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Stability of the Demand for M1 and Harmonized M3 in Finland

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Abstract

We derive a theoretical model for the demand for money using the money-in-the-utility-function approach. The steady-state – utility function – parameters of the model of narrow money (M1) estimated with cointegration techniques are stable over the foreign exchange rate regime shift; whereas in the model of harmonized M3 (M3H) they are *not* stable. The theoretical model fits the M1 data. The adjustment cost parameters of the M1 model describing the dynamics of the demand for money are stable over the sample period. The adjustment cost parameters of the M3H model are not stable. These results suggest that from the Finnish point of view M1 would be a more appropriate intermediate target for monetary policy than harmonized M3.

Keywords: money-in-the-utility-function model – structural breaks – demand for money – narrow money – harmonized M3

JEL classification: C22, C52, E41

Tiivistelmä

Tutkimuksessa selvitetään ovatko suppean rahan (M1) ja harmonisoidun lavean rahan (M3H) kysyntä olleet vakaita Suomessa 1980-luvun alusta lähtien. Rahan kysynnän teoreettinen malli perustuu raha hyötyfunktiossa -lähestymistapaan, jossa taloudenpitäjä maksimoi nykyhetkeen diskontattua odotettua hyötyä, jota hän voi saada kulutuksesta ja rahan hallussapidosta. Mallin pitkän aikavälin tasapainoon liittyvät parametrit estimoidaan käyttäen yhteisintegroituvuusmenetelmää. Tarkasteluissa käy ilmi, että suppean rahan kysynnän pitkän aikavälin tasapainoon liittyvät parametrit ovat vakaita yli otosperiodin kun taas harmonisoidun M3:n vastaavat parametrit vaihtelevat selvästi ajassa. Suppean rahan kysynnän ajassa tapahtuvaan sopeutumiseen liittyvät parametrit ovat myös vakaat. Harmonisoidun M3:n parametrit eivät puolestaan ole vakaat. Näiden tulosten valossa M1 saattaisi Suomen kannalta soveltua paremmin Euroopan keskuspankin välitavoitteeksi kuin harmonisoitu M3H.

Asiasanat: raha hyötyfunktiossa, rakennemuutos, rahan kysyntä, suppea raha, harmonisoitu M3

JEL luokitus: C22, C52, E41

Contents

1 Introduction

As a result of the European currency turmoil in autumn 1992 and again in 1993 and 1995, several European currencies are now floating. The Finnish markka was floated as a consequence of the September 1992 disturbance. In February 1993 the Bank of Finland formalized an explicit target for the inflation rate. In relying on an inflation target instead of an intermediate target, such as a monetary aggregate, Finland followed the earlier examples of New Zealand, Canada, the United Kingdom and Sweden. To increase the transparency of the inflation target, the Bank of Finland announced a list of indicators that it monitors in an effect to anticipate forthcoming inflation.

Articulation of the stance of monetary policy and of the measures applied is based on the evaluation of inflation prospects by means of a number of macroeconomic variables, ie inflation indicators. Money supply is argued to be one of the most important indicators used in the conduct of Finnish monetary policy. Stability of the demand for money is an important condition for the money measure to serve as an inflation indicator or intermediate target.

One possible choice for the monetary policy strategy of the planned European Central Bank (ECB) is to use money as an intermediate target. The stability of different money measures could vary across European countries. The main candidates under investigation are narrow liquid money (M1) and harmonized broad money (M3H)[1]. From the perspective of the monetary union, it is important to find a money measure whose demand is stable in all the participating countries and for which the national money demand parameters are as close as possible to average union values.

The aim of this study is to analyse the stability of the demand for these two money measures. In section 2, we derive the demand for money from the money-in-the-utility-function approach. To estimate the parameters in the presence of integrated variables, we log-linearize the first-order condition. The first-order condition is then linear in the levels of the variables but nonlinear in the parameters of the differenced variables. The steady-state part of the first-order condition can be estimated with cointegration techniques and the other parameters with the generalized method of moments estimator for given estimates of steady-state. The econometrics is overviewed in section 3 and the estimates[2] are reported and the stability evaluated in section 4. The final section concludes and discusses the policy implications of the empirical results.

[1]The European Monetary Institute (EMI) is still working on harmonization rules for various money measures. The present measure of M3H will probably not be the final measure.

[2]The cointegration estimation is done with CATS in RATS by Hansen and Juselius (1995) except for the small sample simulations, which are performed with Gauss utilizing the CIA code by Paolo Paruolo. The GMM estimation is done with Gauss, part of the code is based on the Hansen/Heaton/Ogaki GMM package by Ogaki (1993). I thank Paolo Paruolo and Masao Ogaki letting me to use their code.

2 Theoretical Background: Money-in-the-Utility-Function Model

The theoretical models of the demand for money give us a tool by which to discuss and interpret the estimation results. In the following, we present a standard demand for money model by utilizing the money-in-the-utility-function[3] (MIUF) approach.

We start with an MIUF model in which the *household* optimizes the discounted sum of expected utility from consumption and money:

$$\max E_0 \sum_{t=0}^{\infty} \delta^t \left(u(C_t) + \zeta v \left(\frac{M_t}{P_t} \right) \right). \tag{1}$$

The household allocates its real income y and other earnings among consumption goods (C_t: real value of consumption), bonds (B_t: real value of bonds denominated in units of time t consumption) which pay a gross real return $1 + r_t$ (from time t to time $t+1$), and and real money balances $\frac{M_t}{P_t}$ which pay a gross return $\frac{P_{t+1}}{P_t}$; for some definitions of money, money also pays a nominal return (own-yield of money) $1 + o_t \equiv O_t$. Whenever it adjusts its money balances between period $t-1$ and period t, the household suffers losses (in real terms) $a(M_t, M_{t-1}, M_{t-2})/P_t$. Thus the household's budget constraint is

$$C_t + B_t + \frac{M_t}{P_t} + \frac{a(M_t, M_{t-1}, M_{t-2})}{P_t} \leq y + \frac{O_{t-1} M_{t-1}}{P_t} + (1 + r_{t-1}) B_{t-1}. \tag{2}$$

The household's optimization problem (1) subject to (2) can be written as

$$\begin{aligned} \max E_0 \sum_{t=0}^{\infty} \delta^t \{ u [y + \frac{O_{t-1} M_{t-1}}{P_t} + (1 + r_{t-1}) B_{t-1} - B_t - \frac{M_t}{P_t} \\ - \frac{a(M_t, M_{t-1}, M_{t-2})}{P_t}] + \zeta v \left(\frac{M_t}{P_t} \right) \} \end{aligned} \tag{3}$$

The first-order conditions are

$$\delta E_t u'(C_{t+1}) = \frac{u'(C_t)}{1 + r_t} \tag{4}$$

$$\begin{aligned} 1 + a'_{M_t}(M_t, M_{t-1}, M_{t-2}) = \delta E_t \left\{ \frac{P_t}{P_{t+1}} \frac{u'(C_{t+1})}{u'(C_t)} \left[O_t - a'_{M_t}(M_{t+1}, M_t, M_{t-1}) \right] \right\} \\ + \zeta \frac{v'\left(\frac{M_t}{P_t} \right)}{u'(C_t)} \end{aligned} \tag{5}$$

We *assume* that

$$\operatorname{cov} \left(\frac{P_t}{P_{t+1}}, \frac{u'(C_{t+1})}{u'(C_t)} \right) = 0 \quad \text{and}$$

$$\operatorname{cov} \left(\frac{u'(C_{t+1})}{u'(C_t)}, \left(O_t - a'_{M_t}(M_{t+1}, M_t, M_{t-1}) \right) \right) = 0$$

[3]The empirical papers papers based on a similar approach are Poterba and Rotemberg (1987), Sill (1995) and Lucas (1988), who use the cash-in-advance constraint in their models.

and that the Fisher parity holds, ie $\frac{1}{I_t} \equiv \frac{1}{1+i_t} = \frac{1}{1+r_t} E_t \frac{P_t}{P_{t+1}}$. Then using (4), the condition (5) can be written as

$$\begin{aligned} a'_{M_t}(M_t, M_{t-1}, M_{t-2}) = & -\frac{1}{I_t} E_t a'_{M_t}(M_{t+1}, M_t, M_{t-1}) \\ & -1 + \frac{O_t}{I_t} + \zeta \frac{v'(\frac{M_t}{P_t})}{u'(C_t)}. \end{aligned} \tag{6}$$

Next we parametrize the utility function to the CRRA form as follows

$$u(C_t) = \begin{cases} \frac{1}{1-\rho} C_t^{1-\rho} & \text{if } \rho \neq 1 \\ \log C_t & \text{if } \rho = 1 \end{cases}$$

$$v\left(\frac{M_t}{P_t}\right) = \begin{cases} \frac{1}{1-\omega}\left(\frac{M_t}{P_t}\right)^{1-\omega} & \text{if } \omega \neq 1 \\ \log\left(\frac{M_t}{P_t}\right) & \text{if } \omega = 1. \end{cases}$$

From (6) one obtains

$$\begin{aligned} a'_{M_t}(M_t, M_{t-1}, M_{t-2}) = & -\frac{1}{I_t} E_t a'_{M_t}(M_{t+1}, M_t, M_{t-1}) \\ & -1 + \frac{O_t}{I_t} + \zeta C_t^{\rho} \left(\frac{M_t}{P_t}\right)^{-\omega}. \end{aligned} \tag{7}$$

It is standard practice to estimate such first-order conditions with generalized method of moments (GMM) estimators. However, what is sometimes overlooked — typically in the studies of the early 1980s — is the problem of non-stationarity. Stationarity of stochastic processes is the key assumption of GMM. If that is rejected, as is usually the case for macroeconomic time series, one should use other estimators, which unfortunately exist for linear models only. Thus, it is necessary to linearize the first-order conditions. We use the first-order Taylor approximation around the steady-state. In the steady-state, the stochastic processes should have finite variance, which is not the case if any of the variables in the model are $I(1)$. It is, however, possible that a linear combination of $I(1)$ variables is stationary. If so, the variables are cointegrated. We think that the linearized version of the steady-state solution of the model should represent the stationary linear combination of the variables. This would make it possible to linearize this model also.

In order to log-linearize equation (7), we parametrize the adjustment cost function $a(\cdot)$ as follows:

$$a(M_t, M_{t-1}, M_{t-2}) = \frac{\kappa}{2}\left[(M_t - M_{t-1}) - \nu(M_{t-1} - M_{t-2})\right]^2, \tag{8}$$

where κ and ν are adjustment cost parameters. The adjustment cost function expresses the notion that it is differences in the growth rate of money that affects costs, not the growth rate itself as is typical. However, if the parameter ν is zero, the adjustment cost function is the typical one. First we seek the stationary equilibrium for equation (7) and then use the first-order Taylor approximation around the stationary equilibrium. For the stationary equilibrium, the adjustment costs are zero and $C_t = C$, $I_t = I$, $I_t = I$, $M_t = M$

($\forall t \geq 0$). We denote logarithmic variables in the lower case (eg $\log C \equiv c$) and $I \equiv 1 + i$ and $O \equiv 1 + o$. In the stationary equilibrium, equation (7) reduces to

$$1 - \frac{O}{I} = \zeta C^{\rho} \left(\frac{M}{P} \right)^{-\omega}, \tag{9}$$

which is like the standard demand for money function. The parameter ζ cannot be identified since the chosen scale of C_t influences the estimate of ζ. From the first-order Taylor approximation around the (log of) stationary equilibrium, we obtain the following log-linear Euler equation:

$$\begin{aligned} \Delta m_t = & \frac{1}{I + \nu + \nu^2} \Bigg[(1 + \nu) E_t \Delta m_{t+1} + I \nu \Delta m_{t-1} \\ & + \frac{i - o}{\kappa M} \left((\omega m - \omega p - \rho c + O) - \omega m_t + \omega p_t + \rho c_t - \frac{O}{i - o} (i_t - o_t) \right) \Bigg], \end{aligned} \tag{10}$$

where the constant term $(\omega m - \omega p - \rho c + O)$ can be written as $-\log\left(\frac{I-O}{I}\right) - O - \log(\zeta)$ using information on the steady state. There is no separate constant term in the equation. The observed growth in the money stock should depend on the behaviour of the forcing variables, ie on the marginal processes. Equation (10) can also be written in the following form:

$$\begin{aligned} \Delta m_t = & \frac{1}{I + \nu + \nu^2} \Bigg[(1 + \nu) E_t \Delta m_{t+1} + I \nu \Delta m_{t-1} \\ & + \frac{O}{\kappa M} \left[(i - o) - (i_t - o_t) \right] \\ & + \frac{i - o}{\kappa M} \left((\omega m - \omega p - \rho c) - \omega m_t + \omega p_t + \rho c_t \right) \Bigg]. \end{aligned} \tag{11}$$

Using a number of assumptions and approximations, we arrive at a log-linear Euler equation which can be estimated. In order to describe the dynamics of the demand for money, we assume the existence of adjustment costs. This is not so innocent since money should be the asset which is cheapest to adjust. The specified adjustment cost function is quadratic for *changes* in money stock. One can obtain the standard quadratic adjustment cost function simply using the restriction $\nu = 0$. In such a case, the lagged log-change term Δm_{t-1} drops out of the linearized equation (10). The adjustment costs could be interpreted to proxy or describe the payment technology. Finally, probably the most crucial assumptions of our model concern the covariance restrictions. The first covariance assumption states that inflation and the intertemporal rate of substitution of consumption should be variation free. This can be interpreted as the money neutrality hypothesis. The second covariance assumption states that the intertemporal rate of substitution of consumption and the difference between own-yield of money and marginal adjustment costs should be variation free.

In the case of integrated (of order one) variables, the parametrization (11) suggests that there should be two cointegration vectors in the system $z_t = [m_t, p_t, c_t, i_t, o_t]'$. The first is the net opportunity cost of money $i_t - o_t$ and the second one is 'adjusted' velocity $m_t - p_t - \frac{\varrho}{\omega}c_t$. These lead to the testable hypothesis in the statistical model.

3 Econometric Setup

The econometric methods are briefly described in the following sections. The pricipal tools used in the following statistical analysis of the demand for money are the Euler equation estimation by GMM and cointegration analysis in the FIML framework of Johansen (1991). Under the assumption of non-stationary variables, the theoretical model yields restrictions on the cointegration vectors. Given the estimated cointegration vectors, the estimation of the rest of the parameters of the Euler equation (11) relies on the GMM approach of Hansen (1982).

To illustrate how the estimation can be performed in two steps, we write equation (11) in the following form

$$\Delta m_t = \frac{1}{I + \nu + \nu^2}\left[(1+\nu)E_t\Delta m_{t+1} + I\nu\Delta m_{t-1} + \gamma\beta' z_t\right]$$

where $\gamma \equiv [\frac{O}{\kappa M} \quad \frac{\omega(i-o)}{\kappa M}]'$ and

$$\beta = \begin{bmatrix} 0 & -1 \\ 0 & 1 \\ 0 & \frac{\varrho}{\omega} \\ 1 & 0 \\ -1 & 0 \end{bmatrix} \equiv [\beta_1 \ \beta_2] \quad \text{and} \quad z_t = \begin{bmatrix} m_t \\ p_t \\ c_t \\ i_t \\ o_t \end{bmatrix}.$$

If the variables in z_t are integrated order one, $I(1)$, the model can be interpreted as a sort of forward-looking error correction model, where β represents the cointegration vectors and the rest of the parameters come from the short-run dynamics. Our theoretical model can be reparametrized to indicate at most five independent cointegration vectors. The simplest approach is that of equation (11). If one considers the case of a single cointegration vector (as in parametrization (10)), one makes the implicit assumption that all the forcing variables $[p_t, c_t, i_t, o_t]'$ are integrated of order one ($\sim I(1)$). This leads to a situation wherein one cannot estimate the parameters of the system by GMM, which assumes stationarity of the stochastic processes. According to Dolado, Galbraith and Banerjee (1991), if the forcing variables are integrated of order d ($\sim I(d)$) the endogenous variable m_t is also integrated of the same order. In the case of quadratic adjustment costs, they propose a two-step estimation procedure:

1. The parameters in β can be estimated using the FIML of Johansen (1991), which will be described in the following section. Since the pa-

rameters of the cointegration vectors β are superconsistent one can treat the estimates of $\hat{\beta}$ as fixed in the second stage[4].

2. In the second step all the variables of the model are stationary. In this case, one can estimate the rest of the parameters by GMM.

The next section summarizes FIML estimation of the cointegration vectors and the following section discusses the GMM estimation of the model with special emphasis on the stability of the parameters.

3.1 Johansen's VAR model

We present the FIML estimation within the VAR of the cointegration relations and methods for testing the long-run structural hypothesis. The presentation is based on the papers of Johansen (1988), Johansen (1991) and Johansen and Juselius (1990). The p-dimensional VAR process in levels, $A(L)z_t = \varepsilon_t$ ($\varepsilon_t \sim \mathrm{NID}(0, \Sigma)$), can be written in the difference form

$$\Delta z_t = \Pi z_{t-1} + \Gamma_1 \Delta z_{t-1} + \cdots + \Gamma_{k-1}\Delta z_{t-k+1} + \mu + \Psi D_t + \varepsilon_t, \qquad t = 1, \ldots, T, \tag{12}$$

where $\Pi = -I_1 + \sum_{i=1}^{k} A_i$, $\Gamma_i = -(\sum_{j=i+1}^{k-1} A_j)$, μ is constant and D_t is the vector of deterministic variables. The Π matrix has a reduced rank in the case of cointegration ($\mathrm{rank}(\Pi) < p$). Any reduced rank matrix can be presented as a product of two full-rank matrices:

$$\Pi = \alpha\beta'. \tag{13}$$

Thus, Π is partitioned as α and β'. Matrix β, which defines the long-run relationships of the variables, is called the cointegration matrix. In cointegration, α is called the matrix of loadings of the equilibrium errors of the linear combinations defined by β.

Johansen (1988) and Johansen (1991) show that the ML estimator of the space spanned by β is the space spanned by r canonical variates reflecting the r largest squared canonical correlations between residuals of the least squares regressions of contemporaneous differences on lagged differences and levels on lagged differences.

It is important to note that one can estimate only the space spanned by β, not the individual cointegration vectors. One can give economic interpretation to the cointegration vectors only after identification.

Johansen (1988) derives two tests for testing the number of cointegration vectors. The null hypothesis of the *trace test* can be formulated

$$\mathcal{H}_0 : \mathrm{rank}(\Pi) \leq r \text{ or } \Pi = \alpha\beta',$$

[4]Superconsistency means that the parameters of the cointegration vectors converge much faster rate to the true values than do the parameters of, for example, ordinary least squares regression of stationary variables. Due to this fact, one is able to treat the parameters of the cointegration vectors as (asymptotically) fixed in the subsequent analysis of stationary variables.

where α and β are $p \times r$ matrices. The likelihood ratio test statistic for the hypothesis above is

$$-2\ln(Q) = -T \sum_{i=r+1}^{p} \ln(1-\hat{\lambda}_i^2),$$

where $\hat{\lambda}_{r+1}^2, \ldots, \hat{\lambda}_p^2$ are $p-r$ largest squared canonical correlations. The alternative hypothesis is that the number of cointegration vectors is larger than r. Osterwald-Lenum (1992) has simulated the critical values of these test statistics for $p = 12$. It has been shown in some simulation studies (Eitrheim 1991, Haug 1996, Toda 1995), that using the asymptotical tables might be misleading in small samples. For this study, we have simulated the model, under the null, in order to obtain empirical critical values for the trace tests.

3.2 GMM Estimation of the Euler Equation and Tests of Parameter Stability

Since the parameters $\hat{\beta}$, estimated by cointegration methods, are super-consistent, one can estimate the rest of the model parameters $\theta = [\nu\ I\ 0\ \omega\ \kappa M]'$ with the GMM of Hansen (1982) taking $\hat{\beta}$ as given. We define the 5-dimentional vector $w_t = [\Delta m_t, \Delta m_{t+1}, \Delta m_{t-1}, \hat{\beta}_1' z_t, \hat{\beta}_2' z_t]'$, the parameter vector as

$$\tilde{\theta} = \frac{1}{I+\nu+\nu^2} \begin{bmatrix} I+\nu+\nu^2 \\ 1+\nu \\ I\nu \\ O/\kappa M \\ \omega(I-O)/\kappa M \end{bmatrix}$$

and d as the vector of coefficients of the deterministic variables[5] (D_t) of the empirical model. The total number of parameters is $j = \dim(d) + \dim(\tilde{\theta}) \equiv \dim(\Theta)$, where Θ is the whole parameter space.

Given the instruments[6] set x_t (l-dimensional vector; see again table 3) we define the orthogonality conditions — implied by the Euler equation (11) — as

$$h(\Theta, w_t) = \begin{bmatrix} w_t - d'D_t \\ \left[\tilde{\theta}'(w_t - d'D_t)\right] x_t \end{bmatrix}, \tag{14}$$

where $h(\Theta, w_t)$ is a $b \times 1$ vector-valued function. Let Θ^* denote the true value of Θ such that $E(h(\Theta^*, w_t))=0$, $Y_T \equiv [w_1, \ldots, w_T]$ and $g(\Theta, Y_T) \equiv \frac{1}{T}\sum_{t=1}^{T} h(\Theta, w_t)$. The idea behind GMM is to choose θ so as to make the

[5]The deterministic variables are the constant, centred seasonal dummies and a set of other dummies listed in table 3 and described in the appendix.

[6]Instruments should be chosen so as to correlate as much as possible with Δm_{t+1} but not with the forecast error. The lagged error correction terms, for example, typically contain much information on the endogenous variables involved.

sample moment $g(\Theta, Y_T)$ as close as possible to the population moment. Thus, the scalar

$$Q(\Theta, Y_T) = [g(\Theta, Y_T)]' \, W_T \, [g(\Theta, Y_T)]$$

is to be minimized. W_T is the positive definite weighting matrix, which may be a function of data matrix Y_T. Hansen (1982) shows that the optimal weighting matrix is $\hat{W}_T = \hat{S}^{-1}$, where $\hat{S} = \frac{1}{T}\sum_{t=1}^{T}[h(\hat{\Theta}, w_t)][h(\hat{\Theta}_0, w_t)]'$ if $h(\hat{\Theta}^*, w_t)$ is serially uncorrelated.

When the number of orthogonality conditions b exceeds the number of parameters j, the model is overidentified. Hansen (1982) shows that under certain regularity conditions it is possible to test the overidentification restrictions, since

$$\left[\sqrt{T} g(\hat{\Theta}, Y_T)\right]' \hat{W}_T \left[\sqrt{T} g(\hat{\Theta}, Y_T)\right] \xrightarrow{d} \chi^2(b-j),$$

where $\xrightarrow{d}$ means convergence in distribution.

Due to the financial deregulation and a change in foreign exchange rate regime, we test for the stability[7] of the parameters. We have the advantage that we know the breakpoints of the foreign exchange rate regime and financial deregulation[8]. The total full sample size is T. Let T_0 denote the possible break point, $Y_{T_0} \equiv [w_1, \ldots, w_{T_0}]$, $Y_{T-T_0} \equiv [w_{T_0+1}, \ldots, w_T]$, $g(\Theta_0, Y_{T_0}) \equiv \frac{1}{T_0}\sum_{t=1}^{T_0} h(\Theta_0, w_t)$, $g(\Theta_1, Y_{T-T_0}) \equiv \frac{1}{T-T_0}\sum_{t=T_0+1}^{T} h(\Theta_1, w_t)$ and Θ_0 and Θ_1 the first and second subsample parameters. One can consider, for example, August 1992 ($= T_0$) as a possible break point. According to Hamilton (1994), one approach is to use the first subsample to estimate $\hat{\Theta}_0$ by minimizing

$$Q(\Theta_0, Y_{T_0}) = [g(\Theta_0, Y_{T_0})]' \, \hat{S}_0^{-1} \, [g(\Theta_0, Y_{T_0})]\,,$$

where $\hat{S}_0 = \frac{1}{T_0}\sum_{t=1}^{T_0}[h(\hat{\Theta}_0, w_t)][h(\hat{\Theta}_0, w_t)]'$ if $h(\hat{\Theta}_0^*, w_t)$ is serially uncorrelated. Hansen (1982) shows that

$$\sqrt{T_0}\left(\hat{\Theta}_0 - \Theta_0^*\right) \xrightarrow{d} N(0, V_0).$$

$\hat{V}_0$ can be estimated from

$$\hat{V}_0 = \left(\hat{D}_0 \hat{S}_0^{-1} \hat{D}_0'\right)^{-1},$$

where

$$\hat{D}_0 \equiv \left. \frac{\partial g(\Theta_0, Y_{T_0})}{\partial \Theta_0'} \right|_{\Theta_0 = \hat{\Theta}_0}.$$

[7] Hamilton (1994) and Oliner, Rudebusch and Sichel (1996) survey structural stability tests using the GMM approach. See also Hoffman and Pagan (1989) and Dufour, Ghysels and Hall (1994).

[8] The fixed exchange rate regime collapsed on 8 September, 1992.

One also computes the analogous measures for the second subsample $T - T_0$. Denoting $\pi \equiv \frac{T_0}{T-T_0}$ one can summarize the convergences as

$$\sqrt{T}\left(\hat{\Theta}_0 - \Theta_0^*\right) \xrightarrow{d} N(0, V_0/\pi) \quad \text{and}$$
$$\sqrt{T}\left(\hat{\Theta}_1 - \Theta_1^*\right) \xrightarrow{d} N\left(0, V_1/(1-\pi)\right).$$

Andrews and Fair (1988) suggest the test statistic

$$\mathrm{AF} = T(\hat{\Theta}_0 - \hat{\Theta}_1)' \left\{\pi^{-1}\hat{V}_0 + (1-\pi)^{-1}\hat{V}_1\right\}^{-1} (\hat{\Theta}_0 - \hat{\Theta}_1) \tag{15}$$

to test the null hypothesis $\Theta_0 = \Theta_1$. The test statistic $\mathrm{AF}_T \xrightarrow{d} \chi^2(j)$. If one does not know the date of the possible structural break, one can repeat the test for different choices of T_0 and choose the largest value of the test statistic. Andrews (1993) derives the asymptotic distribution of such a test. The test setup entails the limitation that each of the subsamples sizes should approach infinity. This is also a drawback of the Ghysels and Hall (1990) setup.

Ghysels and Hall (1990) propose a test whereby they estimate the model using the first subsample and then examine whether the orthogonality conditions of the model are satisfied over the second subsample using the parameter estimates obtained from the first subsample. The null and alternative hypotheses for the test are

$$\begin{aligned} H_0 &: E\left(h(\Theta^*, w_t)\right) = 0,\ t = 1, \ldots, T_0 \quad \text{and} \\ &\quad E\left(h(\Theta^*, w_t)\right) = 0,\ t = T_0 + 1, \ldots, T \\ H_1 &: E\left(h(\Theta^*, w_t)\right) = 0,\ t = 1, \ldots, T_0 \quad \text{and} \\ &\quad E\left(h(\Theta^*, w_t)\right) \neq 0,\ t = T_0 + 1, \ldots, T. \end{aligned}$$

The test statistic is defined as

$$\mathrm{GH} = (T - T_0)\left[g(\Theta_0, Y_{T-T_0})\right]' \left(\hat{V}_1\right)^{-1} \left[g(\Theta_0, Y_{T-T_0})\right],$$

where

$$\begin{aligned} \hat{V}_1 &= S_1 + (\pi)^{-1}\tilde{D}_1\left(D_0'S_0D_0\right)^{-1}\tilde{D}_1' & & \\ S_0 &= T_0 g(\Theta, Y_{T_0}) g(\Theta, Y_{T_0})' & S_1 &= (T - T_0) g(\Theta, Y_{T-T_0}) g(\Theta, Y_{T-T_0})' \\ D_0 &= \frac{\partial g(\Theta, Y_{T_0})}{\partial \Theta'} & \tilde{D}_1 &= \frac{\partial g(\Theta_1, Y_{T-T_0})}{\partial \Theta'}. \end{aligned}$$

The test statistic is $\chi^2(b)$ distributed. Oliner et al. (1996) study different choices of weighting matrix V_1. One candidate is the full sample estimate.

4 Estimation Results

In the following two subsections, we present the empirical results for estimation of the parameters of the theoretical model. First we estimate the steady-state part of the theoretical model. Parameters in the steady-state part of the model reflect the parameters of the utility function[9]. That is, we test for cointegration and estimate the restricted[10] cointegration space β implied by the theoretical model. In order to evaluate the stability of the utility function parameters, we test recursively whether the estimated, full-sample cointegration space lies within the space estimated recursively for the period 1985–1995.

We proceed with the given cointegration vectors (estimated from the full sample) and estimate the rest of the parameters of the Euler equation (11). The rest of the parameters in the Euler equation are related to the adjustment cost function. We also test for the stability of these parameters.

We repeat the analysis for both M1 and M3H. The M1 system does not contain the own-yield of money. In the theoretical model, this means that we have restricted $O = 1$ ($o = 0$). The set of deterministic dummy variables differs as between the M1 and M3H models (see table 3). The M3H system is augmented with the dummy MFREST, which is unity for the pre-1987 period, during which the Ministry of Finance restricted banks' CD issues and the Bank of Finland did not use CDs in its open market operations, and zero otherwise. That dummy enters into the cointegration space and is restricted to enter only into the cointegration relations between own-yield and opportunity cost of money.

4.1 Estimates of Steady-State Parameters

We do not test the price homogeneity due to the possibility of I(2)ness of the data (Ripatti 1994). Instead, we impose the price homogeneity restriction[11] on the model by analyzing real money in the steady-state. The adjustment cost function in the theoretical model is parametrized to allow lag length three; $k = 3$ in equation (12). This lag length is long enough to yield zero residual autocorrelations. The vector error correction model is augmented with the centred seasonal dummies and with the set of intervention dummies. These are listed in table 3.

Table (1) reports the trace tests for cointegration rank. According to the trace test and reported 95 per cent *empirical*[12] fractiles, there exists one cointegration vector in the M1 system, as is predicted by the theory. The empirical significance level for the null of no cointegration is less than 0.01.

The determination of the cointegration rank of the M3H system is more problematic. The theoretical model can be parametrized to allow up to five cointegration vectors. The difference between empirical and asymptotical critical values is quite small[13]. Comparison of the trace tests with the empirical critical values indicates that the cointegration rank is one. If we include the dummy variable MFREST in the cointegration space, the trace test value for

[9] The scale elasticity is ρ/ω; In the M1 model, the opportunity cost semi-elasticity is $\frac{1}{\omega i}$.

[10] We test for the restrictions implied by the theoretical model.

[11] Note, however, that we introduce this price homogeneity also into the short-run dynamics. Ripatti (1994) cannot reject long-run price homogeneity in the I(2) system.

[12] The empirical fractiles of the trace test are based on 10 000 replications under the null.

[13] The asymptotical critical values are obtained from Johansen (1995), table 15.3. One should note that the asymptotical critical values are not the correct ones since we have a set of noncentred dummies in the model. However, they are the ones that are given by econometric software packages such as PC-FIML or CATS in RATS.

$r = 0$ is 57.91 while the *asymptotical* 95 per cent fractile is 55.67. For the null hypothesis $r \leq 1$ the trace test value is 21.22 and the the *asymptotical* 95 per cent fractile is 35.71 (the significance level is approximately 0.5). This leads to the conclusion that the cointegration rank is one.

Table 1: Trace Tests of Cointegration Rank

M1[a]					M3H[b]			
λ	Trace test	95 % fractile[c]	95 % asymptotic fractile[d]	H_0	λ	Trace test	95 % fractile	95 % asymptotic fractile
0.186	42.16	31.22	29.38	$r = 0$	0.173	53.81	51.55	47.21
0.013	3.19	15.51	15.34	$r \leq 1$	0.072	17.93	31.63	29.38
0.003	0.64	4.41	3.84	$r \leq 2$	0.016	3.73	17.53	15.34
				$r \leq 3$	0.004	0.73	5.14	3.84

[a]Since we have no measure of the own-yield of money, the dimension of the M1 model is three instead of four.

[b]The dummy MFREST has been included in the deterministic part of the M3H model in the estimation and the simulation of the test statistic.

[c]To obtain empirical distributions for the tests, the trace tests have been calculated, under the null, for 10 000 replications.

[d]The asymptotic fractiles are from Johansen (1995) and are not the correct ones since we include some noncentred dummies in the system.

The normality of residuals is violated in the interest rate equations (table 2). This is due to the excess kurtosis. The autocorrelation figures, which are not reported here, show no residual autocorrelation.

Next we test for the restrictions on the β-space implied by the Euler equation (11). For the M1 model, there are no restrictions in the cointegration space. However, we test for the unit scale elasticity since the free estimate is very close to one (0.95). The restriction is not rejected (with p-value= 0.49). The restricted cointegration vector is

$$\hat{\beta}'_{\mathrm{M1}} z_t = \left[\ (m-p)_t - y_t + 1.807 i_t \ \right].$$

The results contradict the results of Ripatti (1994), where the estimated scale elasticity was significantly below one and the interest rate semi-elasticity only sightly above one. The unit scale elasticity implies that the risk aversion parameters in the utility function are equal, ie $\hat{\rho} = \hat{\omega}$. The recursive estimates of the scale elasticity and opportunity cost semi-elasicity are in figure 1. The graphs indicate that the parameters have been fairly stable during the past ten years. However, the scale elasticity has slightly increased during the 1990s which explains the differences in Ripatti (1994) and this study. The left panel of figure 2 clearly supports the judgement that the parameter estimates of the steady-state are stable[14].

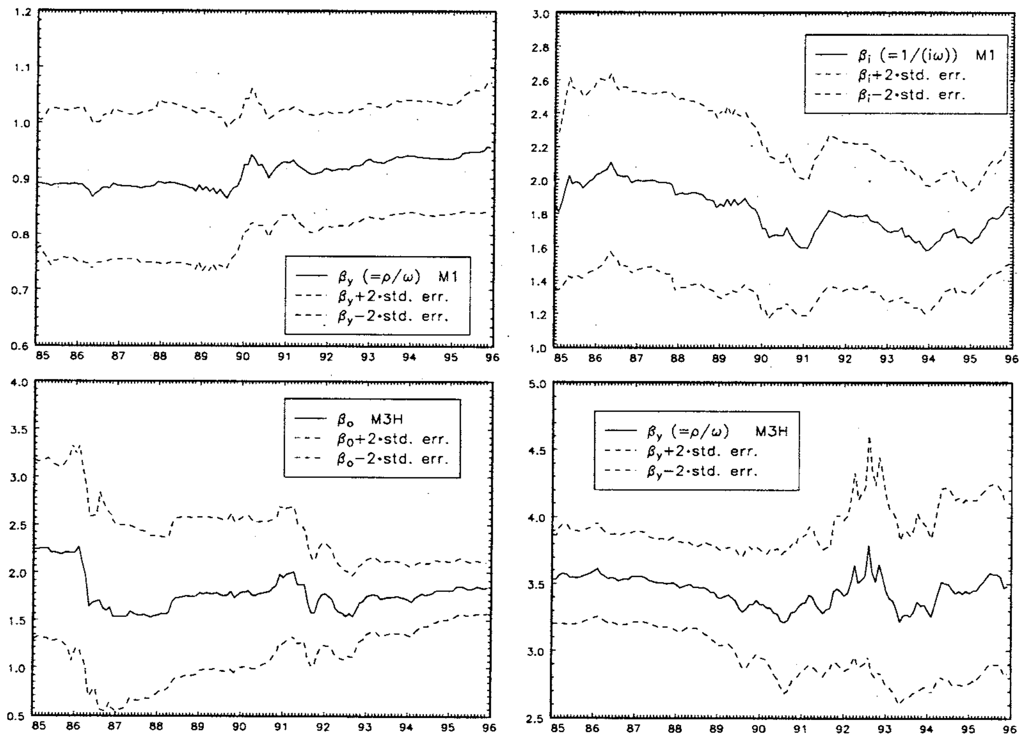
[14]Hansen and Juselius (1995) provides an attractive way to test the stability of the pa-

Table 2: Residual Diagnostics.

M1[a]				M3H		
ARCH(3)	Norm.	R^2	Equation	ARCH(3)	Norm.	R^2
0.03	3.12	0.77	$\Delta(m-p)_t$	2.16	2.09	0.34
9.81	0.04	0.95	Δy_t	10.60	6.45	0.95
1.11	47.54	0.36	Δi_t	6.88	80.27	0.56
			Δo_t	0.59	87.45	0.31
Multivariate LM test for the fourth order residual autocorrelation: p-value= 0.93				Multivariate LM test for the fourth order residual autocorrelation: p-value= 0.25		

[a]The test statistics ARCH(3) for no ARCH of the third degree is $\chi^2(3)$ distributed; Jarque–Bera normality (as null) test statistics is $\chi^2(2)$ distributed. The cointegration space is restricted before computation of residual diagnostics.

Figure 1: Recursive Estimates of the Scale Elasticity (β_y) and the Opportunity Cost Semi-elasticity (β_i) of M1 and the Own-Yield Semi-elasticity (β_o) and the Scale Elasticity (β_y) of M3H



a

[a]Dashed lines 95 % confidence intervals.

In the estimation of the M3H model with no restrictions, the first cointegration vector might be interpreted as the spread between the opportunity cost and the own-yield of money and the second cointegration vector (possibly non-stationary) as the velocity equation with scale elasticity greater than unity. The dummy variable MFREST is restricted to the first cointegration vector. The coefficients of real money and GDP do not differ significantly from zero in the first cointegration vector. According to the theoretical model, one should include the second cointegration vector in the analysis. The trace test does not indicate stationarity, even though it is assumed so in the following procedures.

I restrict the cointegration space in the following way (as implied by the theoretical model):

$$\hat{\beta}'_{\mathrm{M3H}} z_t = \begin{bmatrix} 0 & +0 & +i_t & -o_t & -0.029\mathtt{MFREST} \\ (m-p)_t & -3.26y_t & +0 & +0 & +0 \end{bmatrix},$$

where the coefficients of MFREST in the first vector and y_t in the second vector are estimated freely. These restrictions are not supported by the data (p-value< 0.001). If we estimate the own-yield semi-elasticity freely (1.8 times the opportunity cost semi-elasticity[15]), the restrictions are not rejected (p-value$= 0.2$). If we assume (as the trace test indicates) that there is only one cointegration vector in the M3H model, the test results concerning the first cointegration vector are almost identical. The recursive test statistic for the hypothesis that the estimated full sample cointegration space lies withing the cointegration space for the sub-samples ending in 1985 onwards (figure 2) indicates serious instabilities during the 1990s part of the recursive period.

The coefficient of the dummy MFREST indicates that the banks' interest rate marginal was on the average three percentage points higher during the period of regulation of CD issues, before 1987. The deregulation significantly reduced the banks' margin by boosting the average cost of liabilities. The scale elasticity in the second cointegration vector is much too high to be reliable. It implies that the risk aversion measure of for real money is three times as large as for consumption.

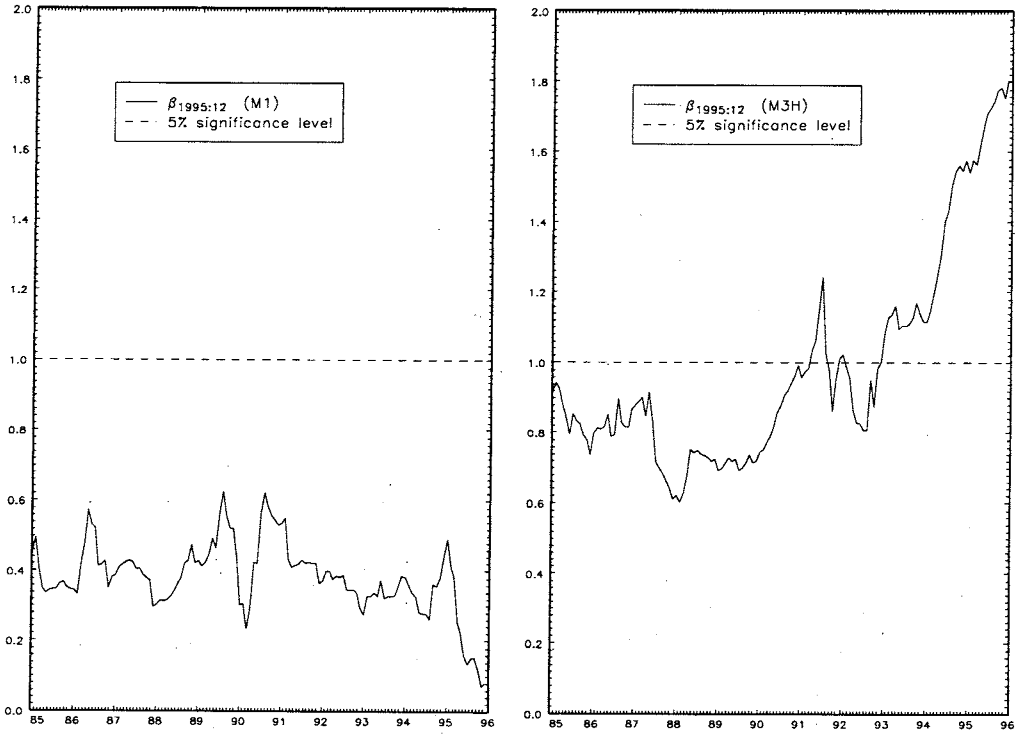
The recursive estimates of the M3H model further illustrate the problem. The own-yield elasticity varies between 1.5 and 2.3 during the recursive period (lower left panel of figure 1); the scale elasticity varies between 3.2 and 3.9 and the confidence interval actually *widens* during the recursive period.

Finally, we augment the M3H model with the variable that is the logarithmic difference between M3H and M1. If the coefficient of the variable in such a modification of the M3H model is unity, the model is a genuine M1 model and

rameters of the cointegration vector. We estimate the cointegration space using the full sample and test recursively whether the estimated sub-sample cointegration space (β_τ, $\tau = T_f + 1, \ldots, T$, where T_f is starting point of recursive testing) contains the full-sample cointegration space $\hat{\beta}_{T_0}$, ie $\mathcal{H}_{\beta_\tau}: \quad \hat{\beta}_{T_0} \in \mathrm{sp}(\beta_\tau), \quad \tau = T_f, \ldots, T$.

[15]We use the after-tax own-yield of money but the ordinary opportunity cost of money. The results do not differ when using the after-tax opportunity cost of money — the coefficient is 1.6. We have chosen to use the ordinary opportunity cost of money instead of the after-tax opportunity cost since firms have the possibility of subtracting interest income from taxes and they can also use foreign subsidiaries in order to avoid paying taxes on interest income.

Figure 2: Recursive Tests for $\hat{\beta}_{1995:12} \in \mathrm{sp}(\hat{\beta}_\tau)$, $\tau = 1985:1, \ldots, 1995:12$



[a]

[a]The 5% significance level scaled to unity. The coefficients of the pre-determined variables (dummies etc) and short-run dynamics are the full sample estimates computed before the recursive test.

the aggregation from M1 to M3H is not valid. The estimated cointegration space with the own-yield of money restricted to zero is as follows (standard errors in parentheses below the coefficients):

$$\hat{\beta}'[z_t' \quad (m3h_t - m1_t)]' =$$
$$\begin{bmatrix} (m3h-p)_t & -0.94y_t & +1.75i_t & +0 & -0.98(m3h_t - m1_t) \\ & (0.13) & (0.19) & & (0.05) \end{bmatrix}.$$

It is clear that the estimated model is the same as the M1 model. Evidently, this suggests that the theoretical model which is consistent with the M1 model is not consistent with M3H model[16].

We summarize this section the fact that the steady-state parameters, ie the utility function parameters of the M1 model, are stable. We can continue on to the estimation of the adjustment cost parameters of the M1 model, ie the dynamics of the M1 system. The parameters of the M3H model are neither stable nor of plausible size. Despite that fact, we continue with the M3H model too. We want to see if the Euler equation estimation provides information on the behaviour of M3H.

4.2 Estimates of Adjustment Cost Parameters

We proceed with the GMM estimation of the first-order conditions. We repeat the analysis for both money measures. The Euler equation in the M3H model is based on equation (11). There is no measure for the own-yield of M1. Thus, we have to modify the Euler equation (10) as follows:

$$\begin{aligned} \Delta m_t =& \frac{1}{I+\nu+\nu^2}\Bigg[i\omega\left(m-p-\frac{\rho}{\omega}c+\frac{1}{\omega i}i\right)+(1+\nu)E_t\Delta m_{t+1} \\ &+I\nu\Delta m_{t-1}-\frac{i\omega}{\kappa M}\left(m_t-p_t-\frac{\rho}{\omega}c_t+\frac{1}{\omega i}i_t\right)\Bigg], \end{aligned} \tag{16}$$

where $\hat{\beta}_1 \equiv \widehat{\rho/\omega} = 1$ and $\hat{\beta}_2 \equiv \widehat{1/(\omega i)} = 1.807$. In the previous section, we did not restrict the constant term to the cointegration space. Therefore, the mean of the error correction term $(m-p-\frac{\rho}{\omega}c+\frac{1}{\omega i}i)$ is non-zero. We will estimate the mean of the error correction term as the average of the error correction term.

We do not estimate the deterministic variables, such as seasonal, strike and other dummies, with the GMM, as suggested in the previous section;

[16] We have tried several other specifications of the M3H system. The deterministic trend in the cointegration space — restricted to the second cointegration vector — yields plausible parameters estimates for fixed exchange rate period. According to the test results, the velocity seems to be trend stationary. However, the forecasting performance of such a model is very unpleasant. The trend does not fit to the cointegration space at all during the floating exchange rate regime. We have also augmented the original variable set with some other variables which might capture the financial deregulation of 1980's and the broken trend in the decline of the velocity in 1990's. An example of this kind of variable is the stock of CDs issued by the banks and the Bank of Finland. The parameter estimates of such models are not plausible and those kinds of variables are not consistent with the theoretical model.

that would be computationally burdensome and would increase the number of instruments needed. However, before the GMM estimation we run two extra OLS regressions in which we condition on the variables listed in table 3. The instrument sets used in the GMM estimation are also listed in table 3.

Table 3: Deterministic Variables and Instruments

	M1	M3H
Δm_t and cointegration relations adjusted[a] for	`CGAINT`, `BSTRIKE1`, `BSTRIKE2`, `TRAF`, `DSPEC`, `REBATE`, `JULY`, `WTAX` and 11 centred seasonals	`CGAINT`, `BSTRIKE1`, `BSTRIKE2`, `TRAF`, `DSPEC`, `REBATE`, `JULY` and 11 centred seasonals
Instruments	Constant, Δm_{t-3}, Δp_{t-j}, Δy_{t-j}, Δi_{t-j} and $(m_{t-j}-p_{t-j}-\hat{\beta}_1 y_{t-j}+\hat{\beta}_2 i_{t-j})$ $(j=2,3)$	Constant, Δm_{t-3}, Δp_{t-j}, Δy_{t-j}, Δi_{t-j}, Δo_{t-j}, $(i_{t-j}-o_{t-j})$ and $(m_{t-j}-p_{t-j}-\hat{\beta}_1 y_{t-j})$ $(j=2,3)$

[a]These are the variables that are used in the separate regressions in order to condition on the seasonality and various tax and strike effects.

M3H model: The M3H model does not fulfil some of the key assumptions: the stationarity is violated by the error correction terms; the error term, which should reflect pure expectational error, contains significant first-order autocorrelation (-0.52). Thus, the parameter estimates of the M3H model might be nonsense (see table 4). None of the original parameters differ significantly from zero, although the point estimates are within the reasonable range[17]. The zero estimate of the adjustment cost parameter ν indicates that the adjustment cost function should be in the standard quadratic form. In that case, the coefficient of the lead term reduces to $1/(1+i)$. The estimates of risk aversion parameters of the utility function are in the typical range[18] of the consumption-based asset pricing models with cross country data (Braun et al. 1993, Roy 1995). The coefficient of the lead term is very close to unity; it does not differ significantly from one. This means that agents are discounting the very distant future. It is also an indication that M3H might be influenced by expected wealth rather the income. Due to the poor statistical properties of the M3H model, one cannot

[17]The i, ie the linearization point of the opportunity cost of money is intended to get negative value in the estimation.

[18]The risk aversion parameters of the capital asset pricing models are typically in the range 0.5–4. For example, the multicountry (Germany, Japan, USA) estimates of Roy (1995) are typically close to the lower bound of the range in the models in which the bond is the only asset. When the set of assets is augmented with stocks, the risk aversion parameter tends to get estimates between 2 and 6. Braun, Constantinides and Ferson (1993) extend the approach, relaxing the time separability of the utility function, ie to allow for habit persistence. Their point estimates for the risk aversion parameter for six large industrial countries vary between 0.35 (Japan) and 12 (Canada). Unfortunately, such studies have not been implemented with Finnish data.

rely on these results. The instability remains the robust feature of the M3H model.

M1 model: The adjustment cost parameter ν is greater and is significant in the M1 model, suggesting that lagged changes in money holdings influence the current change in money holdings. Contrary to the M3H model, the value of the linearization point of the opportunity cost of money is very large in the M1 model. Its standard error is however small. The level of the adjustment costs κM is much higher in the M3H model than in the M1 model. This is partly due to the fact that M3H is almost three times as large as M1; but the components of M3H also include illiquid time deposits and assets that do not have a secondary market. That might also rise the adjustment costs. The risk aversion parameters of the utility function are at the lower bound of the range of multi-country comparisons (see the footnote above). The estimate of the coefficient of the lead term is much smaller in the M1 model than in the M3H model. The value implicates that agents do not care about the distant future. Due to the greater absolute value of ν, the coefficient of the lagged money change is bigger in the M1 model than in the M3H model. The sign of the estimate of the coefficient of the error correction term is correct; the range is also feasible. The uncertainty of the estimate is fairly large in the first sub-sample. The second sub-sample estimate corresponds to the estimate of the error correction term in Ripatti (1994). The J-test indicates no violation of the orthogonality conditions.

In the estimation, we consider two possible structural breaks: the first is the change in the foreign exchange rate regime in September 1992 when the Finnish markka was floated. The second is the end of financial deregulation. The pre-1987 period is characterized by financial deregulation measures. The Bank of Finland started open market operations in March 1987 and the bank quotas for CD issues were abolished simultaneously. So we test for structural breaks at these time points.

The parameter estimates and the test statistic for the structural stability tests are presented in table 4. The parameter stability tests, described in section 3.2, clearly do not display any structural change. Also the general view given by figure 3 is encouraging especially for the samples ending in the 1990s. There is, however, a need to look closer at the parameter estimates of the sub-samples.

The last two columns of table 4 give parameter estimates from the financial deregulation period and the free capital markets period respectively. The estimate of ν in the latter part of the sample is significantly higher than in the first sub-sample. However, the level of adjustment costs is much lower in the period of free capital markets. This might reflect the advances[19] in payment technology for transactions accounts and in banking in general that

[19] In Finland, the number of automatic teller machines (ATM) per capita is among highest in the world. Also other electronic payment systems are very highly developed in Finland. The share of debit card payments and electronic funds transfers at point of sale (EFT-POS) is very high. Giro payments are the most important form of funds transfer. On the other hand, the shares of cheque and currency payments are very low. Cheques are presently used mainly in large-value payments.

Table 4: Parameter Estimates of the Euler Equations for M1 and M3H

Parameter[a]	M3H 1980:5 – 1995:12[b]	M1 1980:5 – 1995:12	M1 1980:5 – 1986:12[c]	M1 1987:1 – 1995:12[d]
$I = (1 + i)$	1.01	2.04	2.15	2.02
	(0.37)	(0.014)	(0.10)	(0.005)
ν	−0.11	−0.34	−0.18	−0.38
	(0.27)	(0.12)	(0.12)	(0.09)
κM	30.55	3.71	15.46	1.77
	(17.35)	(1.30)	(6.39)	(0.36)
ω	1.20	0.53	0.48	0.54
	(7.96)	(0.01)	(0.05)	(0.004)
ρ^e	3.91	0.53	0.48	0.54
	(2.23)	(0.01)	(0.05)	(0.004)
Coefficient of the lead term	0.987	0.36	0.40	0.35
	(0.08)	(0.02)	(0.19)	(0.01)
Coefficient of the lag term	−0.12	−0.38	−0.50	−0.43
	(0.07)	(0.04)	(0.05)	(0.03)
Coefficient of the error correction[f] term	−0.002	−0.08	−0.02	−0.18
	(83.80)	(0.09)	(1.79)	(0.02)
Coefficient of $i_t - o_t$	−0.03			
	(3.02)			
Significance level of the test for the overidentification restrictions	0.35	0.87	0.70	0.13
Significance level of the parameter stability tests				AF[g]: 0.18; GH[h]: 0.53

[a]Standard errors are in parentheses below the parameter value. The standard error of the "derived" parameters, ie parameters that are computed from the original free parameters, are based on linear approximation with respect to the original parameters of the model. However, they do not account for the uncertainty of the cointegration parameters.

[b]Full sample.

[c]Period of the financial deregulation.

[d]Period of free capital markets.

[e]In M1 system $\rho = \omega$ due to the unit scale elasticity.

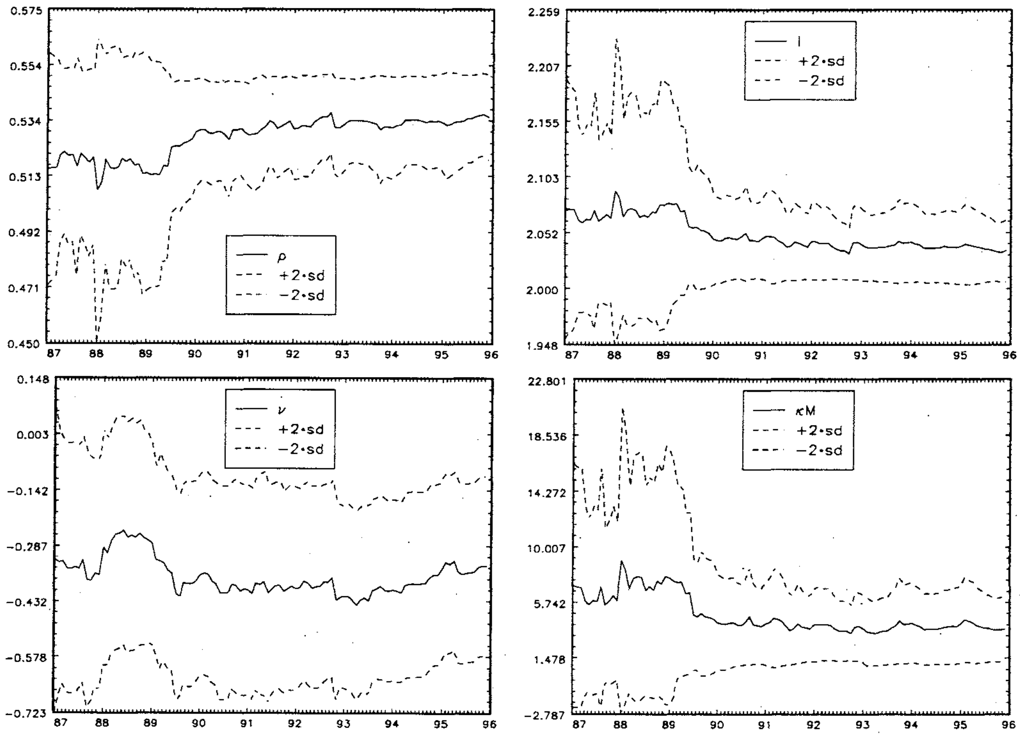
[f]For M1, this is the loading of the single cointegration vector, ie $m_t - p_t - \frac{\rho}{\omega} y_t + \frac{1}{\omega i} i_t$; for M3H, this is the coefficient of the "adjusted velocity", ie $m_t - p_t - \frac{\rho}{\omega} y_t$.

[g]Andrews and Fair (1988) test statistics.

[h]Ghysels and Hall (1990) test statistics, based on the weighting matrices of each sub-sample.

have occurred since the latter part of the 1980s. The linearization point of the opportunity cost of money, i, has essentially been constant over the two sub-periods. An interesting feature however is the significance level of the test of overindentification restrictions. It declines towards the end of the sample period. This might suggest that some instruments are no longer orthogonal (ie exogenous in a sense) with respect to the Euler equation. The evidence is more apparent if we restrict the sample to cover only the floating exchange rate regime[20]. Despite the fact that the null of parameter stability is not formally rejected, there might have been structural change in the system during that period. The low significance level of the J-test indicates that at least one of the instruments might be correlated with the residuals during the floating exchange rate regime.

Figure 3: Recursive Estimates of Some Parameters of the Euler Equation for M1



a

[a]The standard error of ρ is based on the linear approximation of ρ with respect to the original parameters of the model. It does not account for the uncertainty of the cointegration parameters.

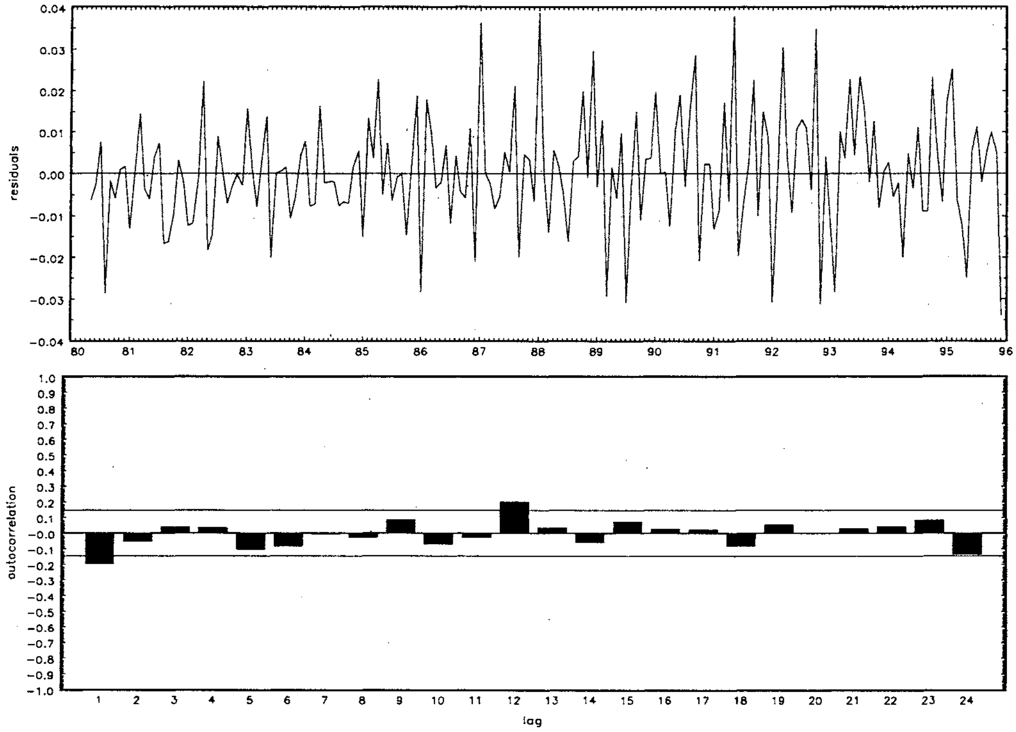
The full-sample estimate of the original, non-linearized Euler equation (7) for M1 augmented with the parametrized adjustment costs (8) is as follows:

$$\frac{3.71}{M}\left(\Delta M_t + 0.34\Delta M_{t-1}\right) = -\frac{2.45}{MI_t}\left(E_t\Delta M_{t+1} + 0.34\Delta M_t\right) - 1 + \frac{1}{I_t} + 0.59C_t^{0.53}\left(\frac{M_t}{P_t}\right)^{-0.53}.$$

[20]One should note that the sample size, 40, for the floating exchange rate regime is extremely small given the fact that the consistency of GMM estimators is based on large sample size. The numbers are not reported here.

The residuals, which reflect pure expectational error, are white noise (see figure 4).

Figure 4: Residuals and Residual Autocorrelation of the Euler Equation



5 Discussion

Starting from the dynamic money-in-the-utility-function model and assuming adjustment costs of changing money holdings, we derived the first-order condition describing the demand for money. We had to assume some covariance conditions: the first can be interpreted as neutrality of money; the second connects the intertemporal marginal rate of substitution and marginal adjustment costs. For integrated (of order one) variables, the log-linearized version of the first-order condition leads to the hypothesis of two cointegration vectors and to the restrictions on those cointegration vectors. The theoretical model is designed for the analysis of the harmonized monetary aggregate, M3H, but it can also be used to the analysis of narrow money, M1.

The estimates of the parameters of the first-order conditions of the M1 model are stable. The test for cointegration rank supports the single cointegration vector. The unit scale elasticity implies that the risk aversion parameters of consumption and money are identical. The interest rate semi-elasticity has reasonable size, 1.8. The recursive estimation of these parameters and the recursive test of the constancy of the full sample cointegration space displays no instability. The GMM estimation of the Euler equation of M1 produces parameters of reasonable size and sign. The system is stable, although there is some indication that adjustment costs have decreased since 1987. This reflects advances in the payment and transfer technologies. The estimate of the risk aversion parameter for the utility function, 0.53, is in the lower range of international comparisons. The residuals can be approximated by white noise, ie they reflect the pure expectational error. The Euler equation passes the formal parameter stability tests. However, there is some indication that the test for overidentification restrictions is alarming for the floating exchange rate period. This might be an indication of the changed endogeneity pattern of the system. Nevertheless, the sample size of the floating exchange rate period, 40, is too small for credible conclusions.

The test statistics for the M3H system do not support the restrictions on the utility function parameters implied by the model: First, the empirical and asymptotic critical values imply a single cointegration vector. Second, this cointegration vector relates the opportunity cost of money and the own-yield of money, but not their difference as the theoretical model predicts. Third, when assuming that there exist two cointegration vectors, the second cointegration vector implies scale elasticity of about three, which is very large value compared with typical international values of between one and two. Finally, the recursive estimation of the scale elasticity betrays significant unsteadiness. The hope for a proper aggregation from M1 to M3H is ruined by the fact that adding the difference between M3H and M1 to the cointegration space of the M3H model leads exactly to the model of M1. The implementation of the Euler equation of M3H naturally yields poor results: hardly any of the estimated parameters differ from zero. The root close to unity in the lead term suggests that M3H might be influenced by wealth rather than consumption or income.

The interest rate measures of monetary policy cause significant changes in asset prices, which influence to the value of wealth. Thus, if wealth is a determinant of money holdings, this creates difficulties for monetary policy. The non-existence of cointegration between price level and M3H implies that the development of M3H and consumer prices might diverge. Then M3H does not fulfil the necessary condition of an good candidate of an inflation indicator. The long-run income elasticity of M3H is approximately twice the magnitude of the European aggregates[21]. The inclusion of the Finnish M3H would increase the aggregation bias in the demand for European-wide M3H. On the other hand, if the ECB chose M3H as an intermediate target and the short-term interest rate as an operational target, instabilities in the demand for M3H could jeopardize the inflation process in Finland. However, the historical instabilities, which are probably caused by the financial deregulation, might disappear before the third stage of EMU.

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A The Data

Empirical counterparts for theoretical variables are the following:

Narrow Money: Narrow monetary aggregate M1, mill. FIM, logarithm. It contains cash held by the public and transactions accounts at the banks.

Harmonized Money: Harmonized monetary aggregate M3H, mill. FIM, logarithm. It contains cash held by the public, all accounts (including foreign currency) at banks and money market deposits and repos at banks.

Prices: Consumer price index (1990=100), logarithm, published by the CSO Finland.

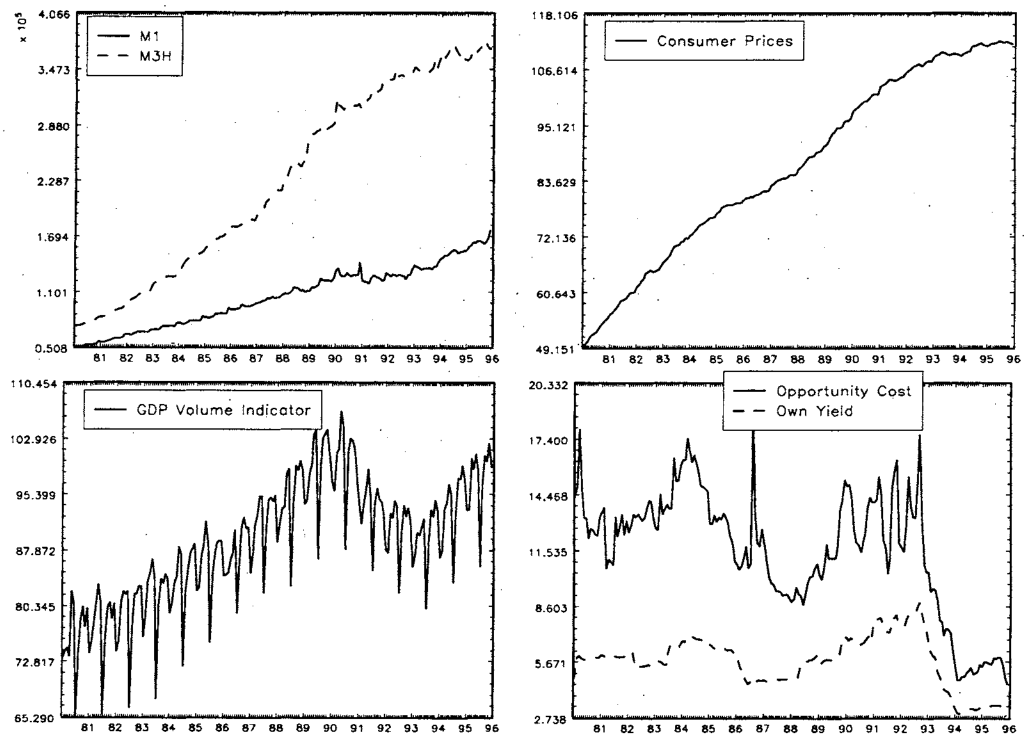
Transactions: Monthly volume indicator of GDP (1990=100), logarithm, published by the CSO Finland. It is a combined index of various indicators such as industrial production, retail sales, consumption of electricity, etc. Since the measure of money contains consolidated money holdings of households and the corporate sector, one cannot use consumption as a scale variable. Instead, we use this GDP indicator and we neglect theoretical consequences of the choice.

Opportunity cost of money: Covered 1-month Eurodollar rate for the markka for the pre-1987 period and 1-month HELIBOR (money market rate) since that, divided by 100, published by the Bank of Finland. For after-tax version, see the explanation below.

Own-yield of M3H: Average after-tax deposit rate (including money market deposits) at banks, divided by 100, published by the Bank of Finland. Most of the bank accounts are tax-exempt, but the situation is changing rapidly. For post-1991 period, we use withdraw tax rate to get after-tax deposit rate for taxable accounts. For pre-1991 period, we use the difference between tax-exempt and taxable bond rates to estimate the level of taxes.

Time period: January 1980 – December 1995. Graphs are presented in figure 5.

Figure 5: Narrow and Harmonized Monetary Aggregate, Consumer Price Index, GDP Volume Indicator, Opportunity Cost of Money and Own-Yield of M3H



There are several exogenous shocks in this data period also. They are modelled with the following dummy variables:

- The seasonal pattern of the GDP volume indicator has changed along with the construction cycle. An extra seasonal variable JULY has been added. It is the ratio of construction to total GDP, where monthly construction is measured by construction licences (CSO Finland). The July value is multiplied by 1 and the August value by –1, while values for the rest of the year are zero.

- Tax rebates are normally paid in December. In the years 1991–1995, the pattern has changed temporarily, and that is modelled by the dummy `REBATE`.

- Devaluation speculation raised interest rates in August 1986 and again in September - December 1991 and finally in April - November 1992, `DSPEC`. Devaluation speculation also measures the *currency substitution* effect.

- Increase of capital gains tax in January 1989 is measured by the dummy `CGAINT`. It is 1 in December 1988, and –1 at end — December 1990, since the special taxfree 24-month time deposit was introduced in December 1988.

- Strike of bank office workers in February 1990 is measured by two dummies. `BSTRIKE1` is 1 in January 1990 and –1 in March 1990, while `BSTRIKE2` is 1 in February 1990. The strike increased cash held by the public and interest rates were frozen. It was not anticipated before the very end of January.

- The introduction of the withholding tax for bank accounts at the beginning of 1991 `WTAX`. A 15% tax on bank accounts stimulated real competition between banks.

- The strike of harbour workers in June 1991 decreased industrial production during that month. The production gap was filled in the following month. That strike is modelled by the dummy `TRAF`.

- During the pre-1987 period, the Ministry of Finance regulated banks' CD issues. `MFREST` has a value of unity during that period and zero otherwise.

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