1995) apply the model to forty-five developing countries and find that, for a majority, the PIH cannot be rejected. Still, for most countries in their study, the results indicate, at least visually, that the actual current account is more volatile than the predicted current account.

The purpose of this paper is to try to account for the failure of the PIH and the related puzzle of an excessively volatile current account by applying Flavin's (1993) excess sensitivity hypothesis (ESH) to the open economy setting. Flavin's work, which follows from Hall's (1978) seminal paper, is concerned with studying consumption and total financial saving of individuals subject to labor income. As Hall shows, given strict restrictions on tastes, the PIH implies that consumption follows a martingale and thus is orthogonal to lagged changes in other variables. By contrast, the ESH posits that consumption exhibits sensitivity to current income. The corresponding income measure in the open economy setting is *national cash flow*, defined as output less investment less government consumption. Consequently, the open economy version of the ESH posits that consumption exhibits sensitivity to current national cash flow.

Section 2.1 extends Flavin's (1993) work on the ESH to the open economy setting, i.e., to the current account. The application is straightforward, since the method of extending the ESH to the current account closely follows that of extending the PIH to the current account. This is because, just as Flavin points out, the PIH can be regarded as a special case of the ESH. The section discusses the main properties of the ESH and then, Section 2.2 shows how these properties translate into testable cross-equation VAR restrictions. It is also shown how the method of constructing the predicted current account directly carries over to the ESH framework.

Section 2.3 shows that the ESH is a generalization of the PIH that can be interpreted as relaxing the assumption of perfect international capital mobility. In

contrast to the PIH, the ESH allows for *any* degree of international capital mobility that is consistent with the fundamental objective of smoothing consumption intertemporally. From this, it follows that the theoretical properties of the ESH can be reconciled with Ghosh's (1995) argument that the failure of the PIH and the excess volatility of the current account is due to "too much" capital mobility. Accordingly, the ESH offers a formal solution to the puzzle of an excessively volatile current account. Furthermore, the ESH suggests an alternative approach to Ghosh's method for measuring the degree of a country's international capital mobility.[2]

Section 3 uses Swedish postwar data to test the PIH and the ESH. Sweden is a small open economy that clearly illustrates what is common when the PIH is applied to an industrialized country's current account.[3] Although the model provides a reasonable fit, the restrictions implied by the PIH can be rejected, and the current account is excessively volatile. By contrast, this paper shows that when it is assumed that consumption is generated by the ESH, the model cannot be rejected, and the actual and predicted current account series are almost identical. That is, when consumption is assumed to be generated by the ESH, the puzzle of an excessively volatile current account is solved. Section 4 provides some concluding remarks.

2. The open economy excess sensitivity hypothesis

2.1 The model

The extension of the ESH to the open economy builds on work by Campbell (1987), Campbell and Deaton (1989), and most notably, by Flavin (1993). Compared to the closed economy setting, the open economy setting focuses on a broader measure of income and a narrower category of assets. The income measure in the open economy

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[2] An approach for assessing a country's degree of international capital mobility that is similar to the one explored in this paper is given by Shibata and Shintani (1998). However, a direct comparison with their work is difficult because they use a slightly different underlying theory for consumption, a different line of theoretical reasoning, and a different empirical method (see f.n. 11).

[3] Obstfeld and Rogoff (1995, 1996) present visual evidence for Sweden and other industrialized countries including Belgium, Canada, Denmark, and the United Kingdom.

setting is national cash flow, Z_t, defined as

$$Z_t = Q_t - G_t - I_t, \tag{1}$$

where Q_t is the level of real per capita output; G_t is the level of real per capita government consumption; and I_t is the level of real per capita investment. It is assumed that national cash flow follows an exogenous stochastic process that is stationary in first differences. The asset category of interest is real per capita net foreign claims, B_t, which evolve over time according to

$$B_{t+1} = (1+r)B_t + Z_t - C_t, \tag{2}$$

where r denotes the world rate of real interest (assumed constant) and C_t denotes the level of real per capita consumption. The current account, CA_t, is then defined as the change in net foreign assets, i.e.,

$$CA_t = B_{t+1} - B_t = rB_t + Z_t - C_t. \tag{3}$$

Taking the first difference and rearranging, gives

$$\Delta C_t = (1+r)CA_{t-1} - CA_t + \Delta Z_t. \tag{4}$$

To test the model, one can impose on the identity (4) the derived form for CA_t implied by the ESH, and then use a VAR for CA_t and ΔZ_t to evaluate the resulting restrictions. The details regarding this are outlined in Section 2.2. Just as with the PIH, it is also possible to use the VAR to construct a prediction of the current account and then compare it to the actual current account series. If the ESH is true, the two series should be identical. However, as Section 2.3 argues, the economic interpretation of this comparison is very different from when consumption is assumed to be generated by the PIH.

In its original form, the ESH is a statement about consumption in levels. In order to derive it, consider first the open economy counterpart of Flavin's (1981)

permanent income, i.e. *permanent national cash flow, Z_t^P*, which is defined as the annuity value of the sum of net foreign assets and the present discounted value of expected future national cash flow. That is,

$$Z_t^P = rB_t + \frac{r}{1+r} \sum_{s=t}^{\infty} \left(\frac{1}{1+r}\right)^{s-t} E_t Z_s, \tag{5}$$

where E_t is the expectations operator, conditional on the agent's complete information set at t, I_t. Next, define *transitory national cash flow, Z_t^T*, as the residual

$$Z_t^T = V_t - Z_t^p = Z_t + rB_t - Z_t^p, \tag{6}$$

where V_t is total cash flow defined as $Z_t + rB_t$, i.e. the sum of national cash flow and interest earned on previously acquired assets. In this open economy setting, Flavin's (1993) excess sensitivity hypothesis of consumption behavior becomes

$$C_t = \beta Z_t^T + Z_t^P, \tag{7}$$

where β is the excess sensitivity parameter. Thus, when β is different from zero, consumption is excessively sensitive in the sense that it depends not only on permanent but also on transitory national cash flow. In the special case when $\beta = 0$, there is no excess sensitivity, and the result is the PIH, i.e.

$$C_t = Z_t^P. \tag{8}$$

It follows from Hall (1978) that the PIH stated as in (8) is the optimal solution for an agent with quadratic utility who sets the subjective discount rate equal to the real interest rate.

While permanent national cash flow follows a martingale under the PIH, this is not true under the ESH. To see this, substitute (7) into (2), lag one period, and then substitute the resulting expression for B_t into (5). After some manipulation, the result is

$$Z_t^P = Z_{t-1}^P - r\beta Z_{t-1}^T + \frac{r}{1+r}\sum_{s=t}^{\infty}\left(\frac{1}{1+r}\right)^{s-t}(E_t - E_{t-1})Z_s. \tag{9}$$

It is evident from (9) that the presence of the term $r\beta Z_{t-1}^T$ destroys the martingale property. However, regardless of whether β is different from zero, the innovation in permanent national cash flow is given by

$$\varepsilon_{zpt} = Z_t^P - E_{t-1}Z_t^P = \frac{r}{1+r}\sum_{s=t}^{\infty}\left(\frac{1}{1+r}\right)^{s-t}(E_t - E_{t-1})Z_s. \tag{10}$$

Because both permanent and transitory national cash flow depend on expectations about future national cash flow, it is difficult to use (7) to determine the restrictions that the fundamental assumption of consumption smoothing behavior places on the excess sensitivity parameter, β. Furthermore, as shown below, transitory national cash flow is by definition proportional to expected *declines* in national cash flow. As a consequence, it can also be difficult to use (7) to determine the exact impact of expected future changes in national cash flow on current consumption. Fortunately, expressing the ESH as a statement about saving, CA_t, allows for a clear interpretation of both the economic and time series properties of the hypothesis.[4] Hence, before investigating the properties of the ESH further, (7) is transformed into a statement about saving.

Intuitively, the fraction of transitory national cash flow that is not consumed in the current period is saved. To see this formally, substitute (6) and (7) into (3), so that

$$CA_t = Z_t^T + Z_t^P - \beta Z_t^T - Z_t^P = (1-\beta)Z_t^T. \tag{11}$$

Transitory national cash flow can be expanded in an informative way. First, substitute (5) into (6) in order to obtain

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[4] For further elaboration on this point, see Campbell (1987).

$$Z_t^T = Z_t - \frac{r}{1+r} \sum_{s=t}^{\infty} \left(\frac{1}{1+r}\right)^{s-t} E_t Z_s. \tag{12}$$

After some manipulation, this yields

$$Z_t^T = - \sum_{s=t+1}^{\infty} \left(\frac{1}{1+r}\right)^{s-t} E_t \Delta Z_s, \tag{13}$$

where $\Delta Z_s = Z_s - Z_{s-1}$. Thus, as claimed above, transitory national cash flow is proportional to expected declines in national cash flow. Substituting (13) into (11), the ESH can be stated as

$$CA_t = (1-\beta)\left[- \sum_{s=t+1}^{\infty} \left(\frac{1}{1+r}\right)^{s-t} E_t \Delta Z_s\right]. \tag{14}$$

Thus, according to the ESH, the current account is equal to the present discounted value of expected declines in national cash flow times a scaling factor that depends on the magnitude of the excess sensitivity parameter, β.

Equation (14) is a generalization of Campbell's (1987) "saving-for-a-rainy-day" equation, and it reveals the restriction that the assumption of consumption smoothing behavior places on the excess sensitivity parameter, β. It follows that consumption smoothing behavior is only consistent with any value of β that is *less* than unity. That is, when $\beta < 1$, (14) implies that when future national cash flow is expected to fall, the current account is in surplus, i.e., the agent saves in order to smooth consumption over time. On the other hand, when the agent expects an increase in future national cash flow, the current account is in deficit, i.e., the agent borrows in order to smooth consumption over time. By contrast, when $\beta = 1$, the current account is closed and can therefore not be used to smooth consumption. When $\beta > 1$, the expectation of, say, a future increase in national cash flow implies a current account surplus. It is evident that saving abroad and consuming relatively less today when future income is expected to *increase,* can never be consistent with at least some degree of consumption smoothing behavior. Hence, in an open con-

sumption smoothing economy, an estimate of β that is equal or larger than unity cannot be justified.

Substituting (14) into (3), the ESH can now be restated as

$$C_t = rB_t + Z_t + (1 - \beta) \sum_{s=t+1}^{\infty} \left(\frac{1}{1+r}\right)^{s-t} E_t \Delta Z_s. \tag{15}$$

Thus, in contrast to (7), (15) shows that consumption can be decomposed into one part that is equal to total *current* income, i.e. total cash flow, $V_t = rB_t + Z_t$, and into one part that is dependent on expectations about *future* income, i.e. the current account deficit. Simply put, present consumption is the sum of total current income and dissaving. Of course, (15) just illustrates the flip side of the saving decision; it explicitly shows how consumption is smoothed by means of the current account whenever the economy is open (i.e., whenever $\beta < 1$). When national cash flow is expected to rise at some point in the future, consumption smoothing behavior implies that present consumption always increases above current total cash flow by the amount that is borrowed abroad. Conversely, when national cash flow is expected to fall, consumption smoothing behavior implies that present consumption always falls below current total cash flow by the amount that is saved abroad.

2.2 Empirical method for evaluating the ESH

Because (14) is conditional on the information set I_t which contains *all* information available to the agent, using it to test the ESH may look impossible. However, this problem can be solved by assuming that the information set used by the econometrician, H_t, is contained in I_t. Taking expectations of (14) conditional on H_t gives

$$
\begin{aligned}
E\left\{CA_t\middle|\, H_t\right\} &= CA_t = E\left\{E\left\{(1-\beta)\left[-\sum_{s=t+1}^{\infty}\left(\frac{1}{1+r}\right)^{s-t}\Delta Z_s\right]\middle|\, I_t\right\}\middle|\, H_t\right\} \\
&= (1-\beta)\left[-\sum_{s=t+1}^{\infty}\left(\frac{1}{1+r}\right)^{s-t}E(\Delta Z_s|H_t)\right],
\end{aligned}
\tag{16}
$$

since CA_t is observable and therefore contained in H_t and $H_t \subseteq I_t$. Thus, in forming expectations about future changes in national cash flow, the agent is likely to use superior information, i.e. information that is not available to the econometrician. But the use of superior information is reflected in the observable current account, and including it when predicting future changes in national cash flow makes it possible for the econometrician to control for information beyond past changes in national cash flow. Hence, an econometric implication of the ESH is that when CA_{t-1} Granger-causes ΔZ_t, there is statistical evidence that the agent has superior information.

The empirical vehicle for evaluating the ESH is the VAR

$$
\begin{bmatrix} \Delta Z_t \\ CA_t \end{bmatrix} = \begin{bmatrix} \Psi_{11} & \Psi_{12} \\ \Psi_{21} & \Psi_{22} \end{bmatrix} \begin{bmatrix} \Delta Z_{t-1} \\ CA_{t-1} \end{bmatrix} + \begin{bmatrix} \varepsilon_{\Delta zt} \\ \varepsilon_{cat} \end{bmatrix},
\tag{17}
$$

where the means of ΔZ_t and CA_t have been removed.[5] Before estimating this VAR, besides verifying that Z_t is I(1), it is standard procedure to verify that the components of the vector (V_t, C_t) are cointegrated such that CA_t is stationary. That is, by theory, V_t and C_t should both be I(1) such that the linear combination

---

[5] For the annual data used in this paper, a one lag VAR is sufficient to capture the time series properties. However, if necessary, the VAR can easily be extended to incorporate several lags.

$(1, -1)(V_t, C_t)' = V_t - C_t = CA_t$ is I(0).[6] In matrix notation, (17) becomes

$$X_t = \Psi X_{t-1} + \varepsilon_t. \tag{18}$$

Following Campbell (1987), there are two equivalent approaches for evaluating the ESH by means of the VAR (17). The first is to impose the ESH (16) on the identity (4), and then test the implied cross-equation restrictions of the VAR. The second is to write the ESH (16) directly in terms of the VAR and then test the resulting cross-equation restrictions.

To show the first approach, use the ESH (16) to write the right-hand side of (4) as

$$
\begin{aligned}
&(1+r)CA_{t-1} - CA_t + \Delta Z_t \\
=\ & \beta \Delta Z_t + (1-\beta) \sum_{s=t}^{\infty} \left(\frac{1}{1+r} \right)^{s-t} [E(\Delta Z_s \,|H_t) - E(\Delta Z_s \,|H_{t-1})].
\end{aligned} \tag{19}
$$

If we define $e_1' = (1, 0)$ and $e_2' = (0, 1)$ such that $e_1' X_t = \Delta Z_t$ and $e_2' X_t = CA_t$, the left-hand side of (19) can be written in VAR notation as

$$
\begin{aligned}
&(1+r)CA_{t-1} - CA_t + \Delta Z_t \\
=\ & (1+r)e_2' X_{t-1} - e_2' X_t + e_1' X_t \\
=\ & ((e_1' - e_2')\Psi + (1+r)e_2')X_{t-1} + (e_1' - e_2')\varepsilon_t.
\end{aligned} \tag{20}
$$

Using the same notation on the right-hand side of (19) yields

$$
\begin{aligned}
&\beta \Delta Z_t + (1-\beta) \sum_{s=t}^{\infty} \left(\frac{1}{1+r} \right)^{s-t} [E(\Delta Z_s \,|H_t) - E(\Delta Z_s H_{t-1})] \\
=\ & \beta e_1' \Psi X_{t-1} + \beta e_1' \varepsilon_t + (1-\beta) \sum_{s=t}^{\infty} \left(\frac{1}{1+r} \right)^{s-t} e_1' \Psi^{s-t} \varepsilon_t.
\end{aligned} \tag{21}
$$

---

[6] Note that (14) implies that CA_t is I(0) since it is a linear combination of expected changes in national cash flow. For additional details of the cointegration properties in this setting, see Campbell (1987).

If the ESH is true, the left-hand side of (20) should be equal to the left-hand side of (21). Consequently, in terms of the VAR, i.e., in terms of the right-hand sides of (20) and (21), the following set of restrictions must be satisfied:

$$(e_1' - e_2')\Psi + (1+r)e_2' = \beta e_1'\Psi \tag{22}$$

$$e_1' - e_2' = \beta e_1' + (1-\beta)\sum_{s=t}^{\infty}\left(\frac{1}{1+r}\right)^{s-t}e_1'\Psi^{s-t}. \tag{23}$$

When $\beta \neq 0$, the restrictions in (22) imply that the change in consumption responds to predictable changes in national cash flow, i.e., consumption responds to lagged changes in national cash flow and lagged current account. They represent the formally testable implications of the ESH and can be evaluated by means of a likelihood-ratio test. As pointed out by Flavin (1993), from the restricted model that underlies the calculation of the likelihood-ratio statistic, an estimate of the excess sensitivity parameter, β, can be obtained. The restrictions in (23) imply that the innovation in the change in consumption depend on both the innovation in the change in national cash flow and the innovation in permanent national cash flow; they are algebraically equivalent to the ones in (22). To see this, note that (22) can be written as

$$(1-\beta)e_1' = ((1-\beta)e_1' - e_2')(I - (1+r)^{-1}\Psi)$$

$$\Leftrightarrow$$

$$(1-\beta)e_1'(I - (1+r)^{-1}\Psi)^{-1} = ((1-\beta)e_1' - e_2'), \tag{24}$$

which is equivalent to (23).[7]

---

[7] In the case of the PIH, the restrictions (22) and (23) are known as the orthogonality and smoothness condition, respectively. That is, if β is forced to zero, the restrictions in (22) guarantee that the change in consumption is independent of either lagged current account or lagged changes in national cash flow. The restrictions in (23) then ensure that the change in consumption is equal to the innovation in permanent national cash flow; if they are satisfied, consumption cannot be "excessively smooth". For an elaboration, see Campbell and Deaton (1989) and Flavin (1993).

To show the second approach for evaluating the ESH, note that the predicted current account, $\widehat{CA}_t$, is calculated from (16) as

$$
\begin{aligned}
\widehat{CA}_t &= (1 - \beta) \left[- \sum_{s=t+1}^{\infty} \left(\frac{1}{1+r} \right)^{s-t} E(\Delta Z_s | H_t) \right] \\
&= -\frac{1 - \beta}{1 + r} e_1' \Psi (I - \frac{1}{1+r} \Psi)^{-1} X_t.
\end{aligned}
\tag{25}
$$

If the ESH is true, the restrictions in (22) hold and can be rearranged to yield

$$
-\frac{1 - \beta}{1 + r} e_1' \Psi (I - \frac{1}{1+r} \Psi)^{-1} = e_2'.
\tag{26}
$$

Substituting this into (25) gives

$$
\widehat{CA}_t = e_2' X_t = CA_t.
\tag{27}
$$

That is, if the restrictions implied by the ESH are satisfied, the predicted and actual current account series are identical. Ghosh (1995) and Sheffrin and Woo (1990), among others, base their evaluation of the PIH on testing the nonlinear restrictions in (26) by means of a Wald test. Although (26) is equivalent to (22), it must be emphasized that nonlinear transformations of restrictions can change the values and power of Wald statistics (see Campbell and Shiller, 1987; and Gregory and Veall, 1985). Section 3 reports Wald test results for both versions of the restrictions.[8]

2.3   Excess sensitivity and international capital mobility

Under the assumption of the PIH, which implicitly includes the assumption of perfect international capital mobility, Ghosh's (1995) *capital mobility hypothesis* posits that the *variances* of the predicted and actual current account series should be equal. In terms of the ESH model, this means forcing the excess sensitivity parameter, β, to zero in (25), and positing that

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[8] The likelihood-ratio statistic is invariant to the transformation and is therefore not reported for the test of the restrictions in (26).

$$var(\widehat{CA_t}) = var(CA_t). \tag{28}$$

If $A = -(1+r)^{-1}e_1'\Psi(I - (1+r)^{-1}\Psi)^{-1}$, then the capital mobility null hypothesis can alternatively be written as

$$Avar(X_t)A' = e_2'var(X_t)e_2, \tag{29}$$

where $var(X_t)$ is the variance-covariance matrix of ΔZ_t and CA_t. The current account is said to be excessively volatile if

$$Avar(X_t)A' < e_2'var(X_t)e_2. \tag{30}$$

The common approach is to calculate the variance or standard deviation *ratio* between the predicted and actual current account series and then test whether it is significantly different from unity. It follows that one case in which the capital mobility hypothesis holds, is when the PIH is true, because then, by (26), $A = e_2'$. Using the PIH framework, Ghosh then argues that if the capital mobility hypothesis fails in a way that makes the actual current account fluctuate more than the predicted current account, as in (30), it indicates that the degree of capital mobility is higher than the degree that is perfect according to the PIH. On the other hand, if the inequality in (30) is reversed so that the actual current account fluctuates less than the predicted current account, the degree of capital mobility is lower than the degree that is perfect according to the PIH. In addition, Ghosh also argues that excessive capital flows can be due to speculation, and that relatively small capital flows can be due to effective barriers to international capital mobility (see also Ghosh and Ostry, 1995).

As Ghosh (1995) notes, the main caveat to his analysis is that if the capital mobility hypothesis is rejected in the PIH framework, its failure can be attributed to *any* of the underlying assumptions, with perfect capital mobility being just one among several others. Differently put, given that the capital mobility hypothesis

fails, assessments about the degree of international capital mobility can in fact be useless because it is entirely possible that the assumption of perfect capital mobility still holds. It is evident that if one sets out to measure the degree of international capital mobility, it would be better to use a hypothesis that, under the null, can allow for any degree of capital mobility, rather than using a hypothesis, such as the PIH, that assumes that the degree of capital mobility is always perfect.

It turns out that the ESH is an alternative to the PIH that allows for any degree of international capital mobility. As such, the ESH can be interpreted as a formalization of Ghosh's argument. In the open economy setting, the excess sensitivity parameter can capture the extent to which the degree of international borrowing and lending, i.e., the degree of international capital mobility, deviates from the degree that is perfect according to the PIH. To clarify, consider the following three cases which can occur when the ESH is true:

If $\beta = 0$, then there is no excess sensitivity, and the result is the PIH. This is the case of optimal consumption smoothing and perfect international capital mobility. For instance, a suddenly expected increase in future national cash flow implies that current consumption is revised upward to *exactly* equal permanent national cash flow by borrowing an amount *exactly* equal to the present discounted value of expected changes in national cash flow. The borrowed amount is the amount needed to smooth consumption *perfectly*, i.e. the amount making expected consumption constant. In this sense, it is natural to consider the PIH a benchmark case.

It should be emphasized that in the PIH model, i.e. when $\beta = 0$, "perfect" international capital mobility is not necessarily synonymous with a capital market with no regulations for international capital flows. In this framework, perfect international capital mobility merely means the specific degree of capital mobility that allows the agent to always smooth consumption perfectly (i.e. optimally) by setting the current account exactly equal to the present discounted value of expected declines in national cash flow. In general, empirical studies using Ghosh's

(1995) capital mobility hypothesis suggest that it is entirely possible to smooth consumption perfectly by means of the current account even in countries where capital markets are subject to mobility constraints.[9]

As a second case, assume that the degree of international capital mobility is *lower* than the degree that is perfect according to the PIH. To see how excess sensitivity is generated, assume that in period t, national cash flow is suddenly expected to rise in the future so that the present discounted value of expected changes in national cash flow becomes positive (and larger than previously expected). As (15) implies, in response to the expected increase, actual consumption is revised upward by means of borrowing abroad. However, due to the relatively lower degree of international capital mobility, it is not possible to borrow and consume as much as would have been possible if the degree of capital mobility had been perfect. Accordingly, *relative* to what would have occurred if the PIH had been true, consumption is *reduced* by the amount that cannot be borrowed, and excess sensitivity is thereby generated in the form of a positive β. Specifically, as shown in (15), the excess sensitivity parameter, β, captures the fraction of the present discounted value of the expected increase in national cash flow that cannot be borrowed and consumed.[10] The lower the degree of international capital mobility, the higher the β, i.e., the higher the fraction that cannot be borrowed. In the borderline case when $\beta = 1$, the current account is closed and present consumption cannot be smoothed and revised upward at all by borrowing abroad. One can use analogous reasoning to conclude that when national cash flow is expected to decrease, consumption smoothing behavior in combination with relatively low capital mobility implies a positive excess sensitivity parameter and less saving relative to what is optimal according to the PIH.

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[9] See for instance Ghosh and Ostry (1995).

[10] That is, according to (15), when future national cash flow is expected to increase, consumption smoothing behavior implies that present consumption increases above total cash flow by $(1 - \beta) \sum_{s=t+1}^{\infty} \left(\frac{1}{1+r}\right)^{s-t} E_t \Delta Z_s$. *Relative* to what would have occurred if the PIH had been true, however, consumption (and borrowing) is reduced by $\beta \sum_{s=t+1}^{\infty} \left(\frac{1}{1+r}\right)^{s-t} E_t \Delta Z_s$.

As a third case, assume that the degree of international capital mobility is *higher* than the degree that is perfect according to the PIH. Assume again that national cash flow is suddenly expected to rise in the future. In response, just as when capital mobility is relatively low, consumption smoothing behavior implies that present consumption increases above total cash flow by the amount that is borrowed. However, since there is more borrowing taking place than the borrowing that is needed for smoothing consumption perfectly, consumption is increased by more *relative* to what would have occurred if the PIH had been true. Thus, excess sensitivity is generated in the form of a negative β. In (15), $-\beta$ measures the fraction of the present discounted value of the expected increase in national cash flow that is borrowed in excess of what would have occurred if the PIH had been true. One can use analogous reasoning to conclude that when national cash flow is expected to decrease, consumption smoothing behavior, in combination with relatively high capital mobility, implies a negative excess sensitivity parameter and more saving relative to what is optimal according to the PIH.

Obviously, the ESH implies a very different economic interpretation of the capital mobility hypothesis (28). From (26) and (27) it follows that (28) holds not only when PIH is true, but also when the ESH is true. However, since the ESH allows for *any* degree of capital mobility that is consistent with consumption smoothing behavior, it no longer takes on the same interpretation as when the PIH is assumed. That is, if the ESH is true, the variance of the predicted current account is equal to the variance of the actual current account *even if the degree of international capital mobility is lower or higher than the degree that is perfect according to the PIH.*

As an example, if the inequality (30) is the outcome under the failure of the PIH, with the true reason being excessive capital flows, then the puzzle of an excessively volatile current account can be explained by the ESH. When the ESH is assumed, the left-hand side of (30) becomes

$$(1 - \beta)^2 Avar(X_t)A'. \tag{31}$$

From the above, we know that when capital mobility is relatively high, $\beta < 0$. Accordingly, expression (31) is always larger than $Avar(X_t)A'$, and the predicted current account series under the ESH fluctuates more than the predicted current account series under the PIH. Thus, if the ESH is true, the left-hand side of (30) is replaced by (31), the inequality becomes an equality, and the puzzle of an excessively current account is solved.

Analogous reasoning can also solve the puzzle of the PIH failing in a way that makes the actual current account too smooth. In such a case, if the failure is due to relatively low capital mobility, β is positive and less than unity and the ESH holds. Consequently, the variance of the predicted current account under the ESH (31) is lower than the variance under the PIH, $Avar(X_t)A'$, and equal to the variance of the actual current account, $e_2'var(X_t)e_2$.

Accordingly, in sum, if one attributes excess sensitivity to the degree to which international borrowing and lending deviates from the degree that is optimal according to the PIH, the ESH suggests an alternative measure of international capital mobility. This measure is based on the sign and magnitude of the excess sensitivity parameter, β. Specifically, β is inversely related to the degree of international capital mobility and has an upper bound equal to unity. The closer β is to unity, the more closed the economy. In the interval $0 < \beta < 1$, the degree of capital mobility is lower than the degree that is perfect according to the PIH. In an optimal consumption smoothing sense, when $\beta = 0$, capital mobility is perfect. Finally, when $\beta < 0$, the

degree of capital mobility is higher than the degree that is perfect according to the PIH.[11]

3. Empirical estimation and results

The annual Swedish national account data used in this section are taken from *International Financial Statistics* and cover the period 1951-99. Specifically, government consumption, G_t, is derived from line 91F; investment, I_t from lines 93E+93I; GNP_t from line 99A; current GDP, Q_t from line 99B; real GDP from line 99BV; and population from line 99Z. Following Ghosh (1995), the current account is calculated as $CA_t = GNP_t - I_t - G_t - C_t$. Thus, it follows from definition (3) above that $GNP_t = rB_t + Q_t$ such that $V_t = GNP_t - G_t - I_t$. All variables are converted into real terms per capita by dividing by the implicit GDP deflator and population.

Before estimation of the VAR (17), it must be verified that V_t and C_t are I(1) and cointegrated such that $V_t - C_t = CA_t$ is I(0). Furthermore, the model also rests on the assumption that Z_t is I(1). To verify non-stationarity of V_t, C_t, and Z_t, augmented Dickey-Fuller (DF) tests are performed (see Dickey and Fuller, 1981). To verify that V_t and C_t are cointegrated, a DF test of a unit root in CA_t is reported. Then, by means of the Johansen (1988) procedure, the DF test for CA_t, V_t, and C_t is supplemented by a cointegration analysis of the linear combination $(\delta_1, \delta_2)(V_t, C_t)'$. In contrast to the augmented DF test, the Johansen (1988) procedure specifies I(0) as a null hypothesis for CA_t, V_t and C_t.

---

[11]The consumption function generated in the interval $0 < \beta < 1$ is similar to the one that Shibata and Shintani (1998) build their model and test of international capital mobility around. In contrast to this paper, they use a slightly different theory for consumption that includes the autarky case $(B_t = 0, C_t = Z_t)$ in which consumption smoothing by means of the current account is impossible. Also in contrast to this paper, they use a different empirical method that leaves out the way of evaluating model fit by comparing actual and predicted current accounts. Although they obtain negative estimates of the excess sensitivity parameter in some cases, they do not consider the possibility that the degree of international capital mobility may be higher than the degree that is perfect according to the PIH.

Table 1
Unit root tests

Variable	DF t-statistic	LM
V_t	-2.93	0.08
C_t	-2.40	1.37
Z_t	-3.45	0.09
CA_t	-2.23^*	0.36
ΔV_t	-7.77^{**}	0.28
ΔC_t	-5.39^{**}	0.98
ΔZ_t	-8.33^{**}	0.21

Notes: LM is the Lagrange multiplier test for residual correlation from lags 1 to 2. "**" and "*" indicate rejection at the 1 and 5 percent level of significance, respectively.

Table 1 reports the results of the augmented DF unit root tests. The validity of the DF method rests heavily on the assumption of uncorrelated residuals, and to verify this, Lagrange multiplier tests for residual autocorrelation are performed. The results are also shown in Table 1. By lag-length selection tests and diagnostic checking, it was concluded that regressing the first difference of each variable V_t, C_t, and Z_t on a constant, a trend, and the lagged level of the variable was enough to avoid residual autocorrelation. In the case of CA_t, the simplest DF regression $\Delta CA_t = \gamma CA_{t-1} + u_t$ appeared to be enough to ensure serially uncorrelated residuals. The null hypothesis of a unit root is not rejected for V_t, C_t, and Z_t, while it is rejected for CA_t. The augmented DF test is also performed for the first difference of V_t, C_t, and Z_t. In each case, the null hypothesis is strongly rejected, which suggests that there is no unit root in any of the differenced series.

The test results from the Johansen (1988) procedure are summarized in Table 2. The first two columns of the table specify the restriction on the vector (δ_1, δ_2). The third column shows the χ^2 statistic from the likelihood-ratio test of the null hypothesis that the specified vector belongs to the cointegrating space. The last column shows this test statistic's p-value. The test result of the main hypothesis, i.e., whether $(\delta_1, \delta_2) = (1, -1)$ belongs to the cointegrating space such that

$(1, -1)(V_t, C_t)' = CA_t$ is I(0), is displayed in the first row of the table: $\chi^2 = 0.80$ and the corresponding p-value 0.372 imply that the null hypothesis cannot be rejected. The last two rows display results from the test of stationarity of V_t and C_t. For each variable, the null hypothesis is strongly rejected. Hence, the results of the unit root and cointegration tests indicate that Z_t is I(1), and that V_t and C_t are both I(1) and cointegrated such that CA_t is I(0).

Table 2

Restricted cointegration tests

H$_0$:			
δ_1	δ_2	$\chi^2(1)$	p-value
1	-1	0.80	0.372
1	0	8.74	0.003
0	1	8.56	0.003

Notes: The null hypothesis is that the vector (δ_1, δ_2) belongs to the cointegrating space such that $(\delta_1, \delta_2)(V_t, C_t)' = \delta_1 V_t + \delta_2 C_t$ is I(0). In the prior VAR estimation, one lag was sufficient to capture the time series properties and to assure no serial correlation among the residuals.

The next step is to estimate the VAR and test the restrictions on the parameters implied by the PIH and the ESH. Lag-length selection tests and diagnostic checking indicated that a one-lag VAR was sufficient to capture the time series properties. The results of the VAR estimation are shown in the top panel of Table 3. The coefficient for CA_{t-1} in the regression for ΔZ_t is negative as expected, but the hypothesis that CA_{t-1} does not Granger cause ΔZ_t cannot be rejected. Thus, statistically, it is not possible to conclude that agents have superior information.

Turning to the bottom panel of the table, recall from (22) that testing the PIH implies forcing β to zero and testing the restrictions

$$
\begin{aligned}
\Psi_{21} &= \Psi_{11} \\
\Psi_{22} &= (1+r) + \Psi_{12}.
\end{aligned}
\tag{32}
$$

On the other hand, if the ESH is true, it follows from (22) that

$$\Psi_{21} = (1-\beta)\Psi_{11}$$

$$\Psi_{22} = (1+r)+(1-\beta)\Psi_{12}. \qquad (33)$$

Following Flavin (1993), both Wald and likelihood-ratio tests for each of the restrictions (32) and (33) are reported, and the results in Table 3 are based on 4 and 14 percent rates of interest.[12] In addition, the Wald statistic of the transformed version of the restrictions (26) for each assumed hypothesis and interest rate is reported.

<div align="center">

Table 3

VAR estimation and tests of restrictions

</div>

VAR estimation					
	ΔZ_t	CA_t			
ΔZ_{t-1}	-0.120	-0.156	t_{NG}	-1.01	
s.e.	(0.150)	(0.156)	p-value	0.319	
CA_{t-1}	-0.107	0.796			
s.e.	(0.106)	(0.111)			

Tests of restrictions					
PIH:	4%	14%	ESH:	4%	14%
$\chi^2_{W1}(2)$	6.55	18.76	$\chi^2_{W1}(1)$	0.24	0.50
p-value	0.038	< 0.001	p-value	0.621	0.478
$\chi^2_{LR}(2)$	5.15	13.22	$\chi^2_{LR}(1)$	0.19	0.35
p-value	0.076	0.001	p-value	0.662	0.557
$\chi^2_{W2}(2)$	8.63	20.67	$\chi^2_{W2}(1)$	0.09	0.13
p-value	0.013	< 0.001	p-value	0.762	0.715
			$\widehat{\beta}$	-0.943	-1.641
			s.e.	0.644	0.958

Notes: In the top panel, column variables are regressed on row variables (standard errors in parentheses). The sample period is 1951-99. t_{NG} is the test statistic for the null hypothesis that CA_{t-1} does not Granger cause ΔZ_t. In the bottom panel, χ^2_{W1} is the Wald statistic for the test of the restrictions implied by the PIH (32) and the ESH (33); χ^2_{LR} is the corresponding likelihood-ratio statistic. χ^2_{W2} is the Wald statistic for the test of the restrictions when they have been transformed according to (26).

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[12]Previous studies of the PIH usually specify r to be somewhere in the interval $0.04 \leq r \leq 0.14$.

Regardless of the interest rate used or the version of restrictions tested, statistical support for the PIH is at best very weak. For the likelihood-ratio test with $r = 0.04$, it is possible to reject the null hypothesis (32) at the 7.6 percent level of significance. For the Wald test with the same interest rate, it is possible to reject at the 3.8 percent level. The Wald test for $r = 0.04$ when the restrictions have been transformed according to (26) (and β set to zero), rejects the null at the 1.3 percent level of significance. When $r = 0.14$, the null is rejected at any conventional level of significance.

On the other hand, when testing the overidentifying restrictions implied by the ESH, the null hypothesis (33) and its equivalent transformation in (26) cannot be rejected at any conventional level of significance. Furthermore, the results for the ESH are robust with regard to the choice of level of interest rate. The table also shows estimates and standard errors of the excess sensitivity parameter, β, generated from the restricted models obtained when calculating the likelihood-ratio test statistics. For each interest rate, the estimated β is negative; it is -0.943 when $r = 0.04$, and -1.641 when $r = 0.14$. This suggests that Sweden's degree of international capital mobility is higher than the degree that is perfect according to the PIH. Statistically, when $r = 0.04$, the estimated β is weakly negative, since it is just 1.46 standard errors from zero. When $r = 0.14$, the estimated β is 1.71 standard errors from zero.

Figures 1 and 2 plot the predicted current account series under the PIH and ESH, respectively, along with the actual current account series; Table 4 shows the estimated standard deviation ratios of the series. Since the standard deviation ratio is a highly nonlinear function of the underlying VAR parameters, caution must be taken when making inference from it (Bekaert et al., 1997; Kilian, 1998). Accordingly, bootstrap simulations are used to construct 95 percent confidence intervals.

The endpoints of each of these intervals are printed within brackets beneath the corresponding standard deviation ratio estimate in Table 4.[13]
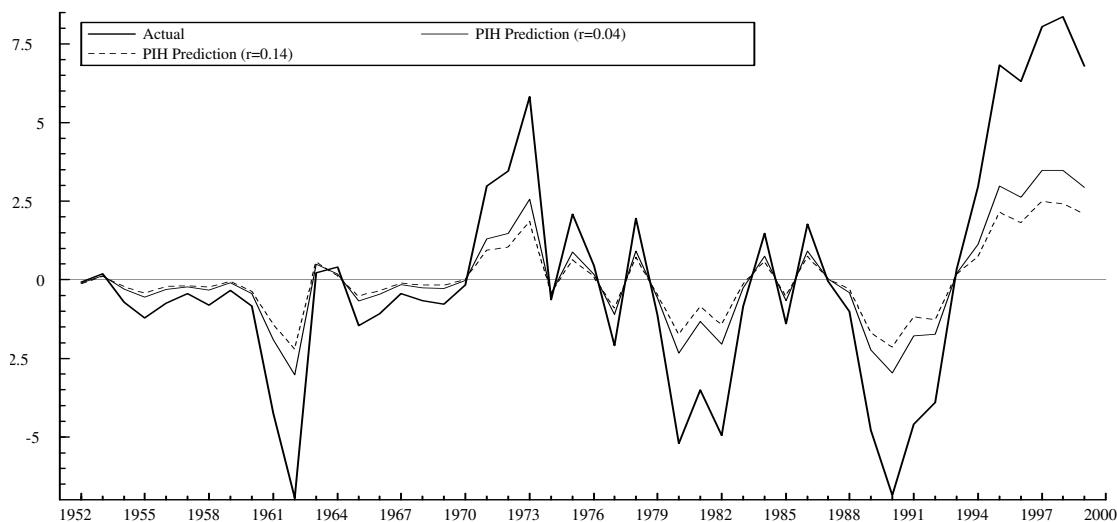

Fig. 1. *PIH predicted and actual current account series, 1952-99.*

Consider first the PIH. As Figure 1 shows, the predicted current account does a good job in capturing the shifts of the actual current account. However, the actual current account fluctuates more than its predicted counterpart. As discussed, this phenomenon is common when the PIH is applied to open economies. Table 4 shows that when $r = 0.04$, the estimated standard deviation ratio is 0.432. The right endpoint of the confidence interval is close to but just under unity, and it is therefore possible to reject the capital mobility hypothesis at the 5 percent level of significance. When $r = 0.14$, the estimated standard deviation ratio is 0.311, and the corresponding confidence interval suggests that the capital mobility hypothesis

[13]Each confidence interval is constructed by means of Kilian's (1998) "bootstrap-after-bootstrap" method. The method is basically divided into two parts. First, based on 1,000 artificial series of ΔZ_t and CA_t, estimation of and correction for bias in the original VAR parameters are made. Then, by use of these bias corrected estimates, another 2,000 artificial series of ΔZ_t and CA_t are produced. For each series of ΔZ_t and CA_t, the VAR parameters are estimated and again corrected for bias. Then, the ratio of standard deviations of the predicted and actual current account series is calculated. Repeating this procedure yields 2000 estimated standard deviation ratios. Finally, a 95 percent confidence interval is constructed from the 2.5th and 97.5th percentile interval endpoints of the resulting empirical distribution. For full details, see Kilian (1998).
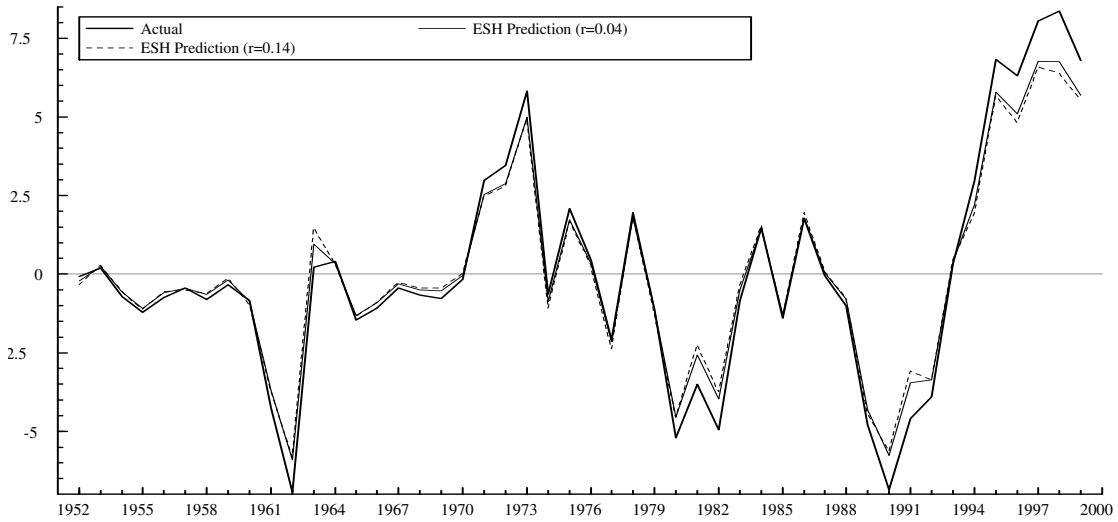
can be strongly rejected.

Fig. 2. *ESH predicted and actual current account series, 1952-99.*

Table 4
Estimated standard deviation ratios and confidence intervals

	PIH		ESH	
	4%	14%	4%	14%
$\sigma(\widehat{CA_t})/\sigma(CA_t)$	0.432	0.311	0.840	0.821
95% c.i.	[0.079 0.989]	[0.060 0.643]	[0.153 1.920]	[0.159 1.699]

In the case of the ESH, Figure 2 shows that, as suggested by the formal tests, the predicted and actual current account series are almost identical. The corresponding estimated standard deviation ratios shown in Table 4 also corroborate the failure to reject the ESH. For both interest rates, the ratios are above 0.8, and unity is easily contained in the corresponding confidence intervals. Hence, for each interest rate, the hypothesis that the ratio of standard deviations is equal to unity, cannot be rejected.

4. Concluding remarks

As emphasized by Flavin (1993), rather than the completely general alternative "any other behaviour other than that predicted by the PIH", the ESH is a *specific* alternative hypothesis to the PIH that, in the open economy setting, implies a mere rescalement of the current account, i.e., a mere rescalement of the degree of international borrowing and lending. Under the assumption of consumption smoothing behavior, the parameter that determines the size of the rescalement, i.e., the excess sensitivity parameter, is restricted to be less than unity. In order to solve the puzzle of an excessively volatile current account, the excess sensitivity parameter should be negative. Conversely, in order to solve the possible puzzle of a too smooth current account, the excess sensitivity parameter should be positive and less than unity.

Using the PIH framework, Ghosh (1995) argues that the puzzle of an excessively volatile current account is due to "too much" capital mobility. But, as he notes, since the PIH implicitly assumes that the degree of international capital mobility is fixed and, in an optimal consumption smoothing sense, perfect, such a conclusion must be interpreted with caution. In general, since several hypotheses are tested jointly, it is not possible to attribute model failure specifically to the assumption of perfect capital mobility. Thus, in the PIH framework, an excessively current account remains a puzzle. However, by showing that the behavioral consumption smoothing properties of the *ESH* framework can be reconciled with Ghosh's argument, a solution to the puzzle of an excessively volatile current account is offered. In contrast to the PIH, the ESH allows for *any* degree of international capital mobility that is consistent with consumption smoothing behavior. It turns out that if the ESH is true when the degree of capital mobility is higher than the degree that is perfect according to the PIH, the excess sensitivity parameter is negative. Hence, according to the ESH, an excessively volatile current account is not a puzzle, but rather what one would expect when the degree of capital mobility is relatively high. Conversely, if the ESH is true when capital mobility is relatively low, the estimated excess sen-

sitivity parameter is positive and less than unity. The excess sensitivity parameter can be used to measure international capital mobility; it is inversely related to the degree of capital mobility and has an upper bound equal to unity.

Annual Swedish data for the period 1951-99 are used to compare the PIH with the ESH. The data illustrate what typically happens when testing the PIH on small open economies: Statistical evidence for the validity of the model is at best very weak, and the current account is excessively volatile. By contrast, when it is assumed that consumption is generated by the ESH, the model cannot be rejected and the actual and predicted current account series are almost identical. Accordingly, empirical estimates of the excess sensitivity parameter are negative, suggesting that Sweden's degree of international capital mobility is higher than the degree that is perfect according the PIH.[14]

Finally, there are some caveats to the analysis. As argued, the existence of excess sensitivity can be reconciled with the view that international capital mobility is not perfect. However, as Deaton (1991) points out, there could be other explanations for excess sensitivity as well, including durability and habit persistence. Thus, detecting excess sensitivity is one thing, attributing it fully to nonperfect capital flows is another, and this is a limitation one must recognize. Still, to the extent that the degree of international capital mobility *does* deviate from the degree that is perfect according the PIH, it should be captured by the excess sensitivity parameter. In this sense, the often observed volatility of international capital markets suggests that a nonperfect degree of capital mobility is a strong candidate for explaining excess sensitivity and the failure of the PIH in industrialized countries. At the same time, it should also be noted that there may also be other alternatives to the PIH that

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[14]It is well-known that, on several occasions during the past decades, the Swedish krona has been subject to devaluation expectations, resulting in massive capital flows. As Ghosh (1995) points out, to the extent that such flows are not absorbed by reserves, they must be reflected in the change in the country's net foreign assets, i.e., in the current account. Thus, Ghosh's additional argument that an excessively volatile current account can be due to speculative flows rather than economic fundamentals, cannot be ruled out.

can improve the model fit.[15] Future research may shed more light on these issues.

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[15]Bergin and Sheffrin (2000) develop a model of the current account that allows for variable interest rates and exchange rates, and then use Campbell's (1987) method to evaluate and test it on Australia, Canada, and the United Kingdom (which all have excessively volatile current accounts). For these countries, the fit of the model is improved when compared to the benchmark PIH model used by previous studies. In another extension of the benchmark PIH model, Ghosh and Ostry (1997) show that precautionary saving motives can be an important determinant of current account surpluses.

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Paper III

The PIH and the standard deviation ratio:
A Monte Carlo study

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Abstract

This paper evaluates the coverage accuracy of small-sample confidence intervals for the functional form of the standard deviation ratio summary statistic that follows from the permanent income hypothesis (PIH). Three methods are considered to construct the confidence intervals: the asymptotic delta method, Runkle's (1987) standard bootstrap method, and Kilian's (1998) bias-corrected bootstrap-after-bootstrap method. Monte Carlo simulations suggest that the asymptotic delta method is unreliable and that researchers should rather use bias-corrected bootstrap confidence intervals when making inference from the standard deviation ratio.

Keywords: Bootstrap; Permanent income hypothesis; Present value model
JEL classification: C15; E21; F32

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1. Introduction

When evaluating the fit of present value models such as the expectations theory for interest rates, the present value model of stock prices and the permanent income hypothesis (PIH), researchers often use a summary statistic called *the standard deviation ratio*. This paper focuses on the small-sample properties of the specific functional form of the standard deviation ratio that follows from the PIH.[1] In the open economy version of the PIH, this statistic has received special attention, as it is also used to test and assess a country's degree of international capital mobility (Ghosh, 1995).

The purpose of this paper is to evaluate the coverage accuracy of small-sample confidence intervals for the standard deviation ratio when they are constructed by means of the asymptotic delta method, Runkle's (1987) standard bootstrap method, and Kilian's (1998) bias corrected bootstrap-after-bootstrap method. The paper is motivated by three important points. First, since the standard deviation ratio is a highly nonlinear function of the parameters of a vector autoregression (VAR), it may not be reliable to use asymptotic theory in small samples when making inference. This is a point that most empirical studies of the open economy version of the PIH have ignored.[2] Second, as argued by Bekaert et al. (1997) and Kilian (1997), adjusting the Monte Carlo methods in order to account for small-sample bias in the VAR coefficients may further improve the coverage accuracy of confidence intervals for statistics such as the standard deviation ratio.[3] Finally, while empirical and Monte Carlo studies of the present value models of the other theories use sample sizes above 100, empirical studies of the open economy PIH often use sample sizes

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[1] The functional forms of the standard deviation ratio that follow from the other theories are very similar.

[2] By contrast, studies of the expectations theory for interest rates and the present value model of stock prices that use Monte Carlo methods to evaluate the small sample properties of the standard deviation ratio are quite common (see for instance Campbell and Shiller, 1989, 1991; Hardouvelis, 1994; and Bekaert et al., 1997).

[3] Recent empirical studies in the present value model VAR framework that use bias-adjusted Monte Carlo methods along the lines suggested by Bekaert et al. (1997) and Kilian (1998), include those by Adler (2001, 2002) and Engsted and Tanggaard (2001, 2002).

ranging from 30 to 50 observations. In this range, little is known about the properties of the standard deviation ratio.

This paper is organized as follows. Section 2 derives the standard deviation ratio, while Section 3 briefly discusses the methods considered to construct the confidence intervals. Section 4 presents the results of a Monte Carlo study. Finally, Section 5 provides some concluding remarks.

2. The standard deviation ratio

As shown by Campbell (1987), the PIH implies the following relation between an income measure, y_{1t}, and a saving measure, y_{2t}:

$$y_{2t} = - \sum_{s=t+1}^{\infty} \delta^{s-t} E_t \Delta y_{1s}, \tag{1}$$

where $\Delta y_{1s} = y_{1s} - y_{1,s-1}$, δ is a known discount factor satisfying $0 < \delta < 1$, and E_t is the mathematical expectation, conditional on an information set, I_t, that generally exceeds the information set available to the econometrician, H_t. It is assumed that current and lagged values of Δy_{1t} and y_{2t} are included in H_t. Furthermore, it is usually assumed that y_{1t} is $I(1)$. Thus, according to (1), y_{2t} is $I(0)$ since it is a linear combination of Δy_{1s}.

Following Campbell (1987) and Campbell and Shiller (1987), the theory as stated by (1) can be evaluated by means of a pth order vector autoregression (VAR (p)) for Δy_{1t} and y_{2t}. The VAR (p) can be written in companion form as the VAR (1)

$$Y_t = \beta_c Y_{t-1} + \varepsilon_t, \tag{2}$$

where Y_t is the $2p \times 1$ companion form vector of regressors, $Y_t = [\Delta y_{1t}, \Delta y_{1,t-1}, ..., \Delta y_{1,t-p+1}, y_{2t}, y_{2,t-1}, ..., y_{2,t-p+1}]$, expressed as deviations from their means; β_c is the companion matrix; and ε_t is a vector of zeros except for the 1st and $p+1$st element which are the Gaussian white noise innovations ε_{1t} of Δy_{1t} and ε_{2t} of y_{2t}, respectively. A first implication for the VAR that follows from (1) is that $y_{2,t-1}$ negatively

Granger-causes Δy_{1t} whenever agents have information beyond that contained in the history of Δy_{1t}. To see the other restrictions that the theory imposes on the VAR, one proceeds in the following way. First, project (1) on H_t:

$$
\begin{aligned}
\widehat{y}_{2t} &= E(y_{2t}|H_t) = E\left\{E\left\{-\sum_{s=t+1}^{\infty}\delta^{s-t}\Delta y_{1s}\,\middle|\,I_t\right\}\,\middle|\,H_t\right\} \\
&= -\sum_{s=t+1}^{\infty}\delta^{s-t}E(\Delta y_{1s}|H_t),
\end{aligned}
\tag{3}
$$

since $H_t \subseteq I_t$. From (2), it follows that

$$
E(\Delta y_{1s}|H_t) = e_1'\beta_c^{s-t}Y_t,
\tag{4}
$$

where e_1 is a column vector with $2p$ elements, where the first element is equal to unity and the others are equal to zero. Substituting (4) into (3) gives

$$
\widehat{y}_{2t} = -\sum_{s=t+1}^{\infty}\delta^{s-t}e_1'\beta_c^{s-t}Y_t = AY_t,
\tag{5}
$$

where $A = -\delta e_1'\beta_c\left(I - \delta\beta_c\right)^{-1}$. The infinite sum in (5) converges as long as the elements of Y_t are stationary. It follows that if the economic theory is true, the time series predicted by the theory, $\widehat{y}_{2t}$, is equal to the actual time series, y_{2t}, since $y_{2t} \in H_t$. In terms of the VAR notation, this implies that

$$
\widehat{y}_{2t} = AY_t = e_2'Y_t = y_{2t},
\tag{6}
$$

where e_2 is a column vector with $2p$ elements, where the $p + 1$st element is equal to unity and the other elements are equal to zero. Thus, if the theory is true, then $A = e_2'$.

The standard deviation ratio is derived from (6). It is defined as

$$
\theta(\beta) = \frac{\sqrt{Avar(Y_t)A'}}{\sqrt{e_2'var(Y_t)e_2}},
\tag{7}
$$

64

where β is a column vector containing the VAR coefficients, and $var(Y_t)$ denotes the variance-covariance matrix of Y_t. Thus, the standard deviation ratio is equal to the ratio of the standard deviation of $\widehat{y}_{2t}$ to the standard deviation of y_{2t}, and as such, it is a highly nonlinear function θ of the VAR parameters β.[4] Researchers are often interested in testing the null hypothesis that $\theta(\beta) = 1$, which follows directly if the PIH is true, i.e., if $A = e_2'$. Furthermore, in the open economy version of the PIH, the null hypothesis that $\theta(\beta) = 1$ is often used to test and assess a country's degree of international capital mobility.[5]

3. The methods considered

3.1 The delta method

The delta method is a common approach to obtain the variance of a statistic that is a nonlinear function of regression parameters (see for instance Lütkepohl, 1993, p. 488). Consider the first order Taylor series expansion of $\theta(\widehat{\beta})$ calculated from a bivariate VAR(p) with T observations and evaluated at β:

$$\theta(\widehat{\beta}) = \theta(\beta) + g(\widehat{\beta} - \beta) + ..., \tag{8}$$

where g is the $4p \times 1$ gradient of θ. As $T \to \infty$, it follows that $\theta(\widehat{\beta})$ is normally distributed with mean $\theta(\beta)$, and an estimator of the asymptotic variance given by $g'var(\widehat{\beta})g$. Accordingly, a 95 percent confidence interval for $\theta(\beta)$ is given by $\theta(\widehat{\beta}) \pm 1.96\sqrt{g'var(\widehat{\beta})g}$.

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[4] I follow Campbell (1987) and Campbell and Shiller (1987) and treat the variance-covariance matrix of Y_t as fixed.

[5] In essence, the test that $A = e_2'$ and the capital mobility null hypothesis that $\theta(\beta) = 1$, amount to a joint test of the PIH and the assumption of perfect international capital mobility (for details, see Ghosh, 1995; and Adler, 2002).

3.2 Bootstrap methods

In contrast to the delta method, bootstrap methods make no distributional assumptions. To obtain confidence intervals based on the bootstrap distribution percentiles, this paper considers two methods: the standard bootstrap method as outlined in Runkle (1987), and Kilian's (1998) bootstrap-after-bootstrap method.

The standard bootstrap is implemented as follows. First, the VAR is estimated in order to obtain $\widehat{\beta}$ and the matrix of fitted residuals $\varepsilon_t^* = (\widehat{\varepsilon}_{1t}, \widehat{\varepsilon}_{2t})$. Then random draws are taken with replacement from ε_t^* and used together with $\widehat{\beta}$ as though they were the population values in order to generate a bootstrap sample, (y_{1t}^*, y_{2t}^*). The VAR is fitted to these simulated data, which gives the bootstrap replications $\widehat{\beta}^*$ that are used to calculate $\theta(\widehat{\beta}^*)$. Repeating this procedure a large number of times gives a distribution of $\theta(\widehat{\beta}^*)$. A 95 percent confidence interval for $\theta(\beta)$ is then constructed by calculating the 2.5th and 97.5th percentile interval endpoints of the distribution of $\theta(\widehat{\beta}^*)$.

Kilian's (1998) bootstrap-after-bootstrap method is an extension of Runkle's (1987) algorithm that corrects for small sample bias in the VAR coefficients. First, the estimated VAR coefficients and the fitted residuals are used to generate a large number of bootstrap replications $\widehat{\beta}^*$ that are used to estimate and correct for bias in $\widehat{\beta}$. The bias-corrected estimates $\widetilde{\beta}$ are then used to generate new bootstrap replications $\widehat{\beta}^*$ that are bias-corrected using the bias estimate from the first step. Repeating this procedure a large number of times gives a large number of bias-corrected bootstrap replications $\widetilde{\beta}^*$. Finally, a 95 percent confidence interval for $\theta(\beta)$ is constructed by calculating the 2.5th and 97.5th percentile interval endpoints of the distribution of $\theta(\widetilde{\beta}^*)$. In addition, the algorithm also includes a procedure for shrinking the bias estimates in case the modulus of the largest root of the companion matrix of the bias-corrected VAR estimates is equal to or greater than unity.[6]

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[6] Typically, this can happen when the true VAR process is highly persistent.

4. The Monte Carlo study

The simulation design is similar to Kilian's (1998), and the population model used throughout is the stationary VAR (1)

$$
Y_t = \begin{pmatrix} \beta_{11} & \beta_{12} \\ \beta_{21} & \beta_{22} \end{pmatrix} Y_{t-1} + \varepsilon_t, \qquad \varepsilon_t \overset{iid}{\sim} N \left(\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} 1 & 0.3 \\ 0.3 & 1 \end{pmatrix} \right), \qquad (9)
$$

where, in accordance with (6), $\beta_{21} = \beta_{11}$ and $\beta_{22} = \delta^{-1} + \beta_{12}$ such that $A = e_2'$, which implies that $\theta(\beta) = 1$ by construction. For simplicity $\delta = 0.96$; $\beta_{11} \in \{-0.2, 0, 0.2\}$ and $\beta_{12} \in \{-0.85, -0.50, -0.15\}$. The sample sizes considered are 30, 50, and 70. For each VAR, denote the modulus of the largest root of the companion matrix as $\mathrm{mod}(\beta_c)$. The parameter settings are deliberately chosen to give a wide spread of the degree of persistence; for any of the nine possible population VARs, $0.19 < \mathrm{mod}(\beta_c) < 0.92$.

For each design point, 500 Monte Carlo trials are generated from the process (9). In each trial, confidence intervals are constructed by means of the asymptotic delta method, the standard bootstrap method, and the bootstrap-after-bootstrap method. For the bootstrap-after-bootstrap, 1000 samples are generated in each trial for the bias estimation. Then, 2000 new samples are generated in order to calculate the 2000 standard deviation ratios used to construct the confidence interval.[7] Accordingly, 2000 samples are also used to construct the confidence interval for the standard bootstrap. Repeating this procedure 500 times for each design point gives 500 confidence intervals for $\theta(\beta)$ for each design point for each of the three methods. The effective coverage rate is then defined as the relative frequency at which the confidence interval covers unity, i.e., the true value of $\theta(\beta)$. Since 95 percent confi-

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[7] The residuals are recentered and rescaled by $\sqrt{(T/(T - 2p))}$ prior to bootstrapping, and the data generating process is initialized by randomly selecting two initial observations for Δy_{1t} and y_{2t} (see Peters and Freedman, 1984; and Stine, 1987). Following Engsted and Tangaard (2001), estimated values of β that are in the nonstationary region before applying the bias adjustment are discarded.

dence intervals are used, an effective coverage of 0.95 would imply perfect coverage accuracy.

Tables A.1 and A.2 in the Appendix display the results of the relative coverage performance of the three methods when the VAR sample size is 30, and the tables also show the corresponding values of $\text{mod}(\beta_c)$ for each design point. Tables A.3. and A.4. summarize the results when the sample sizes are 50 and 70, respectively. In summary, the bootstrap-after-bootstrap interval provides the most accurate coverage, closely followed by the standard bootstrap. The bootstrap-after-bootstrap interval provides better coverage accuracy than the standard bootstrap in 16 of the 27 design points, and it provides better coverage accuracy than the asymptotic interval in all but one case. In general, intervals constructed by means of the asymptotic delta method have the largest coverage error. They provide relatively good coverage accuracy when $\beta_{12} = -0.85$, but they always imply a coverage rate that undercovers the target coverage rate of 0.95.

The most striking feature about the results is the quite dramatic drop of the coverage rate of the asymptotic interval, and to some extent also of the standard bootstrap interval when the persistence of the VAR process increases. Consider the case when the sample size is 30 and $\beta_{11} = -0.2$. Then, from Table A.1, when β_{12} increases from -0.85 to -0.50, $\text{mod}(\beta_c)$ increases from 0.46 to 0.66. This moderate increase of the persistence implies a relatively large drop in the coverage rate of the asymptotic interval as it falls from 0.912 to 0.848. From Table A.2, when β_{12} is further increased to -0.15, $\text{mod}(\beta_c) = 0.92$, and the coverage rate of the asymptotic interval drops off to 0.456. The standard bootstrap interval performs a little better than the asymptotic interval. Given that $\beta_{11} = -0.2$, when β_{12} increases from -0.85 to -0.50 to -0.15, the coverage rate of the standard bootstrap interval drops from 0.954 to 0.920 to 0.692. The bootstrap-after-bootstrap interval is more robust than both the standard bootstrap and the asymptotic intervals. When β_{12} increases from -0.85 to -0.50, the coverage rate of the bootstrap-after-bootstrap interval actually

improves slightly from 0.960 to 0.942. When β_{12} is further increased to -0.15, the coverage rate drops to 0.848. Even though this coverage rate is far better than the corresponding rates for the standard bootstrap and asymptotic intervals, it suggests that when the sample size is small and the degree of persistence is very high, all methods have a problem with the confidence intervals.

Given that $\beta_{12} = -0.50$, when the sample size is increased to 50 and 70 the relative differences between the three methods are small. However, given that $\beta_{12} = -0.15$, even when the sample size is increased to 70 there is still a significant difference between the methods. As shown in the right panel of Table A.4, the coverage rate of the bootstrap-after-bootstrap interval ranges from 0.896 to 0.950, while the coverage rates of the standard bootstrap and the asymptotic intervals range from 0.800 to 0.916 and from 0.752 to 0.910, respectively.

Tables A.1 to A.4 also display the average lengths of the confidence intervals for each design point. In general, the bootstrap-after-bootstrap method generates the widest intervals and the asymptotic delta method the shortest. The only exception is for design points that have $\beta_{12} = -0.15$ and sample sizes of 50 and 70 (i.e. the design points corresponding to the entries in the right panels of Tables A.3 and A.4). For these points, the average length of the standard bootstrap interval is slightly shorter than the average length of the asymptotic interval. When the degree of persistence is relatively low, the standard bootstrap may be preferred over the bootstrap-after-bootstrap, as the shorter interval of the standard bootstrap comes with almost no cost in terms of loss of coverage accuracy. However, as the degree of persistence increases, the relatively wider confidence interval of the bootstrap-after-bootstrap should be of no concern because its corresponding coverage accuracy clearly outperforms the standard bootstrap.

As noted above, it is evident from the tables that the largest difference in relative coverage performance is when the degree persistence is very high, i.e., when $\mathrm{mod}(\beta_c) = 0.92$. In this case, even when the sample size is 70, none of the methods

can produce a coverage rate that is above 0.90. From Table A.4, the relative performance of the bootstrap-after-bootstrap interval is far better than the others, yet it only achieves a coverage rate of 0.896. For this degree of persistence, when do the interval coverage rates of the different methods converge? To investigate this, additional simulations are conducted for the design point that has $\text{mod}(\beta_c) = 0.92$. The results are displayed in Table A.5. Again, there is simulation evidence that shows the relative robustness of the coverage accuracy of the bootstrap-after-bootstrap interval. When the sample size is 100, there are still significant differences among the relative performances of the three methods. The coverage rate of the bootstrap-after-bootstrap interval is 0.926, while the coverage rates of the standard bootstrap and the asymptotic intervals are only 0.856 and 0.832, respectively. It appears that for the asymptotic interval to produce satisfactory results, the sample size should be above 150 observations.

5. Concluding remarks

The Monte Carlo evidence in this paper suggests that in small samples, small changes in parameters can drastically worsen the coverage accuracy of confidence intervals constructed by means of the asymptotic delta method, even when the degree of persistence of the VAR process is moderate. For VAR processes with a low to moderate degree of persistence, both the standard bootstrap interval and the bootstrap-after-bootstrap interval appear to be quite accurate. As the degree of persistence increases, the coverage accuracy of the bootstrap-after-bootstrap interval clearly outperforms the standard bootstrap method. However, in small samples, for highly persistent processes, all methods have a problem. Additional simulations show that in order to achieve satisfactory coverage accuracy for the bootstrap-after-bootstrap interval when the degree of persistence is very high, the sample size should be larger than 70 observations.

Appendix

Table A.1

Effective coverage rates and average lengths of 95 percent confidence intervals

	$\beta_{12} = -0.85$				$\beta_{12} = -0.50$			
β_{11}	*ASY*	*SB*	*BAB*	$\mathrm{mod}(\beta_c)$	*ASY*	*SB*	*BAB*	$\mathrm{mod}(\beta_c)$
-0.2	0.912	0.954	0.960	0.46	0.848	0.920	0.942	0.66
	(1.36)	(1.55)	(1.95)		(1.43)	(1.55)	(2.19)	
0	0.928	0.948	0.954	0.19	0.908	0.948	0.958	0.54
	(1.27)	(1.44)	(1.75)		(1.48)	(1.60)	(2.15)	
0.2	0.930	0.948	0.938	0.46	0.918	0.948	0.960	0.46
	(1.10)	(1.25)	(1.43)		(1.33)	(1.44)	(1.82)	

Notes: Average lengths in parentheses. The results are based on 500 confidence intervals and a VAR sample size of 30. *ASY*=Asymptotic delta method, *SB*=Standard bootstrap, *BAB*=Bootstrap-after-bootstrap.

Table A.2

Effective coverage rates and average lengths of 95 percent confidence intervals

	$\beta_{12} = -0.15$			
β_{11}	*ASY*	*SB*	*BAB*	$\mathrm{mod}(\beta_c)$
-0.2	0.456	0.692	0.848	0.92
	(1.14)	(1.16)	(1.60)	
0	0.692	0.824	0.926	0.89
	(1.38)	(1.37)	(1.96)	
0.2	0.814	0.890	0.946	0.85
	(1.36)	(1.38)	(1.88)	

Notes: Average lengths in parentheses. The results are based on 500 confidence intervals and a VAR sample size of 30. *ASY*=Asymptotic delta method, *SB*=Standard bootstrap, *BAB*=Bootstrap-after-bootstrap.

Table A.3
Effective coverage rates and average lengths of 95 percent confidence intervals

| β_{11} | $\beta_{12} = -0.85$ | | | $\beta_{12} = -0.50$ | | | $\beta_{12} = -0.15$ | | |
	ASY	SB	BAB	ASY	SB	BAB	ASY	SB	BAB
-0.2	0.936	0.954	0.960	0.908	0.954	0.950	0.672	0.740	0.862
	(1.08)	(1.18)	(1.34)	(1.40)	(1.43)	(1.87)	(1.22)	(1.17)	(1.66)
0	0.942	0.958	0.950	0.924	0.946	0.950	0.794	0.860	0.934
	(0.95)	(1.03)	(1.14)	(1.29)	(1.34)	(1.64)	(1.38)	(1.33)	(1.84)
0.2	0.924	0.936	0.940	0.930	0.930	0.936	0.884	0.900	0.962
	(0.81)	(0.87)	(0.93)	(1.03)	(1.09)	(1.24)	(1.37)	(1.36)	(1.83)

Notes: Average lengths in parentheses. The results are based on 500 confidence intervals and a VAR sample size of 50. ASY=Asymptotic delta method, SB=Standard bootstrap, BAB=Bootstrap-after-bootstrap.

Table A.4
Effective coverage rates and average lengths of 95 percent confidence intervals

| β_{11} | $\beta_{12} = -0.85$ | | | $\beta_{12} = -0.50$ | | | $\beta_{12} = -0.15$ | | |
	ASY	SB	BAB	ASY	SB	BAB	ASY	SB	BAB
-0.2	0.908	0.944	0.940	0.916	0.958	0.966	0.752	0.800	0.896
	(0.85)	(0.91)	(0.98)	(1.19)	(1.21)	(1.44)	(1.27)	(1.20)	(1.68)
0	0.918	0.950	0.946	0.922	0.924	0.934	0.856	0.886	0.928
	(0.79)	(0.83)	(0.89)	(1.06)	(1.10)	(1.25)	(1.39)	(1.32)	(1.79)
0.2	0.926	0.924	0.920	0.938	0.954	0.956	0.910	0.916	0.950
	(0.67)	(0.71)	(0.74)	(0.88)	(0.92)	(1.00)	(1.31)	(1.27)	(1.64)

Notes: Average lengths in parentheses. The results are based on 500 confidence intervals and a VAR sample size of 70. ASY=Asymptotic delta method, SB=Standard bootstrap, BAB=Bootstrap-after-bootstrap.

Table A.5
Effective coverage rates and average lengths of 95 percent confidence intervals

VAR sample size	ASY	SB	BAB
100	0.832	0.856	0.926
	(1.27)	(1.19)	(1.62)
150	0.914	0.932	0.950
	(1.24)	(1.17)	(1.50)
200	0.930	0.938	0.956
	(1.09)	(1.04)	(1.26)

Notes: Average lengths in parentheses. The results are based on 500 confidence intervals, with $\beta_{11} = -0.2$ and $\beta_{12} = -0.15$ such that $\text{mod}(\beta_c) = 0.92$. ASY=Asymptotic delta method, SB=Standard bootstrap, BAB=Bootstrap-after-bootstrap.

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Paper IV

Has Sweden's government budget policy been too discretionary? Evidence from a generalization of the tax smoothing hypothesis

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Abstract

Barro's (1979) tax smoothing hypothesis (TSH) assumes that the government is always subject to an "optimal" degree of discretion in budget policy, i.e., optimal in the sense that the welfare costs from taxation are minimized. This paper proposes a generalization of the TSH that relaxes this crucial assumption. Postwar evidence for Sweden indicates that in contrast to the TSH, the generalized model provides close to a perfect fit: Tax smoothing behavior in combination with *more* discretion in budget policy relative to what is optimal, can explain all shifts in the central government's budget balance, including the dramatic shifts during the period 1970-96.

Keywords: Tax smoothing; Discretion; Budget policy; Budget deficits
JEL classification: H21, H61

1. Introduction

As noted by Jonung (2000, p. 3), "The stabilisation policy record of Sweden during the period 1970-1995 is unique. No other OECD-country experimented with as many policy switches or policy reversals as Sweden or did it so drastically." During this period, there were no restrictions on the budget policy, and the Swedish budget balance displayed the largest volatility of all the OECD countries. There were several examples of extreme "tighten and loosen the belt" policies. For instance, the late 1970s was characterized by drastically growing budget deficits as a result of an expansionary fiscal policy aimed at "bridging over" the economic downturn that followed from the first oil price shock. By contrast, in the late 1980s, following a period of high growth and fiscal consolidation, the government ran large budget surpluses. In fact, each year during the period 1987-89, the government had one of the largest budget surpluses in the OECD. But then, once again, the situation changed drastically. In the early 1990s, a severe economic crisis and a major tax reform resulted in growing budget deficits, and in 1993, the central government displayed its largest budget deficit ever. In just four years, Sweden went from having the OECD's strongest government finances to having the weakest. The Swedish government then adopted one of the strictest programs of fiscal consolidation in the OECD. Just a few years later, the budget was back into a surplus where it remained for the rest of the 1990s.

The purpose of this paper is to propose a generalization of Barro's (1979) tax smoothing hypothesis (TSH) and then empirically test whether it can explain the shifts in the Swedish central government's budget balance during recent decades. According to the TSH, it can be optimal for a government to run budget surpluses as well as deficits as long as they are justified by future expectations of changes in government expenditure. In its basic form, the TSH implies that when a government expects a future increase in its expenditure, it increases the tax rate today and runs a budget surplus. Conversely, when the government expects a future decrease in

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expenditure, it lowers the tax rate today and runs a budget deficit. The rationale for this behavior is that the government wishes to smooth the tax rate over time in order to minimize the implied distortionary welfare costs from taxation. As a consequence, when the TSH is true, the expected tax rate is constant over time, or, put in more formal terms, the tax rate follows a martingale.

Huang and Lin (1993) and Ghosh (1995) draw on research done by Campbell (1987) and Campbell and Shiller (1987) in order to explore a useful property of the TSH: By means of a vector autoregression (VAR) for government expenditure and the budget surplus, it is possible to calculate a predicted time path of the budget surplus that, given the validity of the TSH, is optimal for the government to pursue. The optimal budget surplus time series can then be compared to the actual budget surplus time series in order to visually evaluate the fit and the economic significance of the model. If the model is true, the two series should be identical. In addition, the theoretical properties of the TSH translate into cross-equation restrictions on the VAR, and standard statistical testing is therefore easily implemented to formally test the validity of the hypothesis.

Empirical evidence for the TSH is mixed. Huang and Lin (1993) apply a log-linear version of the model to the United States for the period 1929-88. For the full sample period, the TSH is rejected, but it is not rejected for the period 1947-88.[1] Ghosh (1995) applies the model to Canada and the United States for the periods 1962-88 and 1961-88 respectively, and the TSH cannot be rejected for either country. Olekalns (1997) rejects the TSH when applied to Australian data for the period 1964/65 to 1994/95. The TSH has also been applied to a couple of developing countries in Asia (see Cashin et al., 1998, 1999), and here also, evidence is mixed.[2]

A striking feature of the TSH is that it implicitly assumes that the government's

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[1] According to Huang and Lin, the rejection for the full sample period is due to sharp differences in the statistical properties of the data rather than the invalidity of the hypothesis itself.
[2] The same goes for several studies on U.S. data, including Barro (1981) and Sahasakul (1986), that, in contrast to the VAR approach, test the TSH by evaluating the martingale property of the tax rate.

budget policy is always subject to an "optimal" degree of discretion, i.e., optimal in the sense that the welfare costs from taxation are minimized. But given the nature of the budget-making process, it is difficult to imagine that this assumption generally holds. It is often argued that not only expectations about future government expenditure, but also politico-institutional factors such as budgetary institutions and budget laws, are crucial to understanding budget deficits and fiscal policy (Alesina and Perotti, 1995). For instance, Olekalns (1997) finds that the actual Australian budget surplus fluctuates more than its optimal counterpart and argues that "fiscal policy has been too volatile to be consistent with optimal tax smoothing." However, in general, because the model implies that several hypotheses are tested jointly, it is not possible to attribute model failure specifically to the deviation from the optimal degree of discretion without having relaxed this assumption in the first place.

To overcome this shortcoming, this paper proposes a generalization of the TSH that allows for degrees of discretion in budget policy that can differ from the degree that is optimal according to the TSH. As such, the proposed *generalized* tax smoothing hypothesis (GTSH) should be regarded as a specific alternative to the TSH.[3] It is assumed that a specific degree of discretion in budget policy translates into a specific constraint on the government's borrowing and lending capabilities. The constraint can either be stricter or softer than the constraint that corresponds to the degree of borrowing and lending that is optimal according to the TSH. A government that is restricted to sticking very closely to a balanced budget rule does not have much discretion in budget policy, i.e., it is restricted in its borrowing and lending. By contrast, a government with much discretion in budget policy can usually borrow and lend with few restrictions, and can therefore deviate from the balanced budget rule to any extent that it wishes. However, maintaining the assumption of at least some degree of tax smoothing behavior implies that for a government with much discretion, there will be no bias toward letting a budget deficit grow forever.

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[3] The inspiration for the GTSH is drawn from Flavin (1993) who applies the analogous generalization to the theory of individual consumption smoothing behavior.

More specifically, regardless of the degree of discretion, under the assumption of tax smoothing behavior, budget policy implies a symmetrical view of the budget balance. For instance, just as a relatively discretionary budget policy implies a large budget deficit when government expenditure is expected to fall, it also implies a "tighten the belt" policy and a large budget surplus when expenditure is expected to rise.

This paper is organized as follows. Section 2 derives some basic results of the TSH and then shows how the model can be generalized in a way that allows for degrees of discretion in budget policy that can differ from the degree that is optimal according to the TSH, i.e., the GTSH. In Section 3, the empirical method for evaluating the model is outlined. Section 4 tests the TSH and the GTSH on Swedish central government data for the periods 1952-99 and 1970-96. Sweden appears to be an ideal candidate for comparing the TSH to the GTSH. A strong Keynesian tradition has resulted in many available policy instruments when implementing fiscal policy, especially before the budget law that took effect in 1997. Furthermore, so far there appears to be no study that has tested an economic hypothesis empirically in order to try to explain the shifts in Sweden's budget balance during the past decades.[4]

The empirical results presented in Section 4 indicate that for the full period 1952-99, it is not possible to statistically distinguish the GTSH from the TSH. However, for the subperiod 1970-96, the TSH is rejected while the GTSH is not. Visually, for both sample periods, the model when the GTSH is assumed provides close to a perfect fit as the predicted and actual path of the budget surpluses almost always coincide. Section 5 provides some concluding remarks.

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[4] Jonung (1999, 2000) tries to explain Swedish stabilization policy during the period 1970-95 by viewing it as the result of a learning process among the politicians. He also recognizes that the theory of tax smoothing can be used to explain the budget policy record of Sweden during the same period, but he undertakes no formal theoretical or empirical analysis to verify it. Hansson and Hansson (2001) use Bohn's (1998) framework and annual data for the period 1885-1996 in order to test whether the debt-to-GDP ratio in Sweden and five other countries is stationary. For Sweden, it is only when they allow for a structural break in 1973 that the debt-to-GDP ratio is found to be stationary. They argue that their results provide "modest evidence in favor of the tax-smoothing hypothesis".

2. A generalization of the tax smoothing hypothesis

Consider first the basic version of the TSH as derived by Ghosh (1995). The government faces the dynamic budget constraint

$$D_{t+1} = (1+r)D_t + G_t - \tau_t Y_t, \tag{1}$$

where D_t is the stock of real government debt; G_t is real government expenditure; τ_t is the average tax rate; Y_t is real output; and r is the fixed real interest rate. If output grows at a fixed rate equal to n, the dynamic budget constraint can be expressed as

$$(1+n)d_{t+1} = (1+r)d_t + g_t - \tau_t, \tag{2}$$

where each lowercase letter denotes the ratio of the corresponding uppercase letter to output. In the model that follows, the ratio of government expenditure to output, g_t, is assumed to be exogenously given. For simplicity, g_t and d_t are hereafter referred to as government expenditure and debt, respectively. In a stochastic setting, the intertemporal budget constraint states that if a transversality condition on debt is imposed, the sum of the present discounted value of expected government expenditure and initial debt must equal the present discounted value of expected tax rates. That is, solving (2) forward, taking expectations and imposing the transversality condition

$$\lim_{T \to \infty} \left(\frac{1}{1+R}\right)^{T+1} E_t d_{t+T+1} = 0, \tag{3}$$

gives

$$\sum_{s=t}^{\infty} \left(\frac{1}{1+R}\right)^{s-t} E_t g_s + (1+r)d_t = \sum_{s=t}^{\infty} \left(\frac{1}{1+R}\right)^{s-t} E_t \tau_s, \tag{4}$$

where $R = (r-n)/(1+n)$ is the effective net interest rate faced by the government and $E_t = E(\cdot|I_t)$ is the expectations operator, conditional on the government's information set at time t, I_t.

The levying of taxes is assumed to impose distortionary costs such as collection costs and deadweight losses incurred when individuals substitute away from market work. Assuming that these costs are proportional to the square of the tax rate, the government's objective function is

$$V = -(1/2) \sum_{s=t}^{\infty} \beta^{s-t} E_t \tau_s^2 \qquad 0 < \beta < 1, \tag{5}$$

where β is the government's subjective discount rate. The problem is then to maximize (5) subject to (2) and (3). Assuming that $\beta = 1/(1 + R)$, the Euler equation implies that for any $s > t$,

$$E_t \tau_s = \tau_t, \tag{6}$$

i.e., the tax rate follows a martingale, or stated less formally, a random walk. This is a first basic implication of the TSH, which has been tested in several empirical studies including those by Barro (1981) and Sahasakul (1986).

Although (6) neatly captures the notion of tax smoothing, there are several reasons for going beyond it (Campbell, 1987). First of all, the random walk of the tax rate can be the result of a political process that is unrelated to the tax smoothing objective. That is, it is possible that the tax rate follows a random walk yet does not satisfy the TSH. Another reason is that it is difficult to assess the economic significance of a statistical rejection of (6). A third reason is that there are useful time series properties that are not explored when focusing solely on (6). To overcome these shortcomings, Huang and Lin (1993) and Ghosh (1995), among others, apply Campbell's (1987) and Campbell and Shiller's (1987) VAR approach in order to explore and test *all* time series implications of the TSH. In short, the approach is to formulate the TSH as a statement about the *budget surplus,* as it takes into account the full structure of the model, and then use a VAR for government expenditure and the budget surplus to evaluate the implied restrictions. The same approach is used in this paper, and it allows us to assess both the statistical and economic

significance of the model.

Using (6) in (4), the TSH can be written as

$$\tau_t = (r-n)d_t + \frac{R}{1+R} \sum_{s=t}^{\infty} \left(\frac{1}{1+R}\right)^{s-t} E_t g_s = g_t^P. \tag{7}$$

According to (7), the only martingale that satisfies the TSH is the martingale that sets the tax rate exactly equal to the annuity value of the sum of government debt and the present discounted value of expected government expenditure. Thus, the right hand side of (7) is the constant flow of expenditure that is expected to sustain for the remainder of the government's time horizon, i.e., it is the *permanent* government expenditure, g_t^P. Optimal budget policy would then imply to always equal the tax rate to permanent government expenditure.

Define the budget surplus as $sur_t = (1+n)(d_t - d_{t+1})$. The dynamic budget constraint (2) can then be rearranged such that

$$sur_t = \tau_t - (g_t + (r-n)d_t). \tag{8}$$

After substituting (7) into the right hand side of (8), the TSH can be stated as

$$
\begin{aligned}
sur_t &= \tau_t - (g_t + (r-n)d_t) \\
&= (r-n)d_t + \frac{R}{1+R} \sum_{s=t}^{\infty} \left(\frac{1}{1+R}\right)^{s-t} E_t g_s - (g_t + (r-n)d_t) \\
&= \frac{R}{1+R} \sum_{s=t}^{\infty} \left(\frac{1}{1+R}\right)^{s-t} E_t g_s - g_t \\
&= \sum_{s=t+1}^{\infty} \left(\frac{1}{1+R}\right)^{s-t} E_t \Delta g_s.
\end{aligned}
\tag{9}
$$

Equation (9) states that when the TSH is true, optimal budget policy implies that the budget surplus is always set to equal the present discounted value of expected changes in government expenditure. Whenever expenditure is expected to increase, the government runs a budget surplus, i.e., it saves for "a rainy day" (Campbell,

1987). Conversely, when expenditure is expected to fall, the government runs a budget deficit.

Using (9), the TSH can now be restated as

$$\tau_t = g_t^{TOT} + \sum_{s=t+1}^{\infty} \left(\frac{1}{1+R}\right)^{s-t} E_t \Delta g_s, \tag{10}$$

where g_t^{TOT} is total government expenditure, i.e., the sum of current expenditure, g_t, and the effective interest payment on government debt, $(r-n)d_t$. In contrast to (6) and (7), (10) explicitly shows that whenever government expenditure is expected to rise, the tax rate increases above total government expenditure by the amount the government lends, i.e., by the amount of the budget surplus. Conversely, when government expenditure is expected to fall, the tax rate falls below total government expenditure by the amount that is borrowed, i.e., by the amount of the budget deficit.

Note that the budget surplus is always set *exactly* equal to the present discounted value of expected changes in government expenditure. As (6) shows, this is the amount needed to smooth the tax rate *perfectly* and, accordingly, it is the amount that minimizes the distortionary welfare costs from taxation. However, a more general and realistic tax setting scheme would be one that reflects the underlying wish of the government to smooth the tax rate *and* at the same time also recognizes that, given the nature of the budget making process, the government may be subject to a degree of discretion in budget policy that differs from the degree that is optimal according to the TSH. It seems natural to assume that a specific degree of discretion in budget policy translates into a specific constraint on the government's borrowing and lending capabilities. The constraint can either be stricter or softer than the constraint that corresponds to the TSH optimal degree of borrowing and lending.[5]

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[5] As Alesina and Perotti (1995) point out, when considering whether to limit a government's degree of discretion in budget policy, one usually considers regulations that limit the ability to run budget deficits. However, note that for a tax smoothing government, the optimal constraint (rule) is symmetric in the sense that the inclination to borrow (and run a budget deficit) when an expected change in expenditure is negative is as large as the inclination to lend (and run a budget

For instance, a government that has little discretion in budget policy cannot borrow or lend the full amount that is needed in order to smooth the tax rate perfectly. Put differently, a government that must stick close to a balanced budget rule is subject to a relatively strict borrowing and lending constraint. Then, an expected future decrease in government expenditure, say, cannot be accommodated by the full cut in the tax rate and, accordingly, by the full budget deficit that is necessary in order to minimize the welfare costs from taxation. Since the government wishes to smooth taxes, the tax rate will however be cut and the budget will show a deficit, but, due to the "partial" balanced budget rule, not by the full amount that is needed to smooth the tax rate perfectly. On the other hand, a government that has *more* discretion in budget policy relative to what is optimal according to the TSH, is subject to a relatively soft constraint on borrowing and lending. A relatively soft constraint can be a reflection of politico-institutional factors, such as weak budgetary institutions and weak budget laws, and it can open up for influence from interest groups with objectives other than optimal tax smoothing. As a result, given, say, an expected future decrease in government expenditure, a government with relatively much discretion can cut the tax rate by *more* and generate a larger budget deficit than a government that smooths the tax rate perfectly. Unfortunately, although both governments in this example have an underlying wish to smooth the tax rate, the TSH can capture neither the government that sticks close to a balanced budget rule, nor the government with more discretion in budget policy relative to what is optimal. The TSH can only capture the specific intermediate case of a government with a budget policy that is subject to an optimal degree of discretion, i.e., optimal in the sense that it minimizes the distortionary welfare costs from taxation.

In order to capture a tax smoothing government with a degree of discretion

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surplus) when the expected change is positive. Maintaining the assumption of tax smoothing behavior, it is therefore natural to consider any other constraint as also being symmetric, i.e., as applying to *both* borrowing (budget deficits) and lending (budget surpluses) capabilities. Formally, the TSH's symmetric view of the budget balance follows from the transversality condition (3) that rules out overborrowing as well as oversaving.

that can differ from the degree that is optimal according to the TSH, one needs a *specific* alternative to the TSH rather than the completely general alternative "any other behavior than that predicted by the TSH" (cf. Flavin, 1993). The alternative hypothesis should be a generalization along the particular dimension that allows for degrees of borrowing and lending that can differ from the degree that is optimal according to the TSH. To construct such an alternative, consider a linear combination of the balanced budget rule (no degree of discretion) in which $\tau_t = g_t^{TOT}$, and the optimal rule (the optimal degree of discretion) in which $\tau_t = g_t^P$. The *generalized* tax smoothing hypothesis (GTSH) can then be written as

$$\tau_t = (1 - \lambda)g_t^P + \lambda g_t^{TOT}, \tag{11}$$

where λ is an estimable parameter. Obviously, the GTSH incorporates both the balanced budget rule ($\lambda = 1$) and the optimal degree of discretion ($\lambda = 0$) as special cases. However, it should be emphasized that, under the assumption of tax smoothing behavior, one should not view both the optimal degree of discretion and the balanced budget rule as two polar cases. Intuitively, the balanced budget rule should be a polar case as it implies that there is no borrowing or lending whatsoever in response to expected future changes in government expenditure. By contrast, because the optimal degree of discretion just happens to reflect a situation where the amounts of borrowing and lending are just large enough to minimize the welfare costs from taxation, there is nothing that says that it should be a polar case. To formally verify these claims, use the definitions of permanent and total government expenditure and rewrite the GTSH (11) as

$$\tau_t = g_t^{TOT} + (1 - \lambda) \sum_{s=t+1}^{\infty} \left(\frac{1}{1 + R} \right)^{s-t} E_t \Delta g_s. \tag{12}$$

Thus, when the GTSH is true, the budget surplus is given by

$$sur_t = (1 - \lambda) \sum_{s=t+1}^{\infty} \left(\frac{1}{1+R} \right)^{s-t} E_t \Delta g_s. \tag{13}$$

It follows directly from (13) that the assumption of tax smoothing behavior is *only* consistent with *any* value of λ that is less than unity. That is, whenever $\lambda < 1$, an expected increase in expenditure implies that the government increases the tax rate and runs a budget surplus; and an expected decrease in expenditure implies that the government cuts the tax rate and runs a budget deficit. By contrast, when $\lambda = 1$, there is no tax smoothing and the budget is always in balance. When $\lambda > 1$, an expected decrease in government expenditure, for example, implies a budget surplus. It is evident that *increasing* the tax rate and saving in response to an expected future expenditure *decrease* can never be consistent with at least some degree of tax smoothing behavior. Hence, under the assumption of tax smoothing behavior, an estimate of λ that is equal or larger than unity cannot be justified.

Equations (12) and (13) explicitly formalize the idea that a tax smoothing government may be subject to a degree of discretion in budget policy that differs from the degree that is optimal according to the TSH. Assume, for instance, that government expenditure is suddenly expected to decrease in the future so that the present discounted value of expected changes in expenditure becomes negative. From (12), regardless of the degree of discretion, tax smoothing behavior implies that the tax rate is cut and set below total government expenditure in order to generate a budget deficit. If the government has relatively little discretion in budget policy, it cannot implement the full cut in the tax rate that is needed in order to smooth the tax rate perfectly and minimize the distortionary welfare costs. Specifically, in (13), λ becomes positive in order to reflect the fraction of the optimal budget deficit that cannot be implemented due to the "partial" balanced budget rule. The less discretion in budget policy, the higher the λ, and the less the resulting budget deficit. In the extreme and borderline case of no discretion in budget policy, i.e., in the case of a balanced budget rule, $\lambda = 1$ and the government cannot cut and smooth the

tax rate at all. By contrast, when $\lambda < 0$, the resulting budget deficit is larger than the budget deficit that would have occurred if the TSH had been true. That is, when government expenditure is expected to decrease, tax smoothing behavior in combination with more discretion in budget policy relative to what is optimal according to the TSH, lead to budget deficit overshooting. In this case, $-\lambda$ reflects the overshooting fraction of the budget deficit that is optimal according to the TSH.[6] If $\lambda = 0$, there is no overshooting of the optimal budget deficit and, thus, the result is the TSH. One can use analogous reasoning to conclude that an expected increase in government expenditure implies a relatively small budget surplus when $\lambda > 0$, and a relative large budget surplus when $\lambda < 0$.

Thus, in summary, the extent to which the degree of discretion in budget policy differs from the degree that is optimal according to the TSH, should be captured by λ. Specifically, λ is inversely related to the degree of discretion and has an upper bound equal to unity. In the benchmark TSH case, optimal discretion implies that $\lambda = 0$, because then the welfare costs from taxation are minimized. The interval $0 < \lambda < 1$ captures a government with less discretion (more rules) in budget policy relative to what is optimal according to the TSH.[7] By contrast, $\lambda < 0$ reflects a government with more discretion (less rules) in budget policy relative to what is optimal according to the TSH.

It should be emphasized that the case when λ takes on a value that is positive and less than unity does not necessarily reflect a government that in reality is bounded by more rules in budget policy than a government that is subject to an optimal degree of discretion. The reason is that it is theoretically possible for a government with much discretion to stick close to a balanced budget rule. Such a government,

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[6] That is, when $\lambda < 0$, $(1-\lambda) \sum_{s=t+1}^{\infty} \left(\frac{1}{1+R}\right)^{s-t} E_t \Delta g_s < \sum_{s=t+1}^{\infty} \left(\frac{1}{1+R}\right)^{s-t} E_t \Delta g_s < 0$, where $-\lambda > 0$ is the overshooting fraction of the optimal budget deficit.

[7] In their log-linear version of the model, Huang and Lin (1993) briefly formulate an alternative hypothesis that is analogous to the GTSH when $0 < \lambda < 1$. They evaluate this alternative by using preset values of λ. This paper goes further, as the next section shows how to obtain an estimate of λ.

however, behaves *as if* it were bounded by relatively more rules in budget policy. In this sense, λ captures the *effective* degree of discretion for that government. By contrast, because a partially rule-bound government can never deviate from a balanced budget rule to any extent that it wishes, the case when $\lambda < 0$ can only reflect a government with more discretion relative to what is optimal according to the TSH.

Finally, it should be noted that the GTSH retains the symmetric view of the budget balance. For instance, a government with a relatively discretionary budget policy ($\lambda < 0$) generates a large budget deficit in response to an expected decrease in its expenditure, but it also implements a tough "tighten the belt" policy and generates a large budget surplus in response to an expected increase in its expenditure. Intuitively, then the budget surplus time series should be stationary. This and other time series properties of the model are discussed in the next section.

3. Empirical method

As pointed out by Campbell (1987), even though expectations about future expenditure is conditional on the government's information set, I_t, it is still possible for an econometrician with access to only a subset of I_t to calculate the predicted path of the budget surplus from (13). This is because the budget surplus itself contains all information about future changes in government expenditure that is superior to the econometrician. As a consequence, the budget surplus should Granger-cause changes in government expenditure. Hence, by incorporating the budget surplus and changes in expenditure in the econometrician's information set, H_t, (13) *can* be estimated and, as shown below, the GTSH can be tested taking the government's superior information into account. The predicted path of the budget surplus is

$$\widehat{sur}_t = E(sur_t|H_t) = E\left\{E\left\{(1-\lambda)\sum_{s=t+1}^{\infty}\left(\frac{1}{1+R}\right)^{s-t}\Delta g_s\bigg|I_t\right\}\bigg|H_t\right\}$$

$$= (1-\lambda)\sum_{s=t+1}^{\infty}\left(\frac{1}{1+R}\right)^{s-t}E(\Delta g_s|H_t), \tag{14}$$

since $H_t \subseteq I_t$.

The GTSH formulated as a statement about the budget surplus reveals another important time series property. If government expenditure, g_t, contains a unit root, its first difference will be stationary. Accordingly, the budget surplus, sur_t, will be stationary because it is a linear combination of expected changes in expenditure. As noted by Ghosh (1995), although it is entirely possible that g_t in reality is a stationary time series, standard econometric tests generally cannot reject the null hypothesis that it contains a unit root. This is also the case for the Swedish data used below.

Furthermore, if τ_t and g_t^{TOT} are both individually I(1), the stationarity of sur_t implies that (8) defines a cointegration relation. Campbell (1987) shows that the implied error correction model can be expressed in VAR form as

$$\begin{bmatrix} \Delta g_t \\ sur_t \end{bmatrix} = \begin{bmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{bmatrix} \begin{bmatrix} \Delta g_{t-1} \\ sur_{t-1} \end{bmatrix} + \begin{bmatrix} \epsilon_{\Delta gt} \\ \epsilon_{surt} \end{bmatrix}, \tag{15}$$

where the means of Δg_t and sur_t have been removed.[8] After verifying that g_t is I(1), and that τ_t and g_t^{TOT} are both individually I(1) and cointegrated such that $\tau_t - g_t^{TOT} = sur_t$ is I(0), the VAR can be estimated in order to evaluate the GTSH. The procedure is as follows. Write the VAR in matrix notation as

$$X_t = AX_{t-1} + \varepsilon_t. \tag{16}$$

The forecast of a one period change in government expenditure is

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[8] For the annual data used in this paper, a one lag VAR is sufficient to capture the time series properties. If necessary, the VAR can easily be extended to incorporate several lags.

$$E(\Delta g_s | H_t) = \begin{bmatrix} 1 & 0 \end{bmatrix} A^{s-t} X_t. \tag{17}$$

Substituting (17) into (14) gives

$$
\begin{aligned}
\widehat{sur}_t &= (1-\lambda) \sum_{s=t+1}^{\infty} \left(\frac{1}{1+R} \right)^{s-t} \begin{bmatrix} 1 & 0 \end{bmatrix} A^{s-t} X_t \\
&= \begin{bmatrix} 1 & 0 \end{bmatrix} (1-\lambda) \frac{1}{1+R} A \left(I - \frac{1}{1+R} A \right)^{-1} X_t \\
&= \Lambda_1 \Delta g_t + \Lambda_2 sur_t. \tag{18}
\end{aligned}
$$

If the GTSH is true, the predicted budget surplus, $\widehat{sur}_t$, is equal to the actual budget surplus, sur_t, i.e., $\Lambda_1 = 0$ and $\Lambda_1 = 1$. Accordingly, the following overidentifying restrictions must hold for (18):

$$\begin{bmatrix} 1 & 0 \end{bmatrix} (1-\lambda) \frac{1}{1+R} A \left(I - \frac{1}{1+R} A \right)^{-1} = \begin{bmatrix} 0 & 1 \end{bmatrix}. \tag{19}$$

The restrictions in (19) represent the formally testable implications of the overall model, and for both the GTSH and the TSH below, they are evaluated by means of Wald and Likelihood ratio tests.[9] For the GTSH, an estimate of λ is obtained from the restricted model obtained in the calculation of the likelihood ratio statistic (Flavin, 1993). As indicated above, besides formal statistical testing, the fit of the model is also evaluated by calculating the predicted budget surplus according to (18) and then visually comparing it to the actual budget surplus.

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[9] When testing the TSH (i.e., when testing the restrictions in (19) with λ forced to zero), one can postmultiply (19) by $\left(I - \frac{1}{1+R} A \right)$ in order to obtain a simpler, linear expression to evaluate. In general, as noted by Campbell and Shiller (1987) and Gregory and Veall (1985), such a transformation can change the values and power of Wald statistics, and it may therefore be important to consider. However, for the data used in this paper, no conclusions are altered by focusing solely on testing the restrictions as expressed in (19). Test results based on the transformation yield very similar results and are available from the author upon request. Note that the likelihood ratio test statistic is invariant to any transformation of the restrictions.

4. Estimation and results

The data are taken from the *Statistical Yearbook of Sweden* (various issues), the International Monetary Fund's *International Financial Statistics* and *Government Finance Statistics* databases, and refer to the consolidated central government, i.e., the central government units covered by the general budget, and central government units with individual budgets, including the National Debt Office, the Swedish National Social Insurance Board, and regional agencies of the Public Health Insurance Society. Further details regarding the data and the construction of the variables are provided in the Appendix. The sample period is 1952-99, but separate estimation is also made for the sample period of 1970-96 as it is characterized by a relatively volatile budget surplus compared to the 1950s and 1960s. The choice of studying the period 1970-96 rather than the period 1970-99 is based on the fact that in the beginning of 1997, a new budget law aimed at strengthening the budget-making process, took effect.

The first step is to verify that g_t is I(1), and that τ_t and g_t^{TOT} are I(1) and cointegrated such that $\tau_t - g_t^{TOT} = sur_t$ is I(0). Table 1 displays results from augmented Dickey-Fuller (ADF) tests of the null hypothesis that the variables under consideration contain a unit root (for details, see Dickey and Fuller, 1981). It is well-known that the regression that the ADF test is derived from depends critically on the assumption of serially uncorrelated residuals. In order to verify this assumption, each ADF test statistic in Table 1 is supplemented by a Lagrange multiplier test statistic for the null hypothesis of no residual autocorrelation. The null hypothesis of a unit root cannot be rejected for τ_t, g_t and g_t^{TOT}. By contrast, for sur_t, the null is rejected, which suggests that τ_t and g_t^{TOT} are cointegrated. ADF tests are then performed for τ_t, g_t and g_t^{TOT} when they are expressed in their first differences. The null hypothesis of a unit root in the first difference of each series is rejected.

<div align="center">

Table 1

Augmented Dickey-Fuller (ADF) tests

</div>

Variable	1952-99		1970-96	
	ADF	LM	ADF	LM
τ_t	-1.60	0.33	-1.86	0.97
g_t	-1.92	2.11	-1.69	1.02
g_t^{TOT}	-1.95	2.24	-1.69	1.00
sur_t	-4.36^{**}	0.13	-3.36^{**}	0.22
$\Delta\tau_t$	-6.03^{**}	0.42	-4.05^{**}	1.03
Δg_t	-5.07^{**}	0.59	-3.85^{**}	0.71
Δg_t^{TOT}	-4.94^{**}	1.09	-3.86^{**}	0.71

Notes: ADF is the test statistic for the null hypothesis of a unit root. LM is the Lagrange multiplier test statistic for the null hypothesis of no residual correlation from lags 1 to 2. "**" indicates rejection at the 1 percent level of significance.

The ADF test for sur_t, τ_t and g_t^{TOT} is supplemented by a restricted cointegration analysis of the linear combination $(\delta_1, \delta_2)(\tau_t, g_t^{TOT})'$, by means of the Johansen (1988) procedure. In contrast to the ADF test, the Johansen (1988) procedure specifies I(0) as the null hypothesis. The results are summarized in Table 2. The first two columns of the table specify the restriction on the elements of the vector (δ_1, δ_2). The other columns report the test statistic and its corresponding p-value from the test of the null hypothesis that the specified vector belongs to the cointegrating space. Accordingly, the first row displays the test results of the null hypothesis that the vector $(1, -1)$ belongs to the cointegrating space such that $(1, -1)(\tau_t, g_t^{TOT})' = sur_t$ is I(0). As indicated, the null cannot be rejected in either period. The last two rows contain test results of the null hypothesis of stationarity of τ_t and g_t^{TOT}. For each variable in each time period, the null is rejected at the five percent level of significance or better. Thus, in sum, the results indicate that g_t is I(1), and that τ_t and g_t^{TOT} are I(1) and cointegrated such that $\tau_t - g_t^{TOT} = sur_t$ is I(0).

Table 2

Table 2

Tests on the cointegrating space

$\mathbf{H}_0$:		1952-99		1970-96	
δ_1	δ_2	$\chi^2(1)$	p-value	$\chi^2(1)$	p-value
1	-1	0.04	0.846	0.20	0.658
1	0	15.94	< 0.001	3.91	0.048
0	1	15.98	< 0.001	5.76	0.016

Notes: The null hypothesis is that the vector (δ_1, δ_2) belongs to the cointegrating space such that $(\delta_1, \delta_2)(\tau_t, g_t^{TOT})' = \delta_1 \tau_t + \delta_2 g_t^{TOT}$ is I(0). In the prior VAR estimation, for both sample periods, three lags were used in order to capture the time series properties and to assure no serial correlation among the residuals.

Next, the VAR for Δg_t and sur_t is estimated, and the results are shown in Table 3. Diagnostic checking and lag-length selection tests indicated that a one-lag VAR was sufficient to capture the time series properties. The null hypothesis that sur_{t-1} non-Granger causes Δg_t can be rejected at the 2.6 percent level of significance for the full sample period, and at the 1.8 percent level for the period 1970-96. Thus, for both periods, there is statistical evidence that the government has superior information.

Table 3

Estimated VAR coefficients

	1952-99	1970-96
$\widehat{a}_{11}$	0.327	0.257
	(0.137)	(0.187)
$\widehat{a}_{12}$	0.164	0.209
	(0.071)	(0.082)
$\widehat{a}_{21}$	-0.553	-0.717
	(0.172)	(0.280)
$\widehat{a}_{22}$	0.739	0.708
	(0.089)	(0.123)
t_{NG}	2.31	2.54
p-value	0.026	0.018

Notes: Standard errors in parentheses. t_{NG} is the test statistic for the null hypothesis that sur_{t-1} non-Granger causes Δg_t.

Table 4 summarizes the results from the overall tests of the model under each hypothesis. The left panel relates to Wald and likelihood-ratio tests of the TSH, i.e., tests of the restrictions in (19) with λ forced to zero, and the right panel relates to the corresponding tests of the GTSH. The test statistics indicate that it is not possible to statistically distinguish the GTSH from the TSH for the full sample period; neither hypothesis can be rejected. However, this is not the case for the period 1970-96. For this subperiod, the Wald statistic for the test of TSH is equal to 16.14, which implies that the null can be rejected at any conventional level of significance. The corresponding likelihood-ratio statistic is equal to 6.39, which implies that the null can be rejected at the 4.1 percent level of significance. By contrast, for the GTSH, both the Wald and likelihood-ratio statistics imply that the null cannot be rejected at any conventional level of significance.

Table 4

Tests of restrictions and estimates of λ

TSH:	1952-99	1970-96	GTSH:	1952-99	1970-96
$\chi^2_W(2)$	3.29	16.14	$\chi^2_W(1)$	0.03	1.07
p-value	0.193	< 0.001	p-value	0.863	0.302
$\chi^2_{LR}(2)$	2.82	6.39	$\chi^2_{LR}(1)$	0.05	2.27
p-value	0.244	0.041	p-value	0.821	0.132
			$\widehat{\lambda}$	-0.614	-0.899
			s.e. of $\widehat{\lambda}$	0.462	0.482

Notes: χ^2_W is the Wald statistic for the test of the restrictions in (19); χ^2_{LR} is the corresponding likelihood-ratio statistic.

The right panel of Table 4 also displays estimates of λ, which are generated in the estimation of the restricted models that are used in the calculations of the likelihood-ratio test statistics. For both periods, the estimated λ is negative and, accordingly, this suggests that there has been *more* discretion in Swedish budget policy *relative* to what is optimal according to the TSH. For the sample period 1952-99, the estimated λ is equal to -0.614, and is only weakly negative as it falls below 1.5 standard errors from zero. For the sample period 1970-96, the estimated

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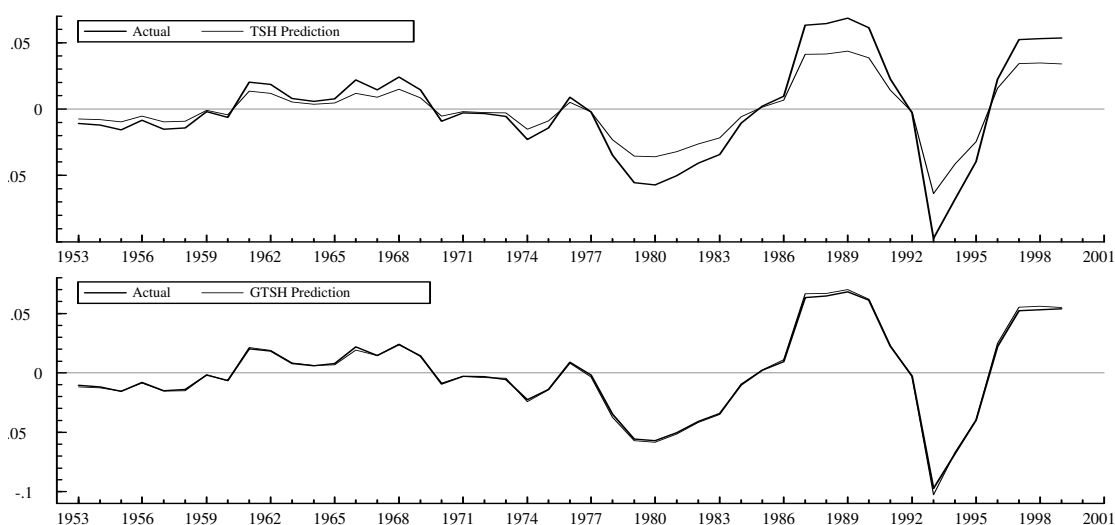λ is equal to -0.899, which is 1.87 standard errors from zero.

Fig. 1. *Predicted and actual budget surpluses, 1953-99.*
Notes: Upper panel: TSH; lower panel: GTSH.

Figure 1 plots the predicted and actual budget surplus time series for the full period. Although it is not possible to statistically distinguish between the two hypotheses, it is evident from the figure that the GTSH provides a better fit. In fact, the fit of the model when the GTSH is assumed is almost perfect. It is virtually impossible to see any difference between the actual and the predicted surpluses. By contrast, when the TSH is assumed, the actual budget surplus fluctuates more than the predicted surplus. Still, as suggested by the statistical tests, the predicted TSH budget surplus captures all shifts of the actual budget surplus.

Figure 2 plots the predicted and actual budget surpluses for the subperiod, and it can be interpreted in more or less the same way as Figure 1. It is worth noting that although the TSH is statistically rejected and the actual budget surplus fluctuates more than the predicted surplus, the latter can at least capture the major shifts in the former.
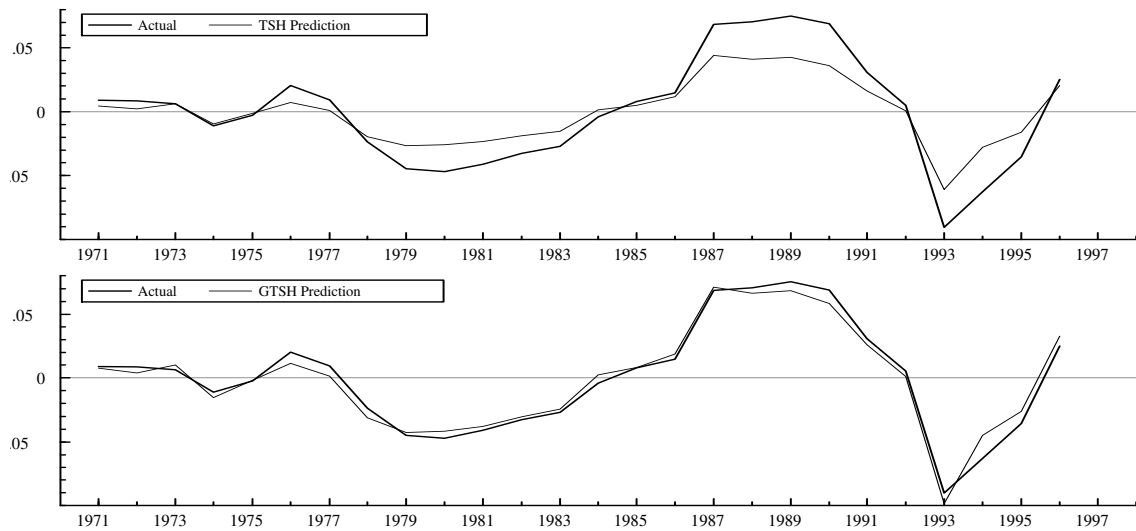
Fig. 2. *Predicted and actual budget surpluses, 1971-96.*
Notes: Upper panel: TSH; lower panel: GTSH.

5. Concluding remarks

Barro's (1979) tax smoothing hypothesis (TSH) assumes that the government is always subject to an "optimal" degree of discretion in budget policy, i.e., optimal in the sense that it minimizes the welfare costs from taxation. By contrast, the *generalized* tax smoothing hypothesis (GTSH) proposed in this paper takes into account the possibility that a tax smoothing government may be subject to a degree of discretion in budget policy that differs from the degree that is optimal according to the TSH.

The results indicate that the GTSH can go a remarkably long way in explaining the shifts in Sweden's central government budget balance. Formally, the GTSH cannot be rejected for either of the periods 1952-99 and 1970-96. The TSH cannot be rejected for the period 1952-99; however, it can be rejected for the period 1970-96. Hence, the result for the period 1970-96 indicates that the statistical power of the tests is high; even though the sample size is reduced, it is possible to distinguish

the GTSH from the TSH. Estimates of the parameter that reflect the degree of discretion, λ, suggest that there has been *more* discretion in budget policy relative to what is optimal according to the TSH. The visual results indicate that the GTSH provides close to a perfect model fit as the predicted and actual budget surpluses are almost identical. Hence, in summary, tax smoothing behavior in combination with more discretion in budget policy relative to what is optimal according to the TSH, can explain all shifts in the central government's budget balance. Accordingly, given a pure objective to smooth the tax rate perfectly, the budget law that was passed in 1997 appears to be a policy measure in line with the results obtained in this paper.

The GTSH should serve as a good platform for future studies on tax smoothing behavior among a wide range of governments. For instance, Strazicich (1996) notes that state governments in the U.S. that stick close to a balanced budget rule could still smooth their tax rates. Nevertheless, he rejects the TSH and concludes that state governments do not smooth tax rates. However, it is entirely possible that state governments *do* smooth their tax rates, but the TSH cannot capture this. By contrast, the GTSH can capture tax smoothing governments that stick close to a balanced budget rule. Thus, testing the GTSH on U.S. state governments and on similar cases may provide new insights and may even alter existing conclusions.

Appendix

All data are annual, and the full sample period is 1950-99. For the period 1950-94, each year refers to the fiscal year ending in June of the same year. For instance, 1950 refers to the budget year starting on July 1, 1949, and ending on June 30, 1950. For the period 1995-99, each year refers to the fiscal year starting on January 1 and ending on December 31.

Data for government expenditure, taxation receipts, GDP, and consumer price

index were taken from the International Monetary Fund's *International Financial Statistics* and *Government Finance Statistics* (GFS) for consolidated central government databases. Calendar year GDP was converted to fiscal year GDP by taking geometric means. Debt and interest payment on debt were taken from various issues of *Statistical Yearbook of Sweden.* Government expenditure is measured by the sum of total expenditure and lending minus repayment minus interest payment on the debt. Taxation receipts is measured by the sum of total revenue and grants. Outstanding debt, government expenditure, and taxation receipts were all divided by GDP in order to obtain $d_{t+1}, g_t,$ and τ_t.

The real interest rate, r, was constructed in a similar manner as in Olekalns (1997). First, the nominal rate on the debt was calculated by dividing interest payment on debt by outstanding debt. Then, the corresponding time period change in the consumer price index was subtracted from the nominal interest rate in order to obtain the budget year real interest rate. The real rate used in the calculations, r, was then set to equal the average of all budget year real interest rates. Following Ghosh (1995) and Olekalns (1997), the growth rate used in the calculations, n, was set to the average of the real GDP growth rates. Total government expenditure, g_t^{TOT}, was calculated as $g_t + (r - n)d_t$, and the budget surplus, sur_t, was calculated in accordance with (8) as $\tau_t - g_t^{TOT}$.

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