WP/00/165

IMF Working Paper

The Effects of Monetary and Fiscal Policy on Aggregate Demand in a Small Open Economy: An Application of the Structural Error Correction Model

Tetsuya Konuki

INTERNATIONAL MONETARY FUND

 WP/00/165

IMF Working Paper

Fiscal Affairs Department

The Effects of Monetary and Fiscal Policy on Aggregate Demand in a Small Open Economy: An Application of the Structural Error Correction Model

Prepared by Tetsuya Konuki [1]

Authorized for distribution by Sanjeev Gupta

October 2000

Abstract

This paper empirically analyzes the short-run effects of monetary and fiscal policy on aggregate demand, using the two-step structural error correction method. This method has an advantage over the standard reduced-form error correction method in providing a meaningful interpretation for impulse responses. The results are in sharp contrast to those of the traditional Mundell-Fleming and Dornbusch models: after the monetary (fiscal) policy is relaxed, the home currency depreciates (appreciates) for a substantial period of time, and the aggregate demand first expands (contracts) then gradually returns toward its original path.

JEL Classification Numbers: C32, E63, F41

Keywords: Monetary and fiscal policy, structural error correction models

Author's E-Mail Address: TKonuki@imf.org

---

[1] The author is grateful to Michael Devereux, Hamid Davoodi, Luiz de Mello, Anastassios Gagales, Eswar Prasad, Marijn Verhoeven, and seminar participants at the Ohio State University and Bloomsbury brown-bag meeting for their insightful comments. Special thanks are due to Masao Ogaki, for his helpful advice.

Contents

I. Introduction

The purpose of this paper is to analyze empirically the short-run effects of monetary and fiscal policy on aggregate demand. The small open economy model under flexible exchange rates is estimated. To identify structural monetary and fiscal disturbances, the two-step structural error correction method is applied as in Ogaki and Yang (1998). Cointegrating vectors are estimated in the first step, and the instrumental variable technique is applied to the system of equations in the second step. This method can overcome the difficulty faced by the standard error correction models (ECMs), as discussed below.

ECMs are widely used when some of the variables in the system are unit root nonstationary and are cointegrated. The standard ECMs are reduced-form models, as Urbain (1992) and Boswijk (1995) illustrate. Cooley and LeRoy's (1985) critique can be applied to the standard estimation methods for ECMs, such as Engle and Granger's (1987) two-step method and Johansen's (1991) Maximum Likelihood (ML) method: structural interpretations of impulse responses are invalid without restrictions on the structural model. To make impulse responses structurally interpretable, it is necessary to recover the structural disturbances of the system. A typical way to recover them is recursive ordering—the structural model is assumed to take lower block triangular form.

Unfortunately, in some cases, recursive ordering cannot be reconciled with an economic model. It is difficult to apply the standard ECM to policy evaluation even when the economic model under consideration suggests the application of the ECM. For example, the small open economy model used in this paper implies cointegration relationships among some variables but cannot take lower triangular form. In this case, the standard ECM cannot provide a meaningful interpretation of impulse responses from the viewpoint of economics. It is frustrating that the standard ECM is of little use in analyzing the relative effectiveness of monetary and fiscal policy under floating exchange rates, which is one of the central issues discussed among economists and policymakers since the seminal work of Fleming (1962).

The structural ECM (SECM), which was introduced by Boswijk (1995), might overcome the difficulty that the standard reduced-form ECMs face. It can provide a meaningful interpretation for impulse responses. However, Boswijk's method for estimating the SECM requires two stringent assumptions: some variables in a system must be exogenous, and the number of cointegrating relationships must equal the number of endogenous variables. These requirements strongly limit the applicability the SECM. Unlike Boswijk's method, the two-step method used in this paper for estimating the SECM does not require these stringent assumptions. It can be applied to a broad range of economic models, as Ogaki and Yang (1998) explain.

In this paper, a small open economy model under floating exchange rates is estimated using the aggregate quarterly data of Switzerland. The results are in sharp contrast to those of the traditional Mundell-Fleming model and the Dornbusch (1976) type of overshooting model: after the monetary (fiscal) policy is relaxed, the home currency depreciates (appreciates) for a substantial period of time, and aggregate demand first expands (contracts) then gradually returns toward its original path.

The results of parameter estimation imply a violation of the uncovered interest parity (UIP). They are consistent with the "forward premium puzzle," a classic result in the exchange rate literature.[2] As pointed out by Eichenbaum and Evans (1995), the forward premium puzzle plays an important role in producing a persistent depreciation induced by monetary expansion, which is in line with the widely observed empirical pattern of overshooting. However, this persistent depreciation is inconsistent with the Dornbusch model, which assumes the UIP and predicts a large immediate depreciation followed by subsequent appreciation.

The estimated aggregate demand function is relatively sensitive to the dynamics of real exchange rates, compared with money supply and fiscal spending. The stimulative effect of monetary relaxation is enhanced by depreciation. The expansionary effect of fiscal relaxation is offset by appreciation by more than 100 percent. This supports empirically the adverse impact of fiscal expansion under a small open economy—see discussion of its theoretical possibility in Devereux and Purvis (1990). The forward premium puzzle and sticky price adjustment lead to long-lasting real depreciation (appreciation). These long-lasting dynamics of real exchange rates produce the persistence of monetary-policy-induced boom and fiscal-policy-induced slump.

The remainder of this paper is organized as follows. Section II discusses the relative advantage of the SECM over the standard reduced-form ECM. Section III outlines the small open economy model to be estimated. Section IV explains the procedure for estimating parameters and deriving a SECM representation. Section V presents the empirical results. Concluding remarks are contained in Section VI.

II. Structural Models and Error Correction Models

This section discusses the relative advantage of the SECMs over the standard reduced-form ECMs. Let $y(t+1)$ be an n-dimensional vector of I(1), difference stationary, random variables. Suppose there exists ρ linearly independent cointegrating vectors, so that $A'y(t)$ is stationary, where A' is a $(\rho\times n)$ matrix of real numbers whose rows are linearly independent cointegrating vectors. Consider a standard ECM,

$$\Delta Z(t+1) = k_0 + \kappa_1 t + GA'Z(t) + F_1\Delta Z(t) + \cdots + F_p\Delta Z(t-p+1) + v(t+1), \qquad (1)$$

where $\Delta Z(t+1)=Z(t+1)-Z(t)$ for any variable $Z(t)$, k_0 and κ_1 are $(n\times 1)$ vector, G is a $(n\times\rho)$ matrix of real numbers, $v(t)$ is a stationary n-dimensional vector of real variables with

[2] For a survey, see Froot and Thaler (1990).

$E[v(t+1) \mid H(t-\tau)]=0$, where $H(t-\tau)$ is the econometrician's information set at $t-\tau$.[3] There are many methods to estimate (1), including Engle and Granger's two-step method and Johansen's ML method.

In many applications, (1) is a reduced-form model. A class of structural models can be written in the form of the SECM as

$$C_0\Delta Z(t+1) = d_0 + d_1 t + BA'Z(t) + C_1\Delta Z(t) + \cdots + C_p\Delta Z(t-p+1) + u(t+1). \qquad (2)$$

By premultiplying both sides of equation (2) by C_0^{-1}, we obtain the standard ECM (1) of a structural model, where $k_i = C_0^{-1}d_i$, $G = C_0^{-1}B$, $F_i = C_0^{-1}C_i$, and $v(t) = C_0^{-1}u(t)$. A standard ECM estimated by Engle and Granger's two-step method or Johansen's ML method is a reduced-form model of a structural model. Impulse response functions based on $v(t)$ cannot be interpreted in a structural way, unless some restrictions are imposed on C_0 to recover the structural disturbance term $u(t)$, as in a VAR.

Typically, C_0 is assumed to be block lower triangular with 1s along the principal diagonal. Then the relevant structural disturbance terms of $u(t)$ can be recovered. However, if C_0 cannot be block lower triangular by any ordering, which is the case in this paper (see Section IV), it is impossible to recover the relevant structural disturbance terms of $u(t)$. The standard reduced-form ECMs face difficulty in interpreting the impulse response functions. A block lower triangular C_0 in the cointegrated system might not be consistent with the underlying economic theory. The SECM can avoid the difficulty confronted by the standard ECMs.

III. Economic Model

An extended version of the Dornbusch (1976) model is employed to examine the response of aggregate demand to monetary and fiscal policy shocks.

The model uses money demand function, where the income elasticity of money is assumed to be one:[4]

$$m(t) = \kappa_m + \theta_m t + p(t) + y(t) - \lambda i(t) + \varepsilon_m(t), \qquad (3)$$

where $m(t)$ is the log of money supply, $p(t)$ the log domestic price, $y(t)$ the log real national income, $i(t)$ the nominal interest rate in the domestic country, θ_m a deterministic trend of

[3] In many applications, τ=0. But this paper uses an application in which τ>0.

[4] As shown in Section V, the statistical evidence against this money demand function is weak; so this assumption on income elasticity is not strong.

money demand, and $\varepsilon_m(t)$ the money demand shock, which is supposed to be stationary and orthogonal to $I(t\text{-}1)$, the economic agents' information set at time $t-1$. The other asset market equilibrium is

$$E[e(t+1)|I(t)] - e(t) = \beta(i(t) - i^*(t)), \tag{4}$$

where $e(t)$ is the log nominal exchange rate (the price of one unit of the foreign currency in terms of domestic currency), $E[\bullet \mid I(t)]$ the expectation operator conditional on $I(t)$, and $i^*(t)$ the nominal interest rate in the foreign country. The UIP, which assumes β=1, will be tested in Section V.

The goods side of the model is characterized as Mundell-Fleming simplifying assumption that the home country is completely specialized in production. This gives us the usual IS relation,

$$y^d(t) = \kappa_d + \theta_d t + bg(t) - \alpha(i(t) - E[\Delta p(t+1) \mid I(t)]) + \eta[e(t) + p^*(t) - p(t)] + \varepsilon_d(t), \tag{5}$$

where $y^d(t)$ is the log of real aggregate demand, $g(t)$ the log of real government spending, $p^*(t)$ the log foreign price, $e(t)+p^*(t)-p(t)$ the log real exchange rate, θ_d a deterministic trend of aggregate demand, and $\varepsilon_d(t)$ the exogenous shock to aggregate demand, which is stationary and orthogonal to $I(t\text{-}1)$.

The next step is the supply-side specification. As in a number of papers[5], labor demand is assumed to depend negatively on the product real wage, $w(t)-p(t)$, where $w(t)$ is the log of the nominal wage. The labor supply is assumed to be positively dependent on the consumption-based real wage, $w(t)-hp(t)-(1-h)[e(t)+p^*(t)]$. Under these assumptions, the equilibrium employment and output can be expressed as an increasing function of $p(t)-e(t)-p^*(t)$. In other words, real appreciation expands aggregate supply. This implies the aggregate supply function,

$$y^s(t) = \kappa_s + \theta_s t + \sigma[p(t) - e(t) - p^*(t)] + \varepsilon_s(t), \tag{6}$$

where θ_s is a deterministic trend of aggregate supply, and $\varepsilon_s(t)$ the aggregate supply disturbance, which is stationary and orthogonal to $I(t\text{-}1)$.

Finally, the model includes a Phillips-curve price adjustment equation,

$$\Delta p(t+1) = \tau + \pi[y^d(t) - y^s(t)] + \varepsilon_p(t+1), \tag{7}$$

[5] For a detailed exposition, see Sachs (1980).

where $\varepsilon_p(t+1)$ is the disturbance to price adjustment, which is stationary and orthogonal to $I(t)$. All parameters are assumed to be positive, except for constants and deterministic trend terms. As in the standard Dornbusch tradition, I proceed with the assumption that out of the long-run steady-state, the output is demand determined, that is, $y(t)=y^d(t)$.

IV. Econometric Methodology

This section explains the procedure for estimating parameters and deriving a SECM representation. To estimate parameters, the two-step procedure is applied as in Ogaki and Yang (1998). In the first step, cointegrating vectors are estimated. In the second step, the instrumental variable technique is applied to the system of equations plugged in the parameters' values estimated in the first step. As cointegrating regression estimators converge at a faster rate than the square root of the sample size, $T^{1/2}$, the first step estimation does not affect the asymptotic distributions of the second step estimators as Engle and Granger's (1987) explain.

The economic model in the previous section implies that some variables are cointegrated. Equation (3) implies that $[m(t)-p(t)-y^d(t),\ i(t)]$ are stochastically cointegrated by $[1, \lambda]$, if money demand is stable in the long run and if both $i(t)$ and $m(t)-p(t)-y^d(t)$ are I(1). Next, equations (6) and (7) provide

$$\Delta p(t+1) = \varphi + \mu t + \pi\left[y^d(t) + \sigma\left[e(t) + p^*(t) - p(t)\right]\right] + u_1(t+1), \qquad (8)$$

where $\varphi=\tau+\pi\kappa_s$, $\mu=\pi\theta_s$, $u_1(t-1)=\pi\varepsilon_s(t)+\varepsilon_p(t+1)$ and $E[u_1(t+1)\,|\,I(t-1)]=0$. Suppose $p(t)$ and $y^d(t)$ are I(1). Then equation (8) implies that $e(t)+p^*(t)-p(t)$ are I(1) and $[y^d(t),\ e(t)+p^*(t)-p(t)]$ are stochastically cointegrated by $[1, \sigma]$, since all variables—except for $y^d(t)$, which is I(1), and a deterministic trend—are stationary. The extended supply-side specification provides theoretical justification for the violation of the PPP which implies the stationarity of $e(t)+p^*(t)-p(t)$—many researchers have found supporting evidence for the violation of the PPP.

Choosing from the many methods for obtaining efficient estimators of cointegrating vectors, this paper employs Park's (1992) Canonical Cointegrating Regressions (CCR) for the four reasons. First, the CCR does not require Gaussian VAR assumption as does Johansen's method. Second, the Monte Carlo experiments in Park and Ogaki (1991) show that the CCR estimators have better small sample properties in terms of the mean square error than Johansen's estimators. Third, Park's (1990) tests of the null of deterministic cointegration and of stochastic cointegration have reasonable size and power, according to Han and Ogaki's (1997) Monte Carlo experiments. Fourth, cointegration is taken as the null hypothesis in the CCR procedure, whereas the standard testing procedures for cointegration—such as Johansen's method—take no cointegration as the null hypothesis. However, the standard tests are known to have very low power against some alternatives and may fail to reject the null of no cointegration with high probability even when the economic model that implies cointegration is actually consistent with data. It is preferable to test the null of cointegration to reduce the risk of rejecting a valid economic model. (Ogaki and Park (1997) discuss in detail the advantage of the CCR procedure over the standard approach for cointegration estimation.)

We can estimate λ, κ_m, θ_m by running CCR as

$$m(t) - p(t) - y^d(t) = \kappa_m + \theta_m t - \lambda i(t) + \varepsilon_m(t). \tag{9}$$

Note that the income elasticity of money demand is assumed to be one, as discussed in Section III. This assumption ensures that the money demand equation implies only one cointegrating vector and allows us to avoid the identification problem of multiple cointegrating vectors.[6] Similarly, σ, κ_s, θ_s can be estimated by CCR as

$$y^d(t) = \kappa_s + \theta_s t - \sigma[e(t) + p^*(t) - p(t)] + \varepsilon_s(t). \tag{10}$$

Plugging in the estimated parameter values obtained in the first step CCR, the instrumental variable technique is applied to estimate the remaining parameters of the system in the second step. Combining equations (4) and (9), the following empirical specification will be estimated:

$$\Delta e(t+1) = \kappa_e + \theta_e t - \frac{\beta}{\lambda}\left[m(t) - p(t) - y^d(t)\right] - \beta i^*(t) + u_2(t+1), \tag{11}$$

where $\kappa_e = \beta/\lambda$, $\theta_e = \beta\theta_m/\lambda$, $u_2(t+1) = \varepsilon_e(t+1) + \beta/\lambda\,\varepsilon_m(t)$, and $\varepsilon_e(t+1)$ is a one-period-ahead forecast error for exchange rate and $E[\varepsilon_e(t+1) \mid I(t)] = 0$.

Equations (3), (5), and (8) provide

$$\begin{aligned}\Delta y^d(t+1) = \theta_y &+ \gamma_1[\Delta m(t+1) - \Delta p(t+1)] + \gamma_2 \Delta g(t+1) \\ &+ \gamma_3[\Delta e(t+1) + \Delta p^*(t+1) - \Delta p(t+1)] + u_3(t+1),\end{aligned} \tag{12}$$

where

[6] Tests for the null of cointegration based on CCR assume that there is only one cointegrating vector and hence cannot be used in the case of multiple cointegrating vectors. Johansen's ML method has an advantage that it allows multiple cointegrating vectors. However, as pointed out by Ogaki (1993a) and among others, cointegrating vectors may not be identified even by the Johansen's ML method.

$$u_3(t+1) = \frac{1}{1-\alpha\pi+\alpha/\lambda}\left[\Delta\varepsilon_d(t+1)+\alpha\pi\Delta\varepsilon_s(t+1)-\frac{\alpha}{\lambda}\Delta\varepsilon_m(t+1)\right],$$

$$\theta_y = \frac{1}{1-\alpha\pi+\alpha/\lambda}\left(\alpha\pi\theta_s+\frac{\alpha}{\lambda}\right)\theta_m, \gamma_1 = \frac{\alpha/\lambda}{1-\alpha\pi+\alpha/\lambda}, \gamma_2 = \frac{b}{1-\alpha\pi+\alpha/\lambda}, \gamma_3 = \frac{\eta+\alpha\pi\sigma}{1-\alpha\pi+\alpha/\lambda}.$$

The empirical specifications for $\Delta m(t+1)$ and $\Delta g(t+1)$ are assumed to be a linear combination of its own past values and the current values of other variables of the system:

$$\begin{aligned}\Delta m(t+1) = {} & \Lambda_0+\Lambda_1\Delta m(t)+\Lambda_2\Delta m(t-1)+\cdots+\Lambda_p\Delta m(t-p+1)\\ & +\theta_1^m\Delta p(t+1)+\cdots+\vartheta_4^m\Delta g(t+1)+\cdots+\vartheta_6^m\Delta p^*(t+1)+u_4(t+1),\end{aligned} \quad (13)$$

$$\begin{aligned}\Delta g(t+1) = {} & \rho_0+\rho_1\Delta g(t)+\rho_2\Delta g(t-1)+\cdots\rho_p\Delta g(t-p+1)\\ & +\vartheta_1^g\Delta p(t+1)+\cdots+\vartheta_4^g\Delta m(t+1)+\cdots+\vartheta_6^g\Delta p^*(t+1)+u_5(t+1),\end{aligned} \quad (14)$$

where $u_i(t+1)$ is a disturbance to each specification and $E[u_i\ (t+1)\,|\,I(t)]=0$, for i=4, 5.

For the remaining two variables, $\Delta i^*(t+1)$ and $\Delta p^*(t+1)$, a small country assumption is imposed: neither domestic variables nor domestic disturbances affect the values of foreign variables. These two variables are specified as a linear combination of its own past values and current values of the other:

$$\Delta i^*(t+1) = \omega_0+\vartheta^{i^*}\Delta p^*(t+1)+\omega_1\Delta i^*(t)+\cdots+\omega_p\Delta i^*(t-p+1)+u_6(t+1), \quad (15)$$

$$\Delta p^*(t+1) = \xi_0+\vartheta^{p^*}\Delta i^*(t+1)+\xi_1\Delta p^*(t)+\cdots+\xi_p\Delta i^*(t-p+1)+u_7(t+1), \quad (16)$$

where $u_i(t+1)$ is a disturbance to each specification and $E[u_i\ (t+1)\,|\,I(t)]=0$, for i=6, 7.

Now we obtain the system of seven equations: (8), (11), (12), (13), (14), (15) and (16). Suppose $e(t)$, $m(t)$, $g(t)$, $i^*(t)$, and $p^*(t)$ are I(1), similar to $p(t)$, $y^d(t)$, $i(t)$, $m(t)-p(t)-y^d(t)$ and $e(t)+p^*(t)-p(t)$. Define $H(t)$ as the econometrician's information set at t which is generated by the current and past values of $\Delta p(t)$, $\Delta e(t)$, $\Delta y^d(t)$, $\Delta m(t)$, $\Delta g(t)$, $\Delta i^*(t)$, and $\Delta p^*(t)$. Obviously, $H(t)$ is a subset of $I(t)$, the economic agents' information set. By the Law of Iterated Expectations, $E[u_i(t+1)\,|\,H(t-1)]=0$ for i=1,2,3, and $E[u_i(t+1)\,|\,H(t)]=0$ for i=4,5,6,7. Define $H^*(t-1)-H(t-1)\cap\{\Delta i^*(t), \Delta p^*(t)\}$. By the small country assumption, $\varepsilon_m(t)$, $\varepsilon_s(t)$, and $\varepsilon_d(t)$ are orthogonal to $\{\Delta i^*(t), \Delta p^*(t)\}$: disturbances to money demand, aggregate supply, and aggregate demand in the home country do not affect foreign variables. This implies $E[u_i(t+1)\,|\,H^*(t-1)]=0$ for i=1,2,3. In addition, only stationary variables are involved in $H(t)$,

$H^*(t-1)$, and the system of seven equations after the parameters λ, θ_m, σ, θ_s are estimated in the first step CCR. We can estimate the remaining parameters by applying the Generalized Method of Moments (GMM) to these seven equations, using the variables in $H(t)$ and $H^*(t-1)$ as instruments in the second step. (See the appendix for the detailed explanation of the GMM procedure used in this paper.)

Note that the system of these seven equations can be written in a form as equation (2): SECM with $Z(t)=[p(t), e(t), y^d(t), m(t), g(t), i^*(t), p^*(t)]$. Suppose neither ϑ_4^m nor ϑ_4^g equals zero. Then it is obvious that this system cannot take block lower triangular form by any ordering. In such a case, there is no way to distinguish between monetary shock, $u_4(t+1)$, and fiscal shock, $u_5(t+1)$, if we estimate the reduced form, as does Johansen's ML method. As explained in Section II, the procedure that this paper follows makes possible the structural interpretation of impulse responses of the ECMs. In addition, this procedure has an advantage over the SECM approach developed by Boswijk (1995): unlike Boswijk's approach, neither the distinction between exogenous and endogenous variables nor the equality between the number of cointegrating relationships and the number of endogenous variables is required to derive and estimate the SECM representation. The two-step technique used in this paper can be applied to a wide range of economic models.

We need to confront one final problem before estimating impulse responses. The specification of $\varepsilon_e(t+1)$, the one-period-ahead forecast error for exchange rate, needs to be defined. Because this forecast error is orthogonal to $I(t)$, it should be assumed to be a linear combination of variables that are orthogonal to $I(t)$. In other words, the one-period-ahead forecast error made by rational economic agents based on $I(t)$ should be a function of the shocks to the system at t+1. In this paper, it is specified as a linear combination of disturbances to other variables, $u_i(t+1)$, $i=4,\ldots,7$, and $\varepsilon_m(t+1)$, all of which are orthogonal to $I(t)$ and observable to econometricians at t+1, plus its own disturbance term $\nu_e(t+1)$. The specification

$$\varepsilon_e(t+1) = \beta_4 u_4(t+1) + \beta_5 u_5(t+1) + \cdots + \beta_7 u_7(t+1) + \beta_m \varepsilon_m(t+1) + \nu_e(t+1),$$

will be estimated prior to impulse response analysis.

V. Empirical Results

In this section, the data, integration property of the data, results of parameter estimation, and dynamic effects of monetary and fiscal disturbances are explained.

A. The Data

The sampling interval of the data is quarterly and the sample period extends from 1975Q4 to 1996Q4. Switzerland is used as the home country and the G7 countries as foreign countries. This choice is motivated by the fact that Switzerland adopts an independent floating exchange system and by its economic size. The nominal exchange rate is the weighted average of the Swiss franc's end-of-period price of one unit of each G7 currency. The trade weight of each G7 country is used as a weight.[7] The CPI is used to measure prices. Note that $m(t)-p(t)$ in the money demand function is a desired real balance and that $e(t)+p^*(t)-p(t)$ is the real exchange rate that consumers face. It is appropriate to use the CPI, rather than the GDP deflator, as the price measure. Foreign price is the weighted (by trade weights) average of G7 countries' CPI. The data for interest rates are money market rates with three-month maturity, but three-month T-bill rates are employed for United Kingdom and United States. Again, the foreign interest rates are calculated using the trade weight of each country with Switzerland. GDP is used as the aggregate demand measure. M2—the sum of M1 and Quasi Money—is used to obtain $m(t)$, as *International Financial Statistics* (IFS) suggests. General government spending is used to measure $g(t)$. GDP and general government spending are deflated by CPI to obtain their real values. The data of GDP and government spending are extracted from the OECD database. The IMF's *Direction of Trade Statistics* is used to calculate the trade weights. All remaining data are from the IFS database. Except for exchange rates and interest rates (of both the home and foreign countries), the data are seasonally adjusted.

B. Integration Properties of the Data and Parameter Estimation

To test the null of unit root nonstationarity for relevant variables which is assumed in the previous section, Park's (1990) J(1,5) test is applied. This test does not require the estimation of the long-run variance and has an advantage over Phillips and Perron's (1988) test and Said and Dickey's (1984) test in that neither the bandwidth parameter nor the order of autoregression needs to be chosen. Park and Choi's (1988) Monte Carlo experiments show that the J(1,5) test has stable size and is not dominated by Phillips and Perron's or Said and Dickey's test in terms of powers in small samples (see Ogaki (1993a)).

Table 1 reports the test results. The null of the difference stationarity of $e(t)$, $p(t)$, $p^*(t)$, $i(t)$, $i^*(t)$, $y^d(t)$, $m(t)$, $g(t)$, $e(t)+p^*(t)-p(t)$, and $m(t)-p(t)-y^d(t)$ is tested against the alternative of their trend stationarity.[8] The null hypothesis is rejected when J(1,5) is small, because it converges to zero under the alternative hypothesis. Except for the nominal exchange rate, the null of difference stationarity cannot be rejected at any reasonable level of significance, which is consistent with the assumption. The J(1,5) test rejects the null of the difference stationarity of $e(t)$ at the 10 percent level. However, it cannot reject the null at the 5 percent level. Thus, the results are in favor of the assumption made in the previous section.

---

[7] The weight for country 1 is calculated as $Weight_{1t} = (X_{1t} + M_{1t}) \Big/ \sum_{i=1}^{7} (X_{it} + M_{it})$, where X_{1t} denotes exports, and M_{1t} imports, for country 1 at period t.

[8] As to $e(t)+p^*(t)-p(t)$, Park's J(0,3) test is applied to test the null of its difference stationarity against the alternative of its stationarity. The test does not reject the null at any reasonable level of significance, which is consistent with the assumption.

Table 1. Tests for Integration Properties of the Relevant Variables

	$e(t)$	$p(t)$	$p^*(t)$	$i(t)$	$i^*(t)$
J(1,5)	0.359*	2.231	36.015	1.147	1.017
	$y^d(t)$	$m(t)$	$g(t)$	$e(t)+p^*(t)-p(t)$	$m(t)-p(t)-y^d(t)$
J(1,5)	2.508	5.832	7.611	0.615	1.289

Note: Critical values of J(1,5) tests for the 1 percent, 5 percent, and 10 percent significance levels are 0.123, 0.295, and 0.452, (Park and Choi (1988)), respectively.
* Significant at the 10 percent level.

Table 2 reports the CCR results[9] of money demand equation, defined by equation (9). H(1,2) and H(1,3) statistics test the null of stochastic cointegration that is assumed in the previous section. The CCR procedure needs estimates of long-run covariance parameters; for this purpose, I employ Park and Ogaki's (1991) VAR prewhitening method with Andrew's (1991) automatic bandwidth parameter. Following the recommendation of Park and Ogaki (1991), I report the CCR estimators based on the third stage. Test statistics for the null of stochastic cointegration from the fourth stage are used, as recommended by Han and Ogaki (1997). The estimate of λ is significantly positive, which is consistent with the economic theory. Neither H(1,2) nor H(1,3) rejects the stochastic cointegrating restriction at any reasonable significance. The evidence against the money demand equation used in this paper is very weak.

The results of CCR as equation (10) are reported in Table 3. As the economic model implies, the sign of estimated σ is positive although the estimation is not precise. H(1,2) rejects the null of stochastic cointegration at the 10 percent level but it does not reject the null at the 5 percent level. H(1,3) cannot reject the null at any reasonable level of significance. These results are in favor of the supply-side and price adjustment specification assumed in Section III.

[9] All CCR estimations in this paper use Ogaki's (1993b) GAUSS CCR package.

Table 2. CCR Results of Equation (9)

θ_m	λ	H(1,2)	H(1,3)
0.004	7.924	1.459	2.834
(0.001)	(1.910)	[0.227]	[0.242]

Note: Asymptotic standard error is in parenthesis, and asymptotic p-values are in brackets. H(1,2) is asymptotically distributed as χ_1^2, and H(1,3) is asymptotically distributed as χ_2^2.

Table 3. CCR Results of Equation (10)

θ_s	σ	H(1,2)	H(1,3)
0.005	0.070	2.753*	3.067
(0.001)	(0.200)	[0.097]	[0.216]

Note: See Table 2.
* Significant at the 10 percent level.

Table 4 summarizes the results of GMM estimation of the system of seven equations: (8), (11), (12), (13), (14), (15), and (16)[10]. CCR estimates of λ, θ_m, σ, θ_s are plugged into equations (8) and (11). The single equation method is applied, as explained in the appendix. The estimators given by this method are more robust than those given by the system method because misspecification in other equations does not affect their consistency.

[10] To save space, some of the parameter estimates are not reported in Table 4. These estimates are available upon request.

Table 4. GMM Results of the System of Seven Equations—(8), (11), (12), (13), (14), (15), and (16)

		Equation (8)	
	π=0.091 (0.015)		J-stat=6.836 [0.654]
		Equation (11)	
	β=–0.842 (0.523)		J-stat=9.361 [0.672]
		Equation (12)	
γ_1=0.290 (0.266)	γ_2=0.552 (1.180)	γ_3=0.440 (0.125)	J-stat=12.318 [0.340]
		Equation (13)	
	ϑ_4^m=–0.634 (0.285)		J-stat=11.748 [0.860]
		Equation (14)	
	ϑ_4^g=–0.116 (0.054)		J-stat=19.931 [0.337]
		Equation (15)	
			J-stat=1.831 [0.608]
		Equation (16)	
			J-stat=3.976 [0.264]

Note: See Table 2. J-stat is asymptotically distributed as χ_9^2 for equation (8), χ_{12}^2 for equation (11), χ_{11}^2 for equation (12), χ_{18}^2 for equations (13) and (14), and χ_3^2 for equations (15) and (16).

Hansen's (1982) J-statistic does not reject any of these seven specifications at any reasonable significance. The first row reports the parameter estimation for equation (8), price adjustment specification. The estimated value of price adjustment, π, is significantly positive, as the economic model predicts. The estimation for equation (11) is reported in the second row. The point estimate of β is negative and the hypothesis of β=1, UIP, can be decisively rejected. This is consistent with the "forward premium puzzle," a classic result in the exchange rate literature. With β<0, the more the domestic interest rate exceeds the foreign interest rate, the more the domestic currency appreciates over the holding period. If economic agents are rational, β<0 implies they are risk-averse. In addition, as Fama (1984) explains, this implies that the variance of risk premium is greater than the variance of both expected appreciation and the interest differential, and that the covariance of expected appreciation and the risk premium is negative.[11]

The third row in Table 4 shows the estimation for equation (12), aggregate demand specification. The point estimates of γ_1, γ_2 and γ_3 suggest the negative interest elasticity of aggregate demand ($-\alpha$), positive fiscal elasticity (b), and positive real exchange rate elasticity (η). These results are consistent with the economic theory assumed in Section III although the estimation is not precise for γ_1 and γ_2. The real exchange rate's multiplier (γ_3) is larger than the monetary multiplier (γ_1). The fiscal multiplier (γ_2) has the largest value among these three multipliers. As seen in the next subsection, the dynamics of real exchange rates imposes a significant impact on the aggregate demand.

The fourth and fifth rows summarize the estimations for equations (13) and (14), specifications for money supply and government spending. The estimates for ϑ_4^m and ϑ_4^g are negative and significantly different from zero. This implies that monetary policy is tightened when the fiscal policy is loosened and fiscal policy is tightened when the monetary policy is loosened. And these counter-acting relations are significant. It is not appropriate to assume block lower triangular form in order to identify these two shocks. Only the estimation based on structural form can distinguish between monetary and fiscal disturbances.

In summary, GMM estimations suggest that the data do not show strong evidence against the economic theory used in this paper: none of the specifications are rejected and signs of parameters are consistent with economic theory assumed in this paper. Only the UIP has been rejected with $\beta<0$; however, the negative value of β might be explained by the existence of risk premiums, as discussed above.

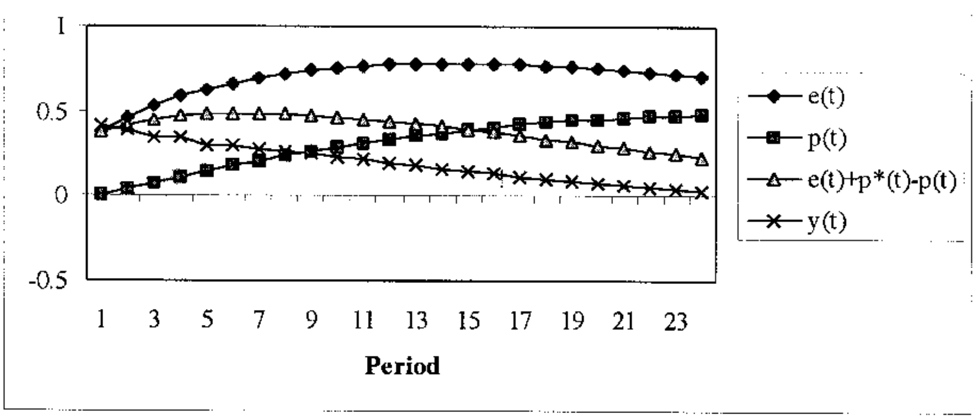
C. Dynamic Effects and Variance Decompositions of Monetary and Fiscal Disturbances

The dynamic effects of monetary and fiscal disturbances are reported in Figures 1 and 2, respectively. The vertical axes denote the responses of the log nominal exchange rate, log domestic price, log real exchange rate, and log real aggregate demand to one-unit impulses in the monetary and fiscal disturbances in the initial period. The horizontal axis denotes time in quarters.

Figure 1 provides a response pattern that is different from the one predicted by the Dornbusch (1976) type models. Positive monetary disturbance leads to a persistent depreciation that magnifies, rather than dampens. The maximal impact on nominal and real exchange rates occurs at 13 and 5 quarters, respectively, after the monetary policy is expanded. This is inconsistent with the simple overshooting models, which assume the UIP.

[11] Discussion on the forward premium bias is beyond the scope of this paper. See Froot and Thaler (1990) and Eichenbaum and Evans (1995) for detailed discussions. Konuki (1999) conducts a race between risk-premium models in terms of their ability to explain observed violations of UIP.

Figure 1: Response to Monetary Shock

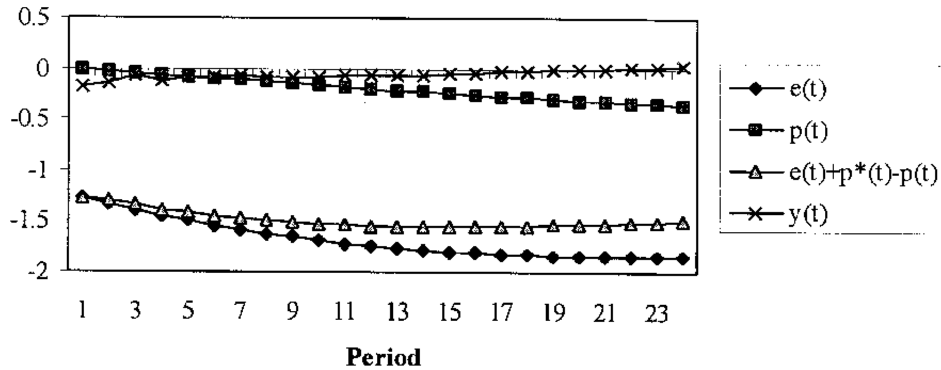


In those models, expansionary monetary policy shock generates large initial depreciation in nominal and real exchange rates, followed by subsequent appreciation. The response pattern of exchange rates in Figure 1 could be viewed as supporting a broader view of overshooting, in which exchange rates eventually appreciate after depreciating for a period of time, as is pointed out by Eichenbaum and Evans (1995). The key to these exchange rates' responses is the negative estimates of β in equation (8): an expansionary monetary shock leads to a fall in $i(t)-i^*(t)$ and a persistent depreciation that magnifies, rather than dampens, the excess returns associated with investing in foreign countries. Forward premium puzzle could be a background to the widely recognized empirical pattern of overshooting.

The domestic price level increases gradually after an expansionary monetary shock. This domestic price dynamics reflects the expansion of aggregate demand through decline in domestic interest rate and real depreciation and the contraction of aggregate supply through real depreciation, both of which produce inflationary pressure. The aggregate demand jumps at the initial period of monetary shock due to the immediate fall in interest rate and (nominal and real) appreciation. Subsequently, the stimulative effect of monetary easing starts to dampen through the price adjustment. The positive impact of monetary policy on aggregate demand is persistent because of the magnifying depreciation and slow adjustment of domestic prices.

Figure 2 displays somewhat surprising results. The response of exchange rates to positive fiscal disturbance is the mirror image of the response to monetary disturbance. Positive fiscal disturbance leads to a persistent appreciation that magnifies, rather than dampens. The maximal impact on nominal and real exchange rates is observed 22 and 15 quarters, respectively, after the fiscal policy is loosened. The surprising aspects are impulse responses

Figure 2: Response to Fiscal Shock



of domestic price and aggregate demand. The domestic price level decreases, rather than increases, gradually after an expansionary fiscal shock. This response pattern of the price level is attributable to the contractionary effect on aggregate demand and expansionary effect on aggregate supply through real appreciation, both of which produce disinflationary pressure. This disinflationary impact dominates the inflationary impact of fiscal expansion.

The aggregate demand drops, rather than jumps, at the initial period of fiscal shock owing to the immediate (nominal and real) appreciation. Subsequently, the contractionary effect of fiscal expansion starts to dampen through the price adjustment. However, slow price adjustment and magnifying nominal appreciation produce the persistent real appreciation, which hinders the recovery of aggregate demand through the fall in price level. Fiscal expansion starts to take positive effect on aggregate demand 21 quarters after the shock.

This response pattern is quite different from the prediction of the original Mundell-Fleming model, which was that a fiscal expansion would be matched by a 100 percent crowding out of aggregate demand through real appreciation and a fall in net exports. On the contrary, it might be supporting evidence for the theoretical possibility of an adverse impact of fiscal expansion on aggregate demand in the short-run, which is discussed by Devereux and Purvis (1990).[12] Note that the system under consideration consists of nonstationary variables. Households and firms could perceive a one-shot fiscal disturbance as a permanent fiscal expansion. A permanent fiscal expansion might lead economic agents to expect that an initial appreciation will persist, and this will produce a further appreciation, which will add to crowding out. (See Krugman and Obstfeld (1997).) The relatively large estimated value of the real exchange rate's multiplier (γ_3) plays an important role in producing this short-run

[12] For detailed discussion on this theoretical possibility, see Devereux and Purvis (1990).

Table 5. Variance Decomposition of Aggregate Demand

Horizon (quarters)	Variance due to *m(t)* / Variance due to *g(t)*
1	28.405
2	34.369
3	43.426
4	42.231
8	48.188
12	47.547
16	47.208
20	47.837
24	47.842

adverse impact of fiscal expansion. This implies that the Swiss economy is so sensitive to real exchange rates that the fiscal expansion cannot stimulate aggregate demand for a substantial period of time. In addition, persistent and magnifying real appreciation makes adverse impact on aggregate demand sticky.

The response pattern of aggregate demand to a positive fiscal shock might partly explain the prolonged recession of the Swiss economy in the early 1990s. Fiscal policy was relaxed at the onset of recession while monetary policy was tightened as the parameter estimates for equation (13) imply. This fiscal relaxation combined with monetary tightening was accompanied by real appreciation and the recession continued for about three years.[13] The switch to an expansionary monetary stance played an important role for the recovery (see IMF (1994)). It is often pointed out that the Swiss economy has a small fiscal multiplier, owing to the high degree of openness. The negative effect of real appreciation could dominate the positive effect of fiscal expansion.[14]

Table 5 shows the variance decompositions of aggregate demand at various horizons. The number in the second column of the table reports the contribution of monetary disturbances to the variance of aggregate demand's forecast errors at each horizon (number in the first column) divided by the contribution of fiscal disturbances. Throughout all of the horizons, the contribution of monetary disturbance is much more important than that of fiscal disturbance. This result reflects the fact that an expansionary monetary policy causes persistent real depreciation, which enhances the stimulative effect. Fiscal policy leads to persistent real appreciation, which offsets the stimulative effect.

---

[13] It is pointed out that external factors, such as a portfolio shift into Swiss francs caused by doubts about the ECU, have exacerbated the appreciation (IMF (1994)).

[14] Some caution is required in interpreting the results of these simulations. The estimated value of real exchange rate's multiplier might be biased upwards due to the specification of aggregate demand.

VI. Conclusion

This paper compares the impulse responses of aggregate demand to monetary and fiscal policy disturbances. The two-step structural error correction method is applied to a small open economy under floating exchange rates. This method has an advantage over the standard reduced-form ECM in providing a meaningful interpretation for impulse responses. Unlike Boswijk's method for estimating SECM, this method does not require the distinction between exogenous and endogenous variables by econometricians; nor does it require the assumption that thc cointegrating rank equals the number of endogenous variables. It can be applied to a broad range of economic models.

The results are in sharp contrast to those of the traditional Mundell-Fleming model and the Dornbusch (1976) type of overshooting model. After the monetary (fiscal) policy is relaxed, the home currency depreciates (appreciates) for a substantial period of time, and the aggregate demands first expands (contracts) then gradually returns toward its original path. The violation of the UIP plays a key role in producing a broader view of overshooting in which the exchange rate appreciates after depreciating for a period of time. The estimated aggregate demand function is relatively sensitive to the dynamics of the real exchange rate, compared with money supply and fiscal spending. The stimulative effect of monetary expansion is magnified by depreciation. The expansionary effect of fiscal expansion is dominated by the negative impact of appreciation in the short run. The fiscal-policy-induced slump continues for a substantial period of time, because of subsequent appreciation and sticky price adjustments.

This paper conducts policy evaluation in a small open economy, using the data of Switzerland. The fiscal multiplier of the Swiss economy, as pointed out, is small, owing to its openness. The methodology used in this paper can be applied to a large open economy whose aggregate demand is relatively insensitive to real exchange rates; for example, using the United States as a home country. Such an extension might provide quite different policy implications for the home country. In addition, it might estimate the policy impact of a large economy on small foreign countries.

Detailed Procedure of GMM Estimation

The detailed procedure of GMM estimation is discussed in this appendix. As discussed in Section IV, $E[u_i(t+1) \mid H^*(t-1)]=0$ for i=1,2,3, GMM disturbances for equations (8), (11), and (12), respectively, while $E[u_i(t+1) \mid H(t)]=0$ for i=4,5,6,7, GMM disturbances for equations (13), (14), (15), and (16), respectively.

Note that $E[u_i(t+1) \mid I(t-1)]=0$ and $u_i(t+1)$ is available in I(t+1) for i=1,2,3. They have a moving average (MA) representation of order one, and serial correlation needs to be taken into account. In the estimation of the long-run covariance matrix of GMM disturbance (Ω) of equations (8), (11), and (12), I applied a truncated kernel estimator, setting the lag truncation number as one. When Ω_T is not positive semidefinite, VAR prewhitening and a QS kernel estimator are employed, because existing Monte Carlo evidence recommends this method. (See Andrews (1991).) Serial correlation is not taken into consideration for GMM disturbances for equations (13), (14), (15), and (16), because $E[u_i(t+1) \mid I(t)]=0$ for i=4,5,6,7.

There are two methods for applying GMM to equations (8), (11), (12), (13), (14), (15), and (16): the single equation method, which estimates equation by equation, and the system method, which estimates these seven equations simultaneously. The single equation method is used in this paper, because it provides more robust estimators than system method — misspecification in other equations does not affect their consistency. (GMM failed to converge when I applied system method.)

I used the Hansen-Heaton-Ogaki GAUSS GMM Package for GMM estimation. In estimating equation (8), {constant, $\Delta e(t-1)$, $\Delta e(t-2)$, $\Delta e(t-3)$, $\Delta y^d(t-1)$, $\Delta y^d(t-2)$, $\Delta y^d(t-3)$, $\Delta i^*(t)$, $\Delta i^*(t-1)$, $\Delta i^*(t-2)$, $\Delta i^*(t-3)$} are used as instruments. For equation (11), {constant, $\Delta p(t-1)$, $\Delta p(t-2)$, $\Delta p(t-3)$, $\Delta y^d(t-1)$, $\Delta y^d(t-2)$, $\Delta y^d(t-3)$, $\Delta m(t-1)$, $\Delta m(t-2)$, $\Delta m(t-3)$, $\Delta i^*(t)$, $\Delta i^*(t-1)$, $\Delta i^*(t-2)$, $\Delta i^*(t-3)$} are used as instruments. For equation (12), {constant, $\Delta m(t-1)$, $\Delta m(t-2)$, $\Delta m(t-3)$, $\Delta g(t-1)$, $\Delta g(t-2)$, $\Delta g(t-3)$, $\Delta i^*(t)$, $\Delta i^*(t-1)$, $\Delta i^*(t-2)$, $\Delta i^*(t-3)$, $\Delta p^*(t)$, $\Delta p^*(t-1)$, $\Delta p^*(t-2)$, $\Delta p^*(t-3)$} are used as instruments. Equations (13) and (14) with lag order p=4 are estimated, and {constant, $\Delta p(t)$, $\Delta p(t-1)$, $\Delta p(t-2)$, $\Delta p(t-3)$, $\Delta e(t)$, $\Delta e(t-1)$, $\Delta e(t-2)$, $\Delta e(t-3)$, $\Delta y^d(t)$, $\Delta y^d(t-1)$, $\Delta y^d(t-2)$, $\Delta y^d(t-3)$, $\Delta m(t)$, $\Delta m(t-1)$, $\Delta m(t-2)$, $\Delta m(t-3)$, $\Delta g(t)$, $\Delta g(t-1)$, $\Delta g(t-2)$, $\Delta g(t-3)$, $\Delta i^*(t)$, $\Delta i^*(t-1)$, $\Delta i^*(t-2)$, $\Delta i^*(t-3)$, $\Delta p^*(t)$, $\Delta p^*(t-1)$, $\Delta p^*(t-2)$, $\Delta p^*(t-3)$} are used as instruments. Equations (15) and (16) with lag order p=4 are estimated, and {constant, $\Delta i^*(t)$, $\Delta i^*(t-1)$, $\Delta i^*(t-2)$, $\Delta i^*(t-3)$, $\Delta p^*(t)$, $\Delta p^*(t-1)$, $\Delta p^*(t-2)$, $\Delta p^*(t-3)$} are used as instruments. In Table 4, I report the GMM estimation results with the fifth iteration on the weighting matrix.

References

Andrews, Donald W.K., 1991, "Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimation," *Econometrica,* Vol. 59 (May), pp. 817–59.

Boswijk, H. Peter, 1995, "Efficient Inference on Cointegration Parameters in Structural Error Correction Models," *Journal of Econometrics,* Vol. 69 (September), pp. 133–58.

Cooley, Thomas F., and Stephen F. Leroy, 1985, "Atheoretical Macroeconometrics: A Critique," *Journal of Monetary Economics,* Vol. 16 (November), pp. 283–308.

Devereux, Michael B., and Douglas D. Purvis, 1990, "Fiscal Policy and the Real Exchange Rate," *European Economic Review,* Vol. 34 (September), pp. 1201–11.

Dornbusch, Rudiger, 1976, "Expectations and Exchange Rate Dynamics," *Journal of Political Economy,* Vol. 84 (December), pp. 1161–76.

Eichenbaum, Martin, and Charles L. Evans, 1995, "Some Empirical Evidence on the Effects of Shocks to Monetary Policy on Exchange Rates," *Quarterly Journal of Economics,* Vol. 110 (November), pp. 975–1009.

Engle, Robert F., and Clive W.J. Granger, 1987, "Cointegration and Error Correction: Representation, Estimation, and Testing," *Econometrica,* Vol. 55 (March), pp. 251–76.

Fama, Eugene F., 1984, "Forward and Spot Exchange Rates," *Journal of Monetary Economics,* Vol. 14 (November), pp. 697–703.

Fleming, J. Marcus, 1962, "Domestic Financial Policies under Fixed and Under Floating Exchange Rates," *IMF Staff Papers,* International Monetary Fund, Vol. 9 (November), pp. 369–79.

Froot, Kenneth A., and Richard H. Thaler, 1990, "Anomalies: Foreign Exchange," *Journal of Economic Perspectives,* Vol. 4 (Summer), pp. 179–92.

Han, Hsiang-Ling, and Masao Ogaki, 1997, "Consumption, Income, and Cointegration: Further Analysis," *International Review of Economics and Finance,* Vol. 6, pp. 107–17.

Hansen, Lars P., 1982, "Large Sample Properties of Generalized Method of Moment Estimators," *Econometrica,* Vol. 50 (July), pp. 1029–54.

International Monetary Fund, SM/94/4, 1994, "Switzerland—Economic Developments and Issues."

Johansen, Soren, 1991, "Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models," *Econometrica,* Vol. 59 (November), pp. 1551–80.

Konuki, Tetsuya, 1999, "Measuring Noise in Exchange Rate Models," *Journal of International Economics,* Vol. 48 (August), pp. 255–70.

Krugman, Paul R., and Maurice Obstfeld, 1997, *International Economics: Theory and Policy* (Massachusetts: Addison-Wesley, 4th ed.).

Ogaki, Masao, 1993a, "Unit Roots in Macroeconometrics: a Survey," *Monetary and Economic 8Studies,* Vol. 11 (November), pp. 131–54.

———, 1993b. "CCR: A User Guide," RCER Working Paper No. 349 (Rochester, New York: University of Rochester).

Ogaki, Masao, and Joon Y. Park, 1997, "A Cointegration Approach to Estimating Preference Parameters," *Journal of Econometrics,* Vol. 82 (December), pp. 107–34.

Ogaki, Masao, and Min-Seok Yang, 1998, "Structural Error Correction Models: Instrumental Variables Methods and an Application to an Exchange Rate Model" (unpublished; Columbus, Ohio: Ohio State University).

Park, Joon Y., 1990, "Testing for Unit Roots and Cointegration by Variable Addition," *Advances in Econometrics,* Vol. 8, pp. 107–33.

———, 1992, "Canonical Cointegration Regressions," *Econometrica,* Vol. 60 (January), pp. 119–43.

Park, Joon Y., and Buhmsoo Choi, 1988, "A New Approach to Testing for a Unit Root," CAE Working Paper No. 88–23 (Ithaca, New York: Cornell University).

Park, Joon Y., and Masao Ogaki, 1991, "Inference in Cointegrated Models Using VAR Prewhitening to Estimate Shortrun Dynamics," RCER Working Paper No. 280 (Rochester, New York: University of Rochester).

Phillips, Peter C.B., and Pierre Perron, 1988, "Testing for a Unit Root in Time Series Regression." *Biometrica,* Vol. 75, pp. 335–46.

Rogoff, Kenneth, 1995, "What Remains of Purchasing Power Parity?", Starr Center Economic Research Report 95/07 (New York: New York University).

Sachs, Jeffrey, 1980, "Wages, Flexible Exchange Rates, and Macroeconomic Policy," *Quarterly Journal of Economics,* Vol. 94 (June), pp. 731–47.

Said, Said E., and David A. Dickey, 1984, "Testing for Unit Roots in Autoregressive Moving Average Models of Unknown Order," *Biometrica,* Vol. 71, pp. 599–607.

Urbain, Jean-Pierre, 1992, "On Weak Exogeneity in Error Correction Models," *Oxford Bulletin of Economics and Statistics,* Vol. 54 (May), pp. 187–207.