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Essays in international finance and open economy macroeconomics

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ESSAYS IN INTERNATIONAL FINANCE AND OPEN ECONOMY MACROECONOMICS

by

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I. INTRODUCTION

This dissertation is composed of four essays. The first two essays cover topics in the area of open economy macroeconomics and the last two attempt to answer questions in the area of international finance. All of them could be categorized as empirical research papers even though the first chapter contains a lot of theoretical derivations. It should be noted that I used a number of statistical packages to come up with the results such as SAS, RATS, and GAUSS. The bayesian unit root and fractional differencing procedures utilized in this dissertation come from programs written by Thomas Doan as I understand and are available with the purchase of RATS. The other classical unit root and co-integration tests used here have been programmed in GAUSS language by Sam Ouliaris. The FORTRAN algorithm for the calculation of the extreme-value parameter in the third essay has been written by Prof. Hosking. Finally, the computer code for the Markov switching model with the EM algorithm has been generously given to me by Prof. J. Hamilton to whom I express my gratitude. Since each essay has its own unique introduction, I refrain from providing further details about each one of them in this general introduction.

I. TESTING AN OPTIMIZING TWO SECTOR MODEL OF CONSUMPTION

A. INTRODUCTION

The purpose of this essay is to evaluate the theory of a representative household exhibiting dynamic optimizing behavior which accounts for the observed fluctuations in major macroeconomic variables such as consumption according to neoclassical business cycle theories for some european countries.

It is well-known that there are two ways to do so. The first way is to use first-order conditions to estimate the deep, structural parameters of the agents' utility function and test the restrictions imposed by the model on the data without obtaining explicit solutions. The second way is to construct an artificial economy and compare the computed equilibrium values to the actual ones. Even though the two approaches are complementary in nature, I chose to proceed with estimation rather than simulation because the latter imposes more structure on the theory than the former.

A number of papers have used the representative agent model to study consumption. Among them, Grossman and

Schiller (1980), Mankiw (1981), Hall (1978, 1981, 1988), Hansen and Singleton (1982, 1983), Mankiw, Rotemberg and Summers (1985), Bernanke (1985), Mankiw and Zeldes (1986) estimate the structural parameters of the utility function without deriving explicit solutions. All of them reject the optimizing model. Others such as Abel (1990), Constantinides (1990), Ferson and Constantinides (1991) and Sundaresan (1989) study consumption by introducing utility functions that exhibit habit persistence. Most of the above papers employ the GMM (Generalized Method of Moments) technique. The GMM allows for the estimation of the parameters of the agent's utility function which are not subject to Lucas' critique (1976). It is known that the Lucas' critique basically states that the parameters of the classical consumption function may be unstable over time since they depend on a host of other macroeconomic factors and policy variables and may not be time invariant with respect to changes in the "forcing variables".

My paper differs from all of the above mentioned papers in a number of respects. First, it allows for the presence of two sectors in the economy, the traded and the nontraded sector. Models of this kind have been developed by Bruno (1976), Dornbusch (1983), Stockman (1983), Dellas and Stockman (1989), Stockman and Tesar

(1990), Froot and Rogoff (1991). However, none of them is concerned with the same type of problem as this chapter does. Second, the countries of interest are Ireland, Greece and Portugal here. In the other empirical studies, I am aware of, the nations of interest were usually the USA and a few other industrialized countries. Third, this paper incorporates recent data and thus uses more observations from the post-Bretton Woods era. Fourth, a variable interest rate has been used instead of a constant one. Most of the other papers take the interest rate as constant.

B. THE MODEL

The model describes a small open economy that is inhabited by identical infinitely-lived agents who maximize expected utility from consumption. We assume that they have a CARA class utility of the form:

$$U(C) = C^H/H$$

where $H = 1-k$

$$C = C_T^{a}\ C_N^{1-a}$$

where C is an index for consumption in traded and non-traded goods and has a Cobb-Douglas functional form. The parameter -a- denotes the share of traded goods in consumption, and the other parameter (1-a) denotes the share of non-traded goods in consumption.

This utility function exhibits both constant relative risk aversion (1-H) which is a measure of the concavity of the utility function or the disutility of consumption fluctuations with constant relative prudence (2-H) defined as (-U'''C / U''). Kimball (1987) is credited with showing that the coefficients of prudence can be used to study the effects of uncertainty on expected marginal utility much like the coefficients of risk aversion can be used to study the effects of uncertainty on expected utility. Unfortunately, the isoelastic form of utility which is prevelant in the macroeconomic asset-pricing and real business cycle literature has the attribute of specifying the agent's risk aversion with the same parameter that characterizes the agent's preferences for intertemporal substitution in dynamic applications under uncertainty assuming time and state separable preferences. Ferson and Constantinides (1991) have shown that for the isoelastic utility function the concavity parameter is approximately equal to the relative risk aversion coefficient but is not the inverse of the elasticity of consumption under habit formation.

The representative individual at time t maximizes the following expected utility function:

$$(1) \qquad E_t[\sum_{0}^{\infty} b^t (C_{Tt}^{a} C_{Nt}^{1-a})^H / H]$$

subject to the budget constraint:

$$(2) \qquad W_{t+1} = (1+r_t) W_t + Y_t - C_{Tt} - p_t C_{Nt}$$

The first-order necessary condition for the maximization of (1) subject to (2) is:

$$(3) \qquad b\, E_t[(1+r_t)(C_t/C_{t+1})^k \, (p_t/p_{t+1})^{1-a}] - 1 = 0$$
$$\text{where } k = (1-H)$$

where -b- is the discount rate, -a- is still the share of traded goods in total consumption, k=(1-H) denotes the degree of relative risk aversion, W and Y stand for total wealth and aggregate output measured in tradeable goods, C denotes aggregate consumption in traded goods terms, r is the real world interest rate expressed in tradeable goods and p is the relative price of non-tradeable goods in terms of tradeable ones.

It is obvious from the above equation that the appropriate real interest rate is the variable world interest rate multiplied by the expected rate of change in the price of non-tradeables to tradeables. So, expected changes in the real exchange rate over time should affect aggregate consumption. The first appendix to this chapter provides the intermediate steps en route to the derivation of the equation (3).

C. DATA AND ESTIMATION

The data used for this study come from the International Monetary Fund IFS tape, the IMF Government Finance Statistics Yearbook, the OECD Economic Indicators and the greek Central Bank's bulletin. It is annual data where all variables have been converted into real by dividing them with the CPI (Consumer Price Index) index of the respective country of interest. The CPI index is the proxy for the tradeable price index while the WPI (Wholesale Price Index) is used as a proxy for the non-tradeable price index. Consequently, all appropriate variables have been divided by the population figure to get per capita terms. Of course, these are imperfect proxies but they have been employed in the literature before as in Wolff (1987), Clements and Frenkel (1980). The world real interest rate has been proxied by the ex-post US real discount rate denominated in traded goods. The latter has been constructed by substracting the US CPI inflation rate from the nominal discount rate. The ex-post rate differs from the ex ante rate, assuming rational expectations of inflation, by the inflation forecast error. Mishkin (1984) has a good discussion of this point.

The Hansen's (1982) generalized method of moments procedure is used. It is very attractive for a number of reasons. First, it allows for direct estimation of the model's parameters without obtaining an explicit solution. Second, it does not make any restrictive distributional assumptions for the model's disturbance terms. Third, it obtains consistent estimators of the parameters and the covariances even in the presence of serially correlated and heteroscedastic errors arising out of situations such as the overlapping of forecast horizons. Fourth, it can be used to test the specification of the model. However, this estimation method is based on the assumption that the variables and the instruments used are stationary and ergodic. Stationarity requires that the properties of a stochastic process are unaffected by changes in the time of origin. Constant mean and variance along with autocovariances and autocorrelations depending just on the time lag for any process define what has come to be called weak stationarity. Since I use annual data I do not have to worry about seasonal integration. Ergodicity refers to sample moments approaching the population moments as the sample size of the particular realization tends to infinity. Although we can test for stationarity by looking at the autocorrelation function or/and conduct unit root tests, it is extremely difficult to test for ergodicity since we are working with a single realization of the stochastic process. That is why ergodicity is being assumed in all the studies I have come across. I follow the same route here. The second appendix to this essay describes the GMM technique in a general form.

Another approach to estimating the deep structural parameters is to put to use the following rule applying to any time series Z. If $E_t(Z_{t+1}) = 1$, then $Z_{t+1} = 1+e_{t+1}$ where the error term is conditionally homoscedastic and $E_t(e_{t+1}) = 0$. Manipulating the first-order condition of the model by transferring 1 to the right-hand side, taking the natural log and employing the previously stated rule gives:

$$(4) \qquad k \log\left(\frac{c_{t+1}}{c_t}\right) = \log(b) + \log(1+r_t) + (1-a)\log(p_t/p_{t+1}) - \log(1+e_{t+1})$$

Simplifying the equation further I get:

$$(5)\qquad \log\left(\frac{C_{t+1}}{C_t}\right) = D+\left(\frac{1}{k}\right)\log(1+r_t)+\left(\frac{1-a}{k}\right)\log\left(\frac{p_t}{p_{t+1}}\right) -\left(\frac{1}{k}\right)\left[\log(1+e_{t+1})+\left(\frac{1}{2}\right)\sigma^2\right]$$

$$\text{where}\qquad D = \left(\frac{1}{k}\right)\left[\log b+\left(\frac{1}{2}\right)\sigma^2\right]$$

Linearizing the log of the error term in equation (5) by means of the Taylor approximation which is exact if the error is log-normally distributed gives:

$$\log(1+e_{t+1}) = e_{t+1} - (1/2)(e_{t+1})^2$$

Then equation (5) takes the form:

$$(6)\qquad \log\left(\frac{C_{t+1}}{C_t}\right) = D+\left(\frac{1}{k}\right)\log(1+r_t)+\left(\frac{1-a}{k}\right)\log\left(\frac{p_t}{p_{t+1}}\right)+v$$

$$\text{where}\quad v_{t+1} = \left(\frac{1}{k}\right)\left[\frac{1}{2}(e_{t+1})^2-e_{t+1}-\frac{1}{2}\sigma^2\right];\qquad E_t(v_{t+1}) = 0$$

$$\text{where}\quad k = (1-H)$$

Since the new error term might be correlated with some explanatory variables, the instrumental variables technique is being applied to the equation. Under the assumption that the expected value of the new error term is zero, the instruments are chosen to be lagged values of these variables. One way to test the validity of this assumption is to find out whether or not the null hypothesis of the coefficient of a time-t variable being zero is rejected. This variable belongs to the agents' information set and is being added to the equation. This instrumental variables approach to estimating first-order conditions in intertemporal models has being adopted by Rotemberg (1983) and Mankiw (1985) among others.

Another complication in estimation arises out of the fact that the expected value of a stochastic process might be different from its time average. Aggregate per capita consumption is measured in reality as an average over a time period but the first-order condition of the model refers to it at a point in time here. This implies that the error term in equation (6) follows a first-order moving average process with coefficient 0.27 as it is known from Flemming's theory (1960). The serial correlation and the endogeneity problem give rise to a much more demanding situation to which the Hayashi-Sims estimator (1983) offers an answer. Hall (1988) takes this into account and proceeds with his computations after correcting for the problem using the same estimator.

The Hayashi-Sims estimator performs an autoregressive transformation to each variable by substracting its future values. Since the residuals follow a MA(1) process here, the required trasformation can be approximated by:

$$(7) \qquad \Delta(Z_t) = \Delta(Z_t) - 0.27\,\Delta(Z_{t+1}) + (0.27)^2\,\Delta(Z_{t+2})$$

where Δ *is the change operator*
while Z_t *is any time series.*

The instrumental variables technique is being used subsequently with instruments the lagged original variables.

D. EMPIRICAL RESULTS

Based on the empirical evidence presented in the following tables, one can see that the the values of the -k- parameter are all negative for Greece and Portugal and for all lags. The range goes from -0.071 to -0.38 for Greece and from -0.36 to -0.746 for Portugal. As instruments are taken to be a constant and the lagged values of the consumption growth ratio, price growth ratio and the U.S. real interest rate for 1,2,4 and 6 lags. The figures remain negative for whatever plausible value of -a-, the share of traded goods in total consumption I imposed. The different values of -a- varied from 0.2 to 0.8 so as not to miss the value implied by the OECD International Sectoral Data Base. The government services, private services, construction, electricity, gas and water as well as finance, insurance and real estate are assumed to belong to the non-traded sector. Classical unit root test based on the ADF methodology found all relative variables to be stationary which permitted GMM estimation. As I point out though in later essays to this dissertation, these results should be viewed with caution given the very small sample at hand. However, the same coefficient turns out to be positive for Ireland after allowing for one and two lags in the instruments. It is negative for the case of four and six lags. The standard errors also get smaller as the number of lags increases for all countries. The discount factor takes on values that range from 0.95 to 0.99 which is less than unity. So, the two sector model considered provides economically meaningful estimates for the discount rate and especially the deep structural parameter of the objective function for Ireland in the case of one and two lags since they are less than 1 being 0.69 and 0.81 respectively.

The chi-square statistic which is used to test the over-identifying restrictions of the model on the data whose number is indicated by df in the tables below shows some interesting patterns. At the 5 percent level of significance, the over-identifying restrictions cannot be rejected when the number of lags in the instruments is 6 for all three countries. The restrictions cannot be rejected for Ireland and Portugal at the same level of significance after allowing for 4 and 2 lags respectively. Interestingly enough, the null cannot be rejected in all cases at the 1 percent level of significance except for Greece at lags 1 and 2 as well as Ireland at lag 1. Note that the weighting matrix used to minimize the appropriate function in the process of the GMM estimation has undergone the transformation recommended by Newey and West (1987) to take into account heteroscedasticity and autocorrelation.

Table-IV below reports the coefficient estimates of equation (6) after performing the necessary transformations called for by the Hayashi-Sims estimator (1983). The only coefficient that is significant other than the constant is that of the price growth ratio in the case of Portugal. On the basis of the regression results the composite parameters for each nation suggest different values for the risk-averse parameter in the same eqaution for different values of the share of traded goods in aggregate consumption at the same time. Thus, no definite conclusion can be drawn as to the meaningfulness of the parameter estimates from this method of parameter estimation. The relevant tables with the empirical results follow in the next few pages.

TABLE-I

GMM ESTIMATES FOR GREECE: YEARLY DATA 1948-1992

NLAG	b*	k*	H	chi-square!	df
1	0.97 (0.015)	-0.38 (0.38)	1.38	10.45 (0.001)	1
2	0.98 (0.008)	-0.07 (0.233)	1.07	14.52 (0.002)	3
4	0.98 (0.005)	-0.15 (0.123)	1.15	17.13 (0.017)	7
6	0.98 (0.003)	-0.15 (0.077)	1.15	17.96 (0.082)	11

* Standard errors in parentheses.

! Significance levels in parentheses.

Newey-West adjustment performed.

TABLE-II

GMM ESTIMATES FOR IRELAND: YEARLY DATA 1948-1992

NLAG	b*	k*	H	chi-square!	df
1	0.99 (0.008)	0.31 (0.299)	0.69	7.03 (0.008)	1
2	0.99 (0.006)	0.09 (0.245)	0.91	8.12 (0.043)	3
4	0.98 (0.005)	-0.11 (0.150)	0.89	12.84 (0.076)	7
6	0.98 (0.003)	-0.04 (0.082)	0.96	15.04 (0.181)	11

* Standard errors in parentheses.

! Significance levels in parentheses.

Newey-West adjustment performed.

TABLE-III

GMM ESTIMATES FOR PORTUGAL: YEARLY DATA 1948-92

NLAG	b*	k*	H	chi-square!	df
1	0.95	-0.75	1.75	3.85	1
	(0.02)	(0.46)		(0.05)	
2	0.96	-0.42	1.42	7.12	3
	(0.01)	(0.17)		(0.07)	
4	0.96	-0.36	1.36	16.44	7
	(0.001)	(0.04)		(0.021)	
6	0.95	-0.41	1.41	17.89	11
	(0.005)	(0.058)		(0.084)	

* Standard errors in parentheses.

! Significance levels in parentheses.

Newey-West adjustment performed.

TABLE-IV

INSTRUMENTAL VARIABLES ESTIMATION
USING HAYASHI AND SIM'S ESTIMATOR.

	GREECE	IRELAND	PORTUGAL
Constant	0.06	0.01	-0.02
	(0.02)	(0.06)	(0.22)
(1-a)/k	-0.63	2.30	-0.93
	(0.69)	(2.88)	(0.24)
(1/k)	0.25	-0.46	6.18
	(1.07)	(1.22)	(0.87)

All numbers have been rounded off to two decimal points.

Standard errors in parentheses.

E. CONCLUSION

I have shown that the two sector model developed above seems to yield economically meaningful and plausible estimates for the risk-averse parameter in the case of Ireland after allowing for a small number of lags in the instruments. Also, the chi-square test results indicate the non-rejection of the restrictions imposed by the model for all countries and for a large number of lags at the 5 percent significance level. The non-rejection of the null holds even stronger at the 1 percent level of significance. Thus, the model is promising and future research should be directed into allowing for other utility functional forms which might include habit persistence, data of a longer time span and/or of a different frequency.

III. DEBT AND THE FEASIBILITY OF PUBLIC FINANCIAL POLICIES

A. INTRODUCTION

The effects of macroeconomic governmental policies on the economy have been the subject of empirical investigation for many years. Most of the research in this area has been directed toward the impact of different budget policies on variables such as inflation, interest rates, GNP growth, the trade balance, private investment spending and others assuming that the government will have no problem in financing its expenditure by borrowing. Papers in this tradition have been published by Barro (1979, 1980), Sargeant (1986), Lucas and Stokey (1983). On the other hand, other empirical work has paid attention to the feasibility of running chronic budget deficits. Papers in the latter tradition have been the ones by Hamilton and Flavin (1986), Hansen, Roberds, and Sargeant (1987), Wilcox (1989), Grilli (1987), Trehan and Walsh (1988), Kremers (1988, 1989) and Smith and Zin (1991). These papers test the intertemporal budget constraint in an attempt to shed light on the question of the sustainability of fiscal policy. Most of them test the U.S. solvency constraint using constant interest rates or stochastic ones with the expected one period ahead rate assumed constant. Grilli (1987) employs E.E.C. data while the Smith and Zin (1991) tests are based on Canadian data.

It is well known that the effects of budget deficit policies on the economy as well as the efficacy of these policies because of the long-run constraints that debt accumulation imposes on their sustainability are related. This essay is concerned about debt accumulation and its effects on fiscal behavior. In particular, it tests the hypothesis that public financial policy satisfies the present-value borrowing constraint for different E.E.C. countries. Starting with Hamilton and Flavin (1986), a number of papers have used direct tests of the intertemporal budget constraint such as the ones by Trehan and Walsh (1988), Grilli (1989), Smith and Zin (1991), Wilcox (1989) and others. This essay uses recent annual data, employs a variable interest rate as Wilcox (1989) and Smith and Zin (1991) did in their respective papers for U.S. and Canada, tests the sustainability of some of the E.E.C. nations' policies as Grilli (1989) did, but under the assumption of a constant interest rate, using a number of different methods of testing for unit roots and common trends, a few of which have not being utilized in the previously mentioned studies.

B. THE MODEL

We start from the identity faced by the government authorities which is expressed as:

$$(1) \qquad B_t = (1+i_{t-1})\ B_{t-1} + G_t - T_t - \frac{(M_t - M_{t-1})}{P_t}$$

B_0 is given and t=0,1,2,...

where B is the value of outstanding government debt adjusted for inflation, i_{t-1} is the ex-post, real interest rate paid on debt, G represents real government expenditures, T represents real tax revenues, M is the monetary base and P is the price level.

This identity is derived from the nominal budget constraint identity and can be treated as a stochastic equation since there are a number of problems associated with timing conventions, errors in measuring variables, and time aggregation problems. A good discussion of the above points can be found in Smith and Zin (1991). This paper follows closely Smith and Zin's (1991) work except for the differences emerging out of utilizing different estimation techniques, time span and countries of interest. Then, the same identity can be presented as:

$$(2) \qquad B_t = (1+i_{t-1})B_{t-1}-S_t+e_t$$

where S stands for the budget surplus inclusive of seigniorage.

Solving forward the above equation, we get:

$$(3) \qquad B_t = E_t\sum_{k=1}^{\infty}(1+r_{t+k-1})^{-1}(S_{t+k}-e_{t+k}) + \\ + \lim_{N\to\infty}E_t(1+r_{t+n-1})^{-1}B_{t+n}$$

in which

$$(1+r_{t+k-1}) = \sum_{l=t}^{t+k-1}(1+i_l)$$

In the above equations, E denotes the expectations of bondholders at time t which are assumed to conform to the rational expectations hypothesis which coupled with a transversality condition yields two equivalent testable hypotheses:

$$(4a) \qquad H_0:\ \lim_{n\to\infty} E_t \frac{B_{t+n}}{(1+r_{t+n-1})}$$

$$(4b) \qquad H_0:\ B_t = E_t \sum_{k=1}^{\infty} (1+r_{t+k-1})^{-1} (S_{t+1} - e_{t+1})$$

The above two hypotheses are equivalent to the extent that S is of exponential order less than (1+i). Equation (4a) implies that the average growth rate of debt cannot exceed the borrowing rate. McCallum (1984) and Hamilton and Flavin (1986) showed that a constant deficit inclusive of interest payments on the debt is consistent with the intertemporal budget constraint when this transversality condition is being satisfied. The null hypothesis also states that bondholders expect the government to balance its budget in present value terms. A rejection of the hypothesis signals either that the current fiscal policy is unsustainable or that the government faces no constraint in borrowing. Smith and Zin (1991) emphasize this point. Even if the hypothesis is not rejected, one should be careful before labeling the public financing policy as sustainable. Satisfying the intertemporal solvency constraint is a necessary but not a sufficient condition for sustainability under certain conditions.

When the borrowing rate falls short of the growth rate of real output, it is generally accepted that the government will be able to service its outstanding debt by issuing new bonds since the base against which the government borrows will increase faster than its borrowing needs. In other words, its taxing capacity will surpass its borrowing undertakings. Under these circumstances, the debt ratio, that is, the ratio of debt to GNP will keep on increasing until it finally reaches a stationary value even though the government runs constant primary deficits all this time. Thus, in general, as long as the growth rate of real income exceeds the borrowing rate, expansionary fiscal policies will not be constrained by a limit on the growth rate of debt. Many researchers such as Carmichael (1982), Hamilton and Flavin (1986), Kremers (1989), Spaventa (1987) and Buiter (1985) have referred to this case. It should be pointed out however that real world tax-evasion and individual as well as business defaults render the main conclusion of this paragraph very shaky even though the real interest rate might be less than the growth rate of real GNP.

The case where the borrowing rate is greater than the growth rate of real output is more interesting indeed.

Kremers (1989) has shown that a fiscal policy observing the intertemporal solvency constraint will not be sustainable in the long run if an upper limit on the debt ratio restricts tax capacity. If such a limit exists, then the average rate of growth of the stock of debt must be less than the borrowing rate as well for a sustainable public financial policy. However, McCallum (1984) has pointed out that as long as the average debt growth rate is less than the sum of the growth rate of real income and the real interest rate, the government should have no problem in rolling over its debt. This is so because interest payments are part of total household income.

C. DATA AND ESTIMATION

Annual data from the IFS tape, the IMF Government Finance Statistics Yearbook are used for the period 1948 to 1992. The countries of interest are Greece, Ireland and Portugal. The budget deficits are equal to the difference between revenue, including grants where applicable, and expenditure as noted in the IFS documentation. The measure of the debt refers to direct and indirect central government debt excluding government guaranteed loans. The real interest rate is the discount interest rate adjusted for inflation using the CPI index. Since the real interest rate is variable, I look into the nature of the comovent between ($r_t B_t$) and S_t as others have done, notably Smith and Zin (1991). In other words, I test whether the variable ($S_t - r_t B_t$) is stationary. In so doing, I employ some unit root tests namely Park and Choi's (1988) J(1,5) test, P.C.B. Phillips' (1987) test. I also test to see whether the budget variables for each of the three countries have some kind of long-run relationship. To do so, the discriminating procedures of Stock and Watson's (1988) common trends statistic as well as Johansen's (1988) trace and maximum eigenvalue statistics are being applied. The above multivariate unit root procedures share the advantage that no pre-determination of a difference stationary time series is necessary for the conduct of the tests. They also help detect the specific number of co-integrating relationships and estimate them. Below, I refer to them.

In this paragraph, I briefly discuss Stock and Watson's test (1988). The best source of information about it remains the original article by Stock and Watson (1988). Given an Nx1 vector of variables, the Stock and Watson procedure tests the hypothesis that this vector has L=<N distinct unit roots. The alternative hypothesis is that the same vector has K<L unit roots. To do so, the (Nx1) vector is being modified in such a way so as to make the

first (N-L) elements stationary and the remaining L elements nonstationary. The eigenvalues of the first-order autoregression of the L nonstationary variables should be equal to 1 if the hypothesis that there are L stochastic trends were correct. Under the alternative hypothesis, only K<L eigenvalues should be equal to 1 implying that the K+1th largest eigenvalue should be less than 1. If we let T denote the sample size and G_{k+1} the K+1th largest eigenvalue, then the test statistic for their procedure becomes $T*(G_{k+1}-1)$. By comparing this statistic with the critical values given by Stock and Watson (1988), one makes an inference about the null hypothesis. Note that this statistic is based on detrended dependent variables in our computations which in turn defines the alternative hypothesis to the null as being stochastic co-integration here. To test for the much stronger hypothesis of deterministic co-integration, one should have proceeded detrending the estimated common trends. In this case, I examine three variables so that N=3 and the null hypothesis is that they have three common stochastic trends, L=3, that is the system is non co-integrated, versus two non-stationary trends, K=2, under the alternative hypothesis of stochastic co-integration.

The other multivariate unit root test used in this essay is Johansen's (1988, 1989, 1990). I briefly describe it as well. The 1988 article by the same author remains the classic source of information. Note that the exact maximum likelihood approach of the author relies on the assumption of a Gaussian VAR of a known order that takes the form of an error correction model. It is clearly a pure parametric technique. Let Y_t be an (Nx1) vector of variables. Then the following two regressions are run:

$$(5) \quad \Delta Y_t = \sum_{i=1}^{p-1} \mathrm{B}_{1i} \Delta Y_{t-i} + R_{1t}$$

$$(6) \quad Y_{t-p} = \sum_{i=1}^{p-1} \mathrm{B}_{2i} \, \Delta Y_{t-i} + R_{2t}$$

where R_{1t} and R_{2t} represent the residual vectors, B_{ji} refer to matrices of coefficient estimates and -p- defines the order of the VAR. These two residual vectors will be used in a second round regression from which the r eigenvectors will be chosen after minimizing the determinant of the covariance matrix of the residuals of the second round regression. According to Johansen(1988), we can employ a likelihood-ratio test based on the r largest eigenvalues to make inferences about the null hypothesis that there are at least (N-k) stochastic trends. Here -k- stands for the maximum number of co-integrating factors. Another way to state the null is that at most k co-integrating factors are present. This so called trace statistic by Johansen is given by:

$$(7) \quad -2 \ln Q_k = -T \sum_{i=k+1}^{N} \ln(1-r_i)$$

Johansen provides the critical values for the trace statistic which is basically a likelihood ratio statistic. These values depend on both the (N-k) stochastic trends and the presence of trends in the series. Note that the absence of any time trend from the above equations translates into testing for the very strong assumption of deterministic co-integration. In addition to this statistic, Johansen's so called maximum eigenvalue statistic is being given here. The latter tests the null hypothesis of k co-integrating relationships to the alternative of (k+1).

Here, the order of the VAR process is taken to be 2, that is, p=2 with N=3 because we have 3 budget variables under consideration while the number of co-integrating factors is allowed to take different values. The same test is performed for 2 budget variables found to be integrated of order one as I discuss it later. The null hypothesis is for no co-integrating vectors.

To supplement the pure parametric approaches to the estimation of co-integrated systems, I employed the non-parametric approach to estimating the short-term dynamics but parametric as far as formulating long-run equilibria goes by Park (1990b) labeled CCR. CCR stands for Canonical Cointegrating Regressions. Park (1990a) shows that all the approaches mentioned above are equivalent in infinite samples. Monte Carlo simulations by Park and Ogaki (1991) have shown the CCR estimator to be superior to Johansen's (1988) maximum likelihood estimator in small samples. In addition to all these, this estimation method does not require any strong distributional assumptions as does Johansen's test. Note that Johansen's procedure is essentially an application of full information maximum likelihood (FIML) to the whole system and as such is vulnerable to equation misspecification. The positive side for these methods is that system estimation incorporates all available information in estimating equation parameters. On the other hand, the CCR procedure is an equation by equation method of estimation of the whole system.

D. EMPIRICAL RESULTS

Unit root tests performed on the gross deficits of Greece, Ireland and Portugal showed some interesting results. First, the univariate unit root tests such as Park and Choi's (1988) J-test as well as P.C.B. Phillips' tests (1987) agree that the Portuguese budget figure does not conform to the unit root hypothesis. This is not so for the other two countries' budgets for which the tests point to the non-rejection of the unit root hypothesis. However, the Z_a test favors the rejection of the non-stationarity for the Greek budget. Once again, the small sample prohibits me from making a definite statement. However, it looks like Portugal's budget is balanced intertemporally. This is the message of the univariate classical unit root tests.

In addition to the univariate tests, I tested the data employing multivariate tests such as Stock- Watson's (1988) and Johansen's (1988). In doing so, I used two systems of variables. A three variable and a two variable system. I did so because the univariate tests before revealed that only two budget variables were integrated of order one. However these multivariate tests do not require that all variables are difference-stationary. Thus, I decided to proceed with the two system estimation. The idea behind these tests is to uncover some long-run relationship between the variables of interest that would allow for a model of the joint determination of government financing policies. The long-run relationship could have been due to the existence of common structural problems and the exposure to the same external shocks. The Stock-Watson test rejected the null of non-cointegration for the two variable system of Greece and Ireland. The value of the statistic was above the reported critical values. It pointed to the presence of one co-integrating factor. However, the same test did not yield the same results when the system was enlarged to three variables.

The other multivariate test, Johansen's test (1988), provided support for the existence of two co-integrating factors when the system contained three variables and one factor when the variables in the system were two. All these based on the trace statistic. On the other hand, the maximum eigenvalue statistic is bigger when the system contains two variables rather than three at all levels of significance and the null hypothesis of two co-integrating factors cannot be rejected at the 5% and 10% levels of significance.

Finally, the CCR and FM estimators are used to compute the co-integrating vector while employing the H(0,5) test statistic to test for deterministic co-integration. The estimates of the two procedures are remarkably close and the H(0,5) test rejects the null hypothesis that the budget variables for Greece and Ireland are co-integrated under both the CCR and the FM (Fully Modified) methods. The following tables provide all the unit root and co-integration results.

TABLE-V

UNIT ROOT TEST RESULTS ON THE BUDGET INCLUSIVE OF INTEREST CHARGES

	GREECE	PORTUGAL	IRELAND
J(1,5)	0.00022**	0.00021	0.00024**
J(-1,5)	133.668	0.146*	13.748
Z_a	3.29*	0.0791*	-44.988
Z_t	4.757	0.05029*	-6.9716

* It indicates rejection of the unit root at the 5% level.

** It indicates rejection of trend stationarity at the 5% level.

The J(1,5) statistic is for the null that the budget variable is non-stationary around a constant and a trend.

The J(-1,5) statistic is for the null of a unit root without drift.

TABLE-VI
TEST FOR COMMON TRENDS ON THE
BUDGET VARIABLE INCLUSIVE OF INTEREST CHARGES

VARIABLES: Greece, Ireland & Portugal

	TEST STATISTIC	CRITICAL VALUES 1%	5%	10%
STOCK-WATSON (3,2)	0.146	-46.7	-38.3	-34.48
STOCK-WATSON (2,1)	-61.125*	-38.5	-30.3	-26.50

Null Hypothesis: NON CO-INTEGRATED SYSTEM
The alternative hypothesis is for stochastic co-integration since the data have been detrended.

(3,2): Three common trends vs. two common trends
(2,1): Two common trends vs. one common trend.

Two variables: Greece and Ireland

TABLE-VII
JOHANSEN'S MULTIVARIATE TEST FOR UNIT ROOTS ON THE BUDGET VARIABLE INCLUSIVE OF INTEREST CHARGES.

VARIABLES: Greece, Ireland & Portugal

(r,N)	MAX STATISTIC	CRITICAL VALUES 1%	5%	10%
(2,3)	0.03	0.004	0.078	0.27
(1,3)	25.64*	2.25	3.35	4.09
(1,2)	2.28*	0.003	0.078	0.27
(0,3)	40.77*	5.69	7.28	8.26
(0,2)	55.45*	2.25	3.359	4.09

Null: r co-integrating vectors.
Alternative: (r+1) co-integrating vectors

N stands for the number of variables in the system.

The order of the VAR is taken to be 2.

Two variables: Greece and Ireland

* Reject the null.

TABLE-VIII

JOHANSEN'S MULTIVARIATE TEST FOR UNIT ROOTS ON THE BUDGET VARIABLE INCLUSIVE OF INTEREST CHARGES.

VARIABLES: Greece, Ireland & Portugal

r	TRACE STATISTIC	CRITICAL VALUES 1%	5%	10%
		9.91	12.29	13.711
0 (N=3)	66.45*			
1 (N=3)	25.67*			
2 (N=3)	0.03			
0 (N=2)	57.74*			
1 (N=2)	2.29			
		2.76	4.06	5.058

Null: r=0 co-integrating vectors.

N=3 stands for the number of variables in the system.

The order of the VAR is taken to be 2.

* reject the null

TABLE-IX
CCR, H(0,5) CO-INTEGRATION TESTS ON THE BUDGET VARIABLE INCLUSIVE OF INTEREST CHARGES.

EQUATION: Greece = a + b Ireland

	CCR ESTIMATES	H(0,5) STATISTIC	P-VALUE
coefficient	1.327	24.335*	0.00018
	(6.115)		
constant	-2.307		
	(-0.805)		

Null: The Greek and the Irish budget variables are co-integrated.

The H(0,5) test is for deterministic co-integration and the statistic follows a chi-square distribution with 5 d.f.

* Reject the null hypothesis.

The PARZEN kernal has been employed in getting the CCR estimates.

TABLE-X

PHILLIPS AND HANSEN'S FM & H(0,5), FM CO-INTEGRATION TESTS ON THE BUDGET VARIABLE INCLUSIVE OF INTEREST CHARGES.

EQUATION: Greece = a + b Ireland

	FM ESTIMATES	H(0,5) STATISTIC	P-VALUE
coefficient	1.323	23.202*	0.0003
	(6.25)		
constant	-2.34		
	(-0.81)		

Null: The Greek and Irish budget variables are co-integrated.

The H(0,5) test is for deterministic co-integration and the statistic follows a chi-square distribution with 5 d.f.

* Reject the null hypothesis.

The PARZEN kernal has been employed for the FM estimates.

E. CONCLUSION

There is conflicting evidence as to whether or not there exists a long-run relationship between the budgetary figures of Greece and Ireland with some tests favoring it and others rejecting it. The same inconclusive evidence emerges out of multivariate co-integration testing with some test statistics favoring the presence of two co-integrating factors and others of one factor. The choice of the number of budgetary variables included in the system under study seems to be crucial to the determination of the number of co-integrating factors. One should cautiously conclude on the basis of the univariate unit root tests that Greece and Ireland have not succeeded in balancing their government financial positions intertemporally. On the other hand, Portugal seems to be on the right avenue. However the satisfaction of the necessary condition for the present-value borrowing constraint should not lead to a strong statement about the feasibility of Portugal's budgetary policies since the sufficient condition for the latter has not been tested here.

IV. UNIT ROOT TESTS AND ESTIMATION OF THE GENERALIZED EXTREME-VALUE DISTRIBUTION OF FOREIGN EXCHANGE RETURNS: EVIDENCE FROM A TURBULENT EMS PERIOD

A. INTRODUCTION

There is a general agreement that the log of spot exchange rates is non-stationary and high frequency data on foreign exchange returns are conditionally heteroscedastic and have unconditional distributions which exhibit fat-tail behavior relative to the normal distribution. Kurtosis has been estimated to be above 3, a number that corresponds to normality, where kurtosis is defined as the ratio of the fourth moment over the square of the second moment. The presence of the so called leptokurtosis in the foreign exchange market has been documented by extensive studies from Westerfield (1977), Boothe and Glassman (1987) as well as Diebold and Nerlove (1989). The nature of the unconditional distribution is important in the sense that we have to use the correct measure of dispersion to account for the effects of exchange rate uncertainty on certain macroeconomic and financial variables as well as to employ the appropriate hypothesis tests and estimation methods to different models of exchange rate determination as pointed out by

Boothe and Glassman (1987). In addition to kurtosis, some evidence of skewness which is defined as the ratio of the square of the third moment to the cube of the second moment has been found in some studies such as in Boothe and Glassman (1987) and Friedman and Vandersteel (1982). However the amount of skewness was found to be relatively small and both papers cited above proceeded on the assumption of distribution symmetry.

It is the empirical results in the above studies that have played a major role in influencing the course of research in exchange rate modelling. Researchers gave up on the open economy Keynesian macro model in the early 70's and focused their attention on the linear structural models such as the flexible and sticky price monetary models developed by Dornbusch (1976), Frenkel (1976, 1979) and Bilson (1978). After enjoying an initial success, the models collapsed subsequently and all efforts to resurrect them by correcting for simultaneity bias and relaxing all kinds of restrictions proved fruitless. The next generation of models was the linear nonstructural models embedding the concepts of co-integration by Engle and Granger (1987) and error correction specification by Hendry et al. (1984) which did not prove of much help in exchange rate forecasting as demonstrated by Meese and Rogoff (1988). However, they revealed the serious misspecifications in the linear structural models as demonstrated by Boothe and Glassman (1987b). In a parallel line of research, new models seeking to exploit the conditional heteroscedasticity found in exchange rate and other asset returns were built. These were different variants of the original ARCH model by Engle (1982). Pursuing this kind of research in nonlinear models, others sought to estimate the functional form linking the exchange rates to their fundamentals by using non-parametric techniques such as Chinn (1991) while others employed similar techniques to uncover nonlinearities in the conditional mean thinking that this might explain conditional heteroscedasticity since otherwise the latter should be an integral feature of the data-generating process. Diebold and Nason's (1990) paper falls into this category. Finding non-linearities in the conditional mean continues to be an area of active research drawing ideas from regime switching, Engle and Hamilton (1990), the ARCH-M model of Engle, Lillien and Robins (1987) and the theory of deterministic chaotic systems, Hsieh (1989). Despite all the efforts to improve exchange rate point predictions by building new models, the results have been poor for out-of-sample forecasting so far. None of them seems to dominate the random walk. So, high-frequency returns in the foreign exchange markets seem to be approximated well by martingale processes.

The stylized fact is that exchange rate changes follow the normal distribution when the data are of the low frequency type, particularly the annual and quarterly one. The evidence is mixed about monthly changes. The normal distribution of course offers the advantage that it can be described by its centrality and dispersion parameters which is not the case with non-normal distributions. Among the candidate distributions that have been considered in the non-normal category are the family of symmetric stable Paretian distributions which incorporates the normal and the Cauchy distributions as special cases, the mixture of normal distributions, the Student distribution and the sum-stable laws.

The mixture of normal distributions is a standard probability distribution which can can be thought of as the superposition of a number of simple normal distributions all of which share the same mean. The latter feature explains its symmetry. The parameters that describe it are its location, the number of its scales and a mixing parameter which basically accounts for the fraction of realizations that were likely to have been drawn from the first distribution. Note that this distribution degenerates back to the normal distribution if all scales (standard deviations) are equal.

The family of symmetric stable Paretian distributions enjoys the advantage of having an unchanged characteristic component as observations are added but bears the disadvantage of nct having the second and higher-order moments as pointed out by Boothe and Glassman (1987). The normal distribution which is a member of this family of distributions has for a value of 2 of the characteristic component the second and higher moments but the presence of many outliers in the exchange rate yields data renders it unattractive.

The Student distribution, although not a stable one, has become very popular in the literature. Its main attraction is a combination of finite variance for degrees of freedom greater than 2 and fat-tailness. As the degrees of freedom go to infinity, it approaches the normal distribution but other than this it is flatter and has longer tails than the latter distribution.

The sum-stable laws combine the fat-tailness with the additivity property which characterize financial market yields.

All the above mentioned distributions are non-nested which creates a problem in picking the one that best fits the data. The typical likelihood ratio test is misleading since the ranges of the likelihood functions are unequal for these distributions. The Cox test for non-nested hypotheses cannot be used in most cases since a number of the previously mentioned distributions do not have a second moment as noted by White (1982), and the Pearson chi-square goodness of fit test that Boothe and Glassman (1987) used in their work suffers from the fact that observed and expected frequencies have to be compared on the basis of arbitrarily chosen intervals. Faced with the empirical regularity of fat-tailness and the problems in identifying the specific distribution, Koedijk, Schafgans and Vries (1990), KSV from now on, employed extreme value theory which takes into account the fat-tail property of returns and tried to answer questions related to the distribution that best fits EMS (European Monetary System) data for the Belgian Franc, the French Franc, the Italian Lira, the Dutch Guilder, the British Pound and the Danish Krone quoted vis-a-vis German mark spot rates from 1979 to 1989, as well as the stability of the parameters of the distribution over different subsamples and finally the effects of time aggregation on the distribution. In so doing, they estimated the tail index which ends up characterizing the limit law of the distribution of the maxima and used it to discriminate against some of the distributions. Their results favored the sum-stable hypothesis but did not reject the Student-t distribution either. They also found that the EMS did not have much of an effect in reducing extreme volatility during the period up to 1989. This method of estimation centers on discovering just the distribution of the tails but allows for the construction of confidence intervals exploiting the asymptotic normality of the reciprocal of the estimated tail index.

I follow a similar route here to the KSV study. I estimate the shape parameter of the generalized extreme-value distribution (GEV) developed by Jenkinson (1955) on daily ERM (Exchange Rate Mechanism) data for the turbulent period August 20, 1992 to February 20, 1993. In addition to this, I perform a number of unit root tests on levels and differences to determine whether or not they conform to the stylized facts mentioned above.

B. THE MODEL

The non-stationarity hypothesis is tested by means of classical, Bayesian and other tests. All the tests used are well-known and a short description is provided either in this chapter or in a previous one. The Park and Choi's J(1,5) and the Sims' slightly modified Bayesian test are being complemented by the Geweke-Porter-Hudak estimation of the fractional differencing parameter and the inspection of the response of each exchange rate series to univariate shocks. Last but not least, Johansen's multivariate unit root test is employed to provide evidence on the number of co-integrating vectors on level data.

The properties of extreme values have long been appreciated in areas such as ocean engineering, highway traffic, hydrology, meteorology, pollution control and others. In this essay I look into the nature of the approximate distribution that the extreme maxima events in the EMS took during one of its recent upheaval. A brief exposition to the extreme-value theory is given in the appendix to this chapter. I consider Jenkinson's (1955) generalized extreme-value distribution, GEV from now on, which incorporates all three Fisher-Tippett (1928) types of extreme-value distributions using the maximum-likelihood approach to derive estimates of the shape parameter as in Hosking, Wallis and Wood (1985). Maximum-likelihood estimation is well established on grounds of asymptotic theory and seems to perform better than the sextile method of Jenkinson (1969) and equally well as the method of probability-weighted moments of Hosking, Wallis and Wood (1985) in small and medium-size samples. The numerical method used was that of Newton-Raphson in the context of Hosking's algorithm (1985).

Let Y be a random variable with distribution function F(y). The GEV distribution which engulfs all three types of limiting distribution for extreme values is described by:

$$(1) \quad F(y) = \exp\,[-(1-a(y-b))/c^{1/a}], \quad for\ a \neq 0$$

$$= \exp\,[-\exp\,(-(y-b)/c)], \quad for\ a = 0$$

where a, b, c are the shape, location and scale parameters respectively. If a>0, then y is bounded by (b+c/a) from above and if a<0, then y is bounded by (b+c/a) from below. The inverse distribution function is given by:

$$(2) \quad y(F) = b + c\,(1-(-\log F)^{a}]/a, \quad for\ a \neq 0$$

$$= b - c\,\log(-\log F), \quad for\ a = 0.$$

The log-likelihood that is being maximized takes the form as in Hosking (1985):

$$(3)\ L(y;k) = -n \log c-(1-a) \sum_{j=1}^{n} z_j - \sum_{j=1}^{n} e^{z_j}$$

where $z_j=-a^{-1} \log(-a(y_i-b)/c)$. As Hosking, Wallis and Wood (1985) note, the log-likelihood will take a large value by setting a>1 and making the (b+c/a) upper bound almost equal to the biggest observation by previously picking the appropriate values for b and c. On the other hand though, it is possible that the function might not converge if even a local maximum cannot be found.

The values taken by the shape parameter determines the type of the extreme-value distribution. For a=0, the Fisher-Tippett type I distribution or Gumbel distribution is appropriate. A number of well-known distributions belong to this family such as the normal, the log normal, the exponential and even the gamma distribution. For a>0 and a<0, the Fisher-Tippett types III and II distributions are appropriate.

C. DATA AND ESTIMATION

The data set includes the 100 times the first log differences of daily bilateral exchange rate bid prices for the period August 20, 1992 to February 20, 1993. All of them are quoted in DM per other currency terms and represent London closing prices. Initially, all exchange rates were expressed in terms of the US dollar. The main reason for using the DM is the fact that Germany represents the biggest economy in the EEC and the DM is being regarded as the anchor currency in the EMS. It also enjoys the status of a reserve currency along with the dollar and the Japanese yen. The other currencies examined were the British pound, the Irish punt, the Dutch guilder, the Belgian franc, the Danish krone, the Portuguese escudo, the Spanish peseta, the Italian lira, the French franc and finally the ECU. Although some of them dropped out of the ERM of the EMS for the period I study such as the pound and the Italian lira, I think that it is important to look at their distributions as well since there is speculation that they will rejoin the system in some fashion in the future and it is useful to have some information with regard to their behavior in a non-tranquil environment like the period I look into. Hoskings' algorithm was used to come up with the maximum likelihood estimates of the relevant parameters a, b, c.

The initial value for the shape parameter -a- was set to be equal to 0 for all exchange rates under study. The initial values for the location and scale parameters were set to be equal to the sample mean and standard deviation for each exchange rate series.

Tables-XI through XIV help shed light on the question of unit roots in the series at hand. Table-XI presents evidence on the non-stationarity property of the log in levels for the ten exchange rates under study. Table-XII reports the results of the unit root tests on the 100 times the log of first differences in each series. The unit root tests have been chosen so as to represent the traditional statistical hypothesis testing procedure as well as the Bayesian one. These are the J(1,5) test devised by Park and Choi (1988) and the Sims' Bayesian odds ratio test (1988) with a small adjustment. Below, I discuss them briefly. Also, given evidence by Diebold and Rudebusch (1991) that typical classical statistical unit root tests such as the Dickey-Fuller one fail to reject a unit root even if the true data-generating process is fractionally integrated, I use the Geweke-Porter-Hudak (1985) spectral regressions technique to estimate the fractional differencing parameter for each of the series. Table-XIII sheds light on the number of co-integrating relationships using Johansen's test and table-XIV into the impulse-response function of each exchange rate following a unit shock.

The J(p,q) test, where p is the order of the time polynomial in the null hypothesis and q is the time polynomial in the fitted regression, is used to reject or not the null hypothesis here that the series possesses a unit root around a p-th order polynomial time trend. The same test can be formulated in a multivariate context under the null hypothesis of co-integration or no co-integration. The J(1,5) test was chosen for a number of reasons. First, no selection of the order of the AR process is necessary. It is known that parametric tests such as the Dickey-Fuller (1979) and the Said and Dickey (1984) depend on the chosen AR order. Second, no estimate of the long-run variance is required. Third, Monte-Carlo experiments performed by Park and Choi (1988) provide evidence that the J(1,5) test is superior, in terms of size adjusted power, to the Augmented Dickey-Fuller and Phillips and Perron tests in small samples and suffers from less size distortion than the Phillips and Perron test (1988).

It should be noted however that it is impossible to distinguish a unit root from a stationary process in small samples because either process can be approximated closely by the other. So, the choice of the deterministic trends is critical to the traditional difference-stationarity tests as pointed out by Campbell and Perron (1991). But there is no guidance as how to go about handling the deterministic trend and this has led Christiano and Eichenbaum (1989) to suggest that we forget about testing for unit roots. To make things worse, Perron (1989) showed that the conventional tests can be fooled into not rejecting a unit root in certain time series if we do not allow for a "breaking trend" or any other non-linear trend that might be indeed present. Ghysels and Perron (1990) have also found out that seasonal adjustment methods induce some bias against rejecting the difference stationary hypothesis.

To complement the classical approach, I also ran unit root tests under the Bayesian tradition. The Bayesian tests are concerned with estimating the probability of the unit root hypothesis relative to the alternative trend-stationary. The posterior odds ratio is the yardstick that compares the competing hypotheses and is defined as:

$$P_{12} = \frac{P(H_1)\ P(Data|H_1)}{P(H_2)\ P(Data|H_2)}$$

Assigning a 50% probability for each hypothesis, the posterior odds ratio becomes:

$$P_{12} = \frac{P(Data|H_1)}{P(Data|H_2)}$$

In other words, the posterior odds ratio gives the probability that the observations would have occured on the condition that H_1 were true versus the probability that the observations would have occured if H_2 were true. Each conditional probability incorporates the prior and the likelihood function. The odds ratio test developed by Sims (1988) is employed with a small adjustment to include a log(2phi) term missing from Sims' equation. The values of the test statistic are not shown but the value at which the posterior odds ratio for and against the unit root are even, called Marginal Alpha, in the tables are reported. If the number is relatively large, this is interpreted as evidence in favor of the unit root hypothesis. Unfortunately, even this family of tests is not immune to criticism. The Bayesian tests based on the calculation of posterior odd ratios seem to be sensitive to the selection of the priors. Informative priors render the tests non-objective and non-informative priors make the tests favor sharp null hypotheses.

D. EMPIRICAL RESULTS

From the unit root tests results for levels reported in table-XI, one can point out that the J(1,5) test fails to reject the null of stationarity at the 5% level for the Portuguese escudo, the ecu, the Belgian franc and the French frank. At the 1% level of significance, the French frank is the only currency that continues to support the null of stationarity around a linear time trend. On the other hand, Sims' Bayesian test favors the difference stationary hypothesis only for the British pound, the Irish punt and the Italian lira. The results changed very little when the I reduced the prior probability on the stationary values of the relevant coefficient from 0.8 to 0.5 and the lower limit for the stationary prior from 0.5 to 0.3 . Finally, the estimate of the fractional differencing parameter based on the Geweke-Porter-Hudak technique is close to 1, the number is 0.9934, only for the Irish punt. From the other currencies, only the British pound and the Italian lira have differencing parameters above 0.9, that is 0.932 and 0.930 respectively. Judging from the results of table-1, I am tempted to conclude that the Irish punt, the British pound and the Italian lira exchange rates must be driven by unit root processes. It is no surprise that the pound was allowed to float against the other currencies early on in the sample at hand while the lira and the punt were devalued vis-a-vis the DM during the same period. However, the sample is relatively small containing just 123 observations and therefore one should be very cautious in drawing conclusions about the presence of unit roots from such a sample. Also, the observations are London closing prices and there is a problem of synchronous recording.

Table-XII presents evidence for difference stationary processes of the 100 times the first log differences of the ten exchange rate series examined here. All tests agree that the differences are stationary, something which is in accord with findings of past studies focusing on dollar exchange rates during the floating exchange rate regime. Note that the Bayesian marginal alpha takes the value of 0 for all series. The J(1,5) test fails to reject the null of stationarity at the 1%, 5% and 10% levels.

Finally, table-XIII reports the value for skewness, kurtosis and the shape parameter of the GEV distribution for exchange rate changes. The values of the other two GEV parameters, namely the location and the scale, are not included because they are very close to their sample counterparts and are uninteresting. The skewness figures are small except for the Irish punt which is -4.825. On the other hand,the kurtosis numbers are large especially for the Irish punt, the Danish krone, the Spanish peseta and the French franc. The respective numbers being 39.870, 16.139, 16.357 and 53.918 . This leads me to conclude that daily EMS returns in the period under study had an excess amount of kurtosis which fits well with the stylized facts about high frequency exchange rate changes. The most important result of table-3 is that the shape parameter of the generalized extreme-value distribution is 0 for all rates except the DM/British pound series which has a maximum likelihood estimate of 0.9982. It is known that the pound floated against the DM starting in the middle of September 1992 and the Italian lira dropped out a little bit later while all the other currencies stayed in the system despite reallignments by the peseta, the escudo and the punt during this time span. However, it should be pointed out that the Italian authorities did not take the same approach as the British, namely to reduce drastically their interest rates and allow their national currency to be exposed to the market winds. The positive number for the shape parameter is indicative of a Fisher-Tippett II distribution and should lead one to conclude that the changes in the DM/pound series are characterized by a non-normal distribution. The same cannot be said about the other series since more or less behaved as if they belonged to the fixed exchange rate mechanism and this might explain the fact that the maximum likelihood estimate of the shape parameter is equal to 0 which characterizes the Fisher-Tippett I distribution.

In chapter-III, I discussed the multivariate unit root test by Johansen (1988). Here I search for unit roots in the data set using the maximimum eigenvalue statistics generated by his test to determine the number of co-integrating vectors in the system. The major problem is the fact that I had 10 time series at hand whereas the tabulated critical values for the statistic depend on the number of stochastic common trends and cannot be less than 1 or more than 5. This limits the room for maneuver since the minimum number of co-integrating vectors I could assign under the null is 5 while still utilizing the full system information. In table-XIV, I report the results of these tests with the number of vectors in the null ranging from 5 to 9. The same procedure is tried on different sub-systems without leading me to reconsider my conclusions.

Finally, Figures-1 through 10 graph the impulse-response function of each of the series on levels in another attempt to find a clue about unit roots. To accomplish that the information criteria for lag length selection of Akaike (1974) and Schwarz (1978) were used to find out the order of the autoregression. The whole idea is to unit shock each price and look at the impulse-response function of the variable subsequently. A unit root process should have a non-zero limit in its impulse-response function. However, univariate shocks can be different than the true shocks. Hansen and Sargent (1991) refer to the interpretation difficulties inherent in VARs due to the nature of the shocks. The order of autoregression is found to be three for the guilder, krone, peseta and Belgian franc on the basis of the AIC and Schwarz criteria. The ecu and the pound have an order of two whereas the French franc, the lira and the punt have one. The exchange rates that come close to have a non-zero limit in their impulse-response functions are the DM/punt first and the DM/pound second. These are the same rates for which all other tests suggest the presence of a unit root. The remaining DM rates have their responses converge to zero.

TABLE-XI: UNIVARIATE UNIT ROOTS TESTS ON EMS LOG LEVELS

Exchange Rates in (DM/others)	J(1,5) where H_0:I(0)	Bayesian Marginal Alpha	Fractional Differencing Parameter
UK	5.154	0.603	0.932
IRL	2.193	0.831	0.993
HOLLAND	0.565	0.000	0.884
BELGIUM	0.181*	0.000	0.595
DENMARK	3.940	0.000	0.222
PORTUGAL	0.183*	0.035	0.852
SPAIN	3.940	0.259	0.936
ITALY	1.027	0.813	0.929
FRANCE	0.098*	0.000	0.123
ECU	1.253	0.158	0.740

Notes: The asterisk denotes that the test fails to reject the null of stationarity at the 5% significance level.
A high figure for the Marginal Alpha favors the non-stationarity hypothesis.
The fractional differencing parameter estimates are based on the Geweke-Porter-Hudak method of spectral regressions.
All numbers have been rounded off to three decimal points.

TABLE-XII: UNIVARIATE UNIT ROOTS TESTS ON EMS
100 TIMES THE LOG OF 1ST DIFFERENCES

Exchange Rates in (DM/others)	J(1,5) where H_0:I(0)	Bayesian Marginal Alpha	Fractional Differencing Parameter
UK	0.064	0.000	0.416
IRL	0.028	0.000	0.030
HOLLAND	0.004	0.000	-0.210
BELGIUM	0.002	0.000	-0.424
DENMARK	0.002	0.000	-0.448
PORTUGAL	0.002	0.000	-0.521
SPAIN	0.016	0.000	-0.039
ITALY	0.069	0.000	0.032
FRANCE	0.000	0.000	-0.722
ECU	0.010	0.000	-0.156

Notes: The asterisk denotes that the test fails to reject the null of stationarity at the 5% significance level.
A high figure for the Marginal Alpha favors the non-stationary hypothesis.
The fractional differencing parameter estimates are based on the Geweke-Porter-Hudak method of spectral regressions.
All numbers have been rounded off to three decimal points.

TABLE-XIII: MAXIMUM LIKELIHOOD ESTIMATION OF THE EXTREME VALUE PARAMETER

Exchange Rates in (DM/others)	Skewness	Kurtosis	Extreme Value Parameter
UK	-0.699	4.579	0.998
IRL	-4.825	39.870	0.000
HOLLAND	0.310	2.666	0.000
BELGIUM	-0.028	2.183	0.000
DENMARK	0.369	16.139	0.000
PORTUGAL	-0.475	1.980	0.000
SPAIN	0.319	16.357	0.000
ITALY	-1.603	8.999	0.000

TABLE-XIV: MAXIMUM EIGENVALUE STATISTIC

JOHANSEN'S TEST

FOR THE EMS EXCHANGE RATE SERIES: $100*LOG(S_t/S_{t-1})$

H_0: r = co-integrating vectors
H_1: r+1 = co-integrating vectors

SIGN. LEVEL—	r-values	1%	5%	10%
r=5	18.500	13.908	16.011	17.447
r=6	14.253	9.689	11.550	12.714
r=7	9.557	5.694	7.286	8.265
r=8	4.033	2.252	3.360	4.097*
r=9	0.020	0.004	0.079*	0.274*

Notes: An asterisk indicates that the test has failed to reject the H0.
The test allows for no deterministic terms.
The number of lagged difference terms used for estimation is taken to be 2 here.
All numbers heve been rounded off to three decimal points.

Figure-1

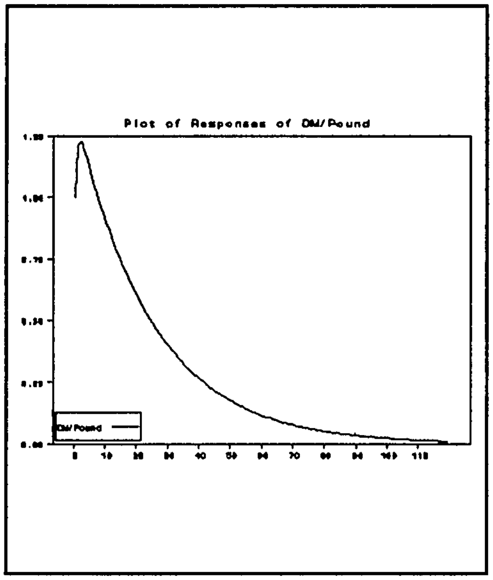


Figure-2

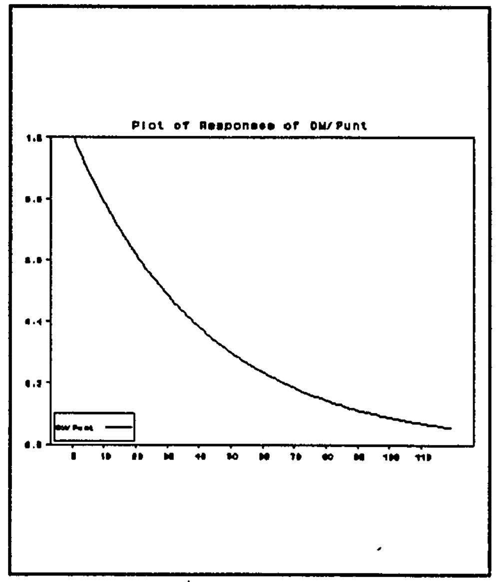
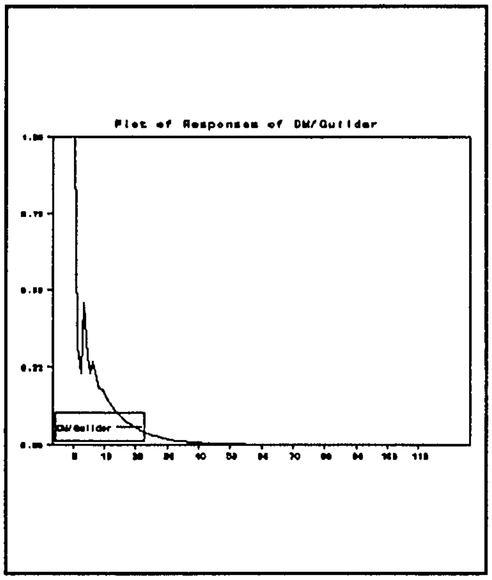


Figure-3



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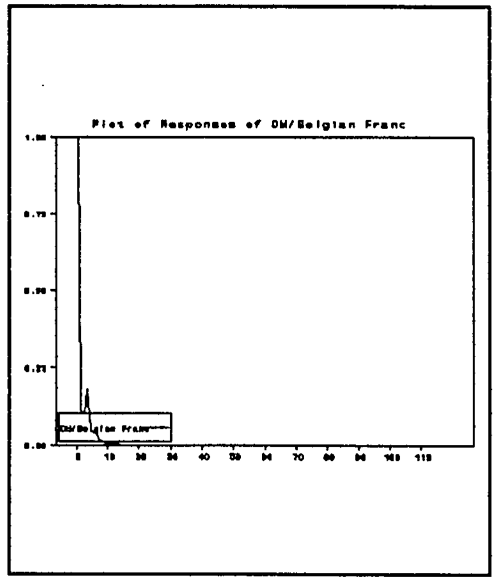

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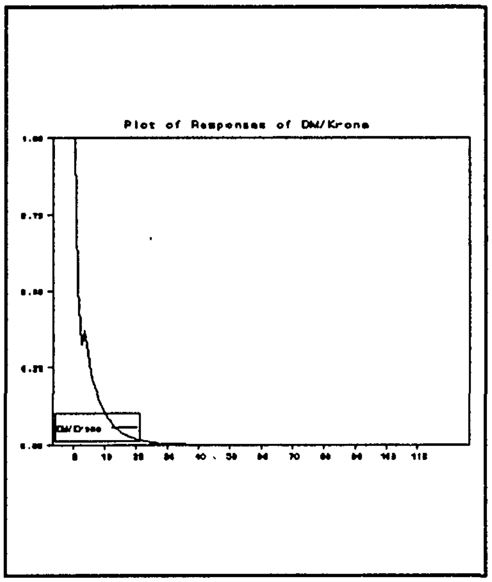
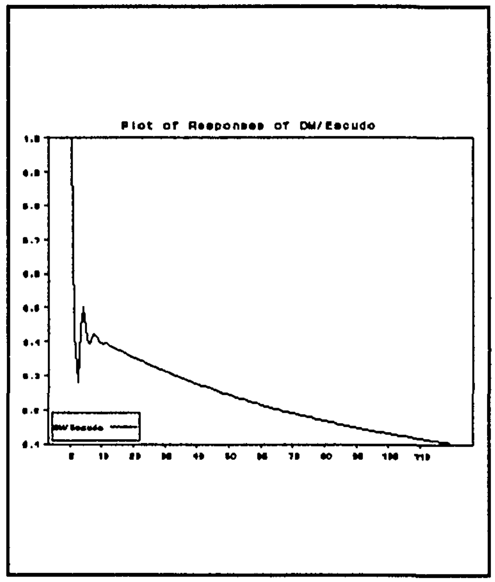


Figure-6



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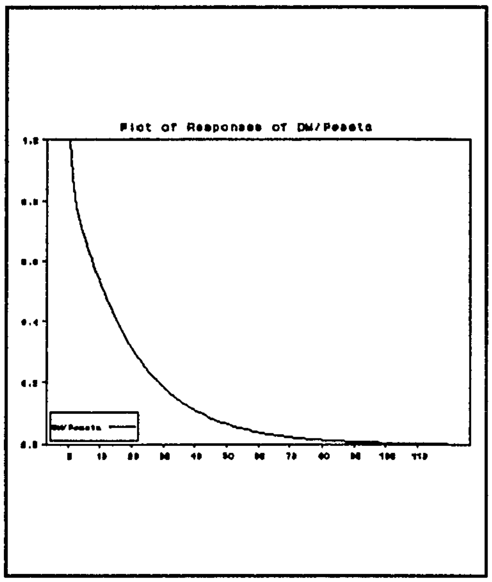

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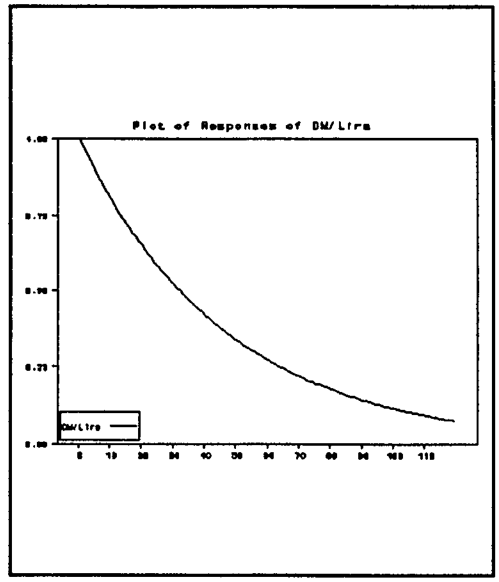


Figure-9

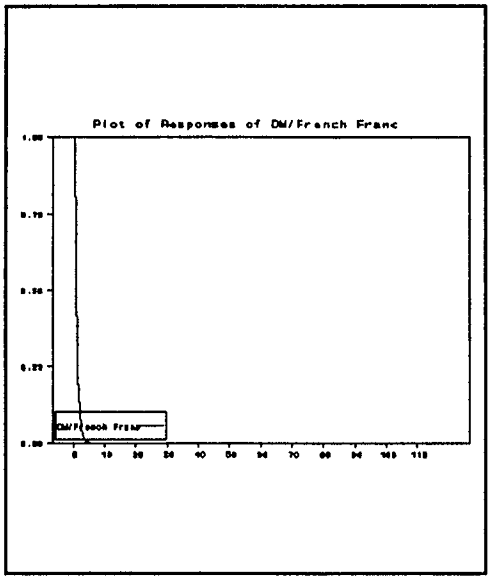
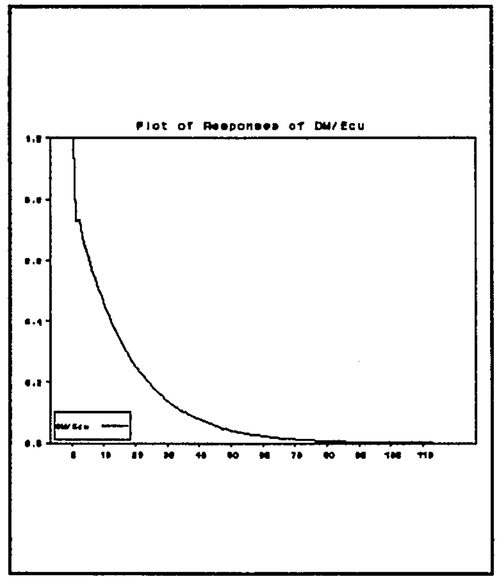


Figure-10



E. CONCLUSION

On the basis of the statistical evidence presented in this essay, the following conclusions are drawn:
First, all unit root tests agree that exchange rate changes of the DM vis-a-vis the other currencies in the EMS during the period August 20, 1992 to February 18, 1993 were stationary. The same univariate tests appear to give conflicting evidence as to whether a unit root exists in their log level representations. However, they are all in accord that the DM/punt, DM/pound and DM/lira series are difference stationary. The limit of the impulse-response function of the DM/punt rate seems to be non-zero implying a unit root process. This is contrary to the other series which seem to have a zero limit in their impulse-response function. Further support for the presence of a unit root in the DM/punt level series is given by the computed value of the fractional differencing parameter which is extremely close to one.
Second, the maximum likelihood estimate of the shape parameter of Jenkinson's generalized extreme-value distribution is different than 0 for the DM/pound first difference series. This clearly demonstrates that the distribution of the particular time series is of the non-normal type.

It should be pointed out again that the sample at hand is relatively small and the quality of the data suspect. This means that the above conclusions should be viewed more as temporary answers to complex questions rather than as definite ones.

V. AN APPLICATION OF THE MARKOV SWITCHING MODEL WITH HAMILTON'S EM ALGORITHM ON THE FINEX DOLLAR INDEX

A. INTRODUCTION

In a fascinating paper written by Engle and Hamilton (1990), the dollar vis-a-vis the mark, the pound and the French franc exchange rates found to be predictable since they were well described by stochastic segmented trends. The random walk model was rejected. In this essay, I apply the simple, two-state Hamilton's model (1988a, 1988b, 1989) without assuming any other autoregressive dynamics to daily opening price data for the dollar index futures contract traded at FINEX. This U.S. dollar index is a measure of the dollar's composite value against other currencies and it is actively being traded for quite some time. The idea is to see whether or not the Markov model can be used to describe the time series behavior of the opening prices of the index. Also, unit root tests are being performed on the data both on levels and the 100 times the log of first differences of the opening and closing prices. The results from the unit root tests on the differences are not being reported because they are uninteresting. They do not conform to the unit root hypothesis. I also supply the results of the Wald test devised by Engle and Hamilton (1990) to discriminate between a martingale and a segmented time trends.

Goldfeld and Quandt (1973) first attempted to model non-linearities arising out of discreet shifts in regime. Their work though did not allow for endogenous state determination. Models to allow for endogenous determination of states in the Markov switching regression tradition have been developed by Diebold and Rudebusch (1987) and Neftci (1984) as well. Neftci's model referred to the U.S. GNP but the transition between the two states was the outcome of a second-order Markov process. Unemployment was the index marking the state of the economy.

B. THE MODEL

According to Hamilton (1989), a time series can be decomposed into two parts. One part follows a random walk with drift and evolves according to a two-state Markov process and the other follows an autoregressive process with a unit root. The two components are unobserved and independent of each other. The general model takes the form in levels:

$$Y_t = X_t + Z_t$$
$$X_t = X_{t-1} + a_0 + a_1 S_t$$
$$Z_t = b_1 Z_{t-1} + b_2 Z_{t-2} + ...b_r Z_{t-r} + e_t$$

In this essay, I simplify the model by focusing on the simple two-state model without any autoregressive dynamics. The transition between states is the outcome of a first-order Markov process of the form:

$$Prob[S_t = 1 | S_{t-1} = 1] = P11$$
$$Prob[S_t = 2 | S_{t-1} = 1] = P21$$
$$Prob[S_t = 1 | S_{t-1} = 2] = P12$$
$$Prob[S_t = 2 | S_{t-1} = 1] = P22$$

In addition to this I use Hamilton's EM algorithm in obtaining the maximum likelihood estimates of the model. The advantages of the EM algorithm developed by Dempster, Laird and Rubin (1977) are its numerical robustness which lets it avoid overstepping when the initial parameter values are far away from their maximum likelihood counterparts and finally the fact that no calculation of the log-likelihood function and its Hessian at each iteration is necessary. The disadvantage relates to its slowness in convergence in the neighborhood of the maximizing values. However, Hamilton (1990) points out that the big steps taken at the beginning more than compensate for the latter. Ruud (1988) has an excellent survey of the EM algorithm.

The 100 times the log of first differences in opening prices for the U.S. dollar index are assumed to be normally distributed with mean Mi and variance VAR(i) where i=1,2 denotes the two states. Each state is determined endogenously and depends on past data and state through the previous period's state. This scheme can encompass a number of situations. First, positive or negative mean values point to upward or downward price moves respectively. Second, big or small absolute mean values point to sharp or gradual price moves respectively. Big or small values for each probability, Pii, indicate drawn out and short regime duration. If the values for P11 and P22 are large and the two means have opposite signs, Engle and Hamilton (1990) claim this to be evidence of the long swings hypothesis. I modify this a bit here saying that large probability values by P11 and P22 are being viewed as evidence for long swings.

The model is quite parsimonious since the parameters of interest are the two means, the two variances and the two probabilities. The expectation and maximization steps of the EM algorithm provide the maximum likelihood estimates of the parameters. The econometrician's task is to draw probabilistic inferences about the unobserved state on the basis of the available information. An initial value for the parameter vector and knowledge of the data can help draw "smoothed" inferences about the probability of a specific regime at a particular point in time. The parameters are being revised consequently in an iterative way until certain convergence criteria are being satisfied. The related formulas are given by Hamilton (1990) and are not being repeated here. Hamilton imposes a pseudo-Bayesian prior for the parameters of the two regimes to avoid certain problems that plague maximum likelihood estimation of Markov switching models. The mean of each regime is shrunk toward zero by including a specific number of additional zero observations from each state and the variance of each regime is being modified toward (b/a) as if one had (2a) observations from each regime. Setting (b/a) too high might not let the data speak for them. Setting (b/a) too low might make P22 be equal to zero.

C. DATA AND ESTIMATION

The data for the closing and opening prices for the U.S. dollar index futures contract come from Genesis Financial Data Services. They cover twelve contracts for the period 1986-1989. Knowing that models such as this one encounter multiple local maxima, I gave twenty different starting values to the parameters for each contract. The estimates turned out to be the same regardless of the initial guesses for all contracts. To avoid the problem of having the likelihood function go to infinity which is possible if the mean of one regime is equal to any observation and the variance of that regime goes to zero, I follow Hamilton (1990) in imposing a Bayesian prior. This helps alleviate the problem with the multiple local maxima as well. The Bayesian prior had a=0.2, b=1.0 and c=0.3 for all contracts. The adjustment takes place in the mean and variance but the probability formulas are left intact. The Wald test suggested by Engle and Hamilton (1990) for the null of a regime at time-t being independent of the regime at time (t-1) is performed on all contracts. The non-linear filter provides both filter and smoothed inference about the regime at time-t. The filter probabilities refer to the probability that the price change was at state-1 using currently available information. The figures shown below refer to this kind of inference. The smoothed probabilities on the other hand are being calculated using all ex post available information to draw an inference about the state at time-t. These are not reported here.

D. EMPIRICAL RESULTS

As noted before a number of unit root tests on changes in the log of the prices and levels were performed as well. All tests for the log levels are being reported below but not for the changes in the log which conformed to stylized facts.

From the unit root test on opening and closing prices for the March dollar index contract one can observe that the evidence is overwhelmingly for unit roots under the ADF test. The Bayesian test favors the same hypothesis in most cases except for the closing and opening prices of the 1987 contract. The G-H-P fractional differencing parameter is close to a unit for many contracts but exceeds one in the case of the March 1989 opening price contract.

The June contract prices seem to follow unit roots as evidenced by the Said and Dickey's (1984) ADF test. The Bayesian tradition favors the same hypothesis except in the case of the 1987 contract. The G-H-P parameter explodes for all June contracts.

The story is similar to the June contract for the September contract. The fractional differencing parameter takes values above one, the ADF test favors the unit root hypothesis and Sim's (1988) slightly modified test does the same except for the 1987 contract.

The December contract seems to break away from the norm. Although the ADF test still points to unit roots, this is not so for the Bayesian test which favors for the trend-stationary hypothesis in the case of the 1987 contract but also the opening price of the December 1988 one. The G-H-P parameter either displays explosive behavior or takes very low values. Once again, one should view the results of the unit root test with skepticism. In small samples, the difference stationary and trend-stationary alternatives are equivalent.

The MLE of the parameters of the model are more interesting indeed. As far as the March contract is concerned, the means take negative values except for the March 1989 contract which has positive means in both states of the world. However the values taken by P11 and P22 are high except for the 1989 contract. This might be interpreted as evidence in favor of the long swings hypothesis.

The June contracts for the 1986-1988 period display persistence as indicated by high probability values. The mean in regime-2 of the 1987 contract becomes positive as does the mean in regime-1 of the 1988 contract.

The September contracts exhibit long swings as well except for the 1988 contract which has a positive mean for state-2. All other mean values are negative.

The December contracts are characterized by large probability values for all periods with negative means except for the mean in state-2 of the 1988 contract which turns out to be positive.

The Wald test performed on each contract rejected the null hypothesis of a martingale in favor of the stochastic, segmented trends model in all cases except for the March 1989 and September 1988 contracts. This lead me to conclude that long swings might be an integral feature of the opening prices futures data I looked into. Below, I give Engle and Hamilton's (1990) formula used to compute the chi-square(1) statistics. Note that the critical value of chi-square(1) at the 5% level of significance is 3.84 as indicated in the table.

$$\frac{[P11-(1-P22)]^2}{[var(p11)+var(p22)+2cov(p11,p22)} = \chi^2(1)$$

All variances and covariances are estimated from the negative of the matrix of second derivatives of the generalized objective function given by Engle and Hamilton (1990).

TABLE-XV: UNIT ROOT TESTS

Opening Prices of the March FINEX

Dollar Index Contracts: 1987-89

	ADF Test	Bayesian Alpha	Geweke-Hudak Porter
1987	0.94	0.26	0.99
1988	0.97	0.72*	0.91
1989	0.98	0.88*	1.01

Closing Prices of the March FINEX

Dollar Index Contracts: 1987-89

	ADF Test	Bayesian Alpha	Geweke-Hudak Porter
1987	0.94	0.24	0.99
1988	0.97	0.71*	0.93
1989	0.98	0.88*	0.99

Notes: 1. The Dickey-Said's ADF test is formulated on the null of a non-stationary series around a linear trend.

2. The asterisk indicates the Bayesian test favors unit roots.

3. All numbers have been rounded off to two decimal points.

TABLE-XVI: UNIT ROOT TESTS

Opening Prices of the June FINEX

Dollar Index Contracts: 1986-88

	ADF Test	Bayesian Alpha	Geweke-Hudak Porter
1987	0.94	0.48	1.02
1988	0.96	0.66*	1.10
1989	0.98	0.82*	1.22

Closing Prices of the June FINEX

Dollar Index Contracts: 1986-88

	ADF Test	Bayesian Alpha	Geweke-Hudak Porter
1987	0.93	0.39	1.07
1988	0.96	0.67*	1.10
1989	0.98	0.83*	1.25

Notes: 1. The Dickey-Said's ADF test is formulated on the null of a non-stationary series around a linear trend.

2. The asterisk indicates the Bayesian test favors unit roots.

3. All numbers have been rounded off to two decimal points.

TABLE-XVII: UNIT ROOT TESTS

Opening Prices of the September

FINEX Dollar Index Contracts:

1986-88

	ADF Test	Bayesian Alpha	Geweke-Hudak Porter
1987	0.93	0.20	1.01
1988	0.98	0.80*	1.04
1989	1.00*	0.88*	1.20

Closing Prices of the September

FINEX Dollar Index Contracts:

1986-88

	ADF Test	Bayesian Alpha	Geweke-Hudak Porter
1987	0.93	0.15	1.14
1988	0.98	0.81*	1.05
1989	1.00*	0.88*	1.15

Notes: 1. The Dickey-Said's ADF test is formulated on the null of a non-stationary series around a linear trend.

2. The asterisk indicates the Bayesian test favors unit roots.

3. All numbers have been rounded off to two decimal points.

TABLE-XVIII: UNIT ROOT TESTS

Opening Prices of the December

FINEX Dollar Index Contracts:

1986-88

	ADF Test	Bayesian Alpha	Geweke-Hudak Porter
1987	0.96	0.45	1.14
1988	0.97	0.06	0.64
1989	0.98	0.88*	0.87

Closing Prices of the December

FINEX Dollar Index Contracts:

1986-88

	ADF Test	Bayesian Alpha	Geweke-Hudak Porter
1987	0.96	0.36	1.16
1988	0.97	0.69*	0.82
1989	0.98	0.86*	1.08

Notes: 1. The Dickey-Said's ADF test is formulated on the null of a non-stationary series around a linear trend.

2. The asterisk indicates the Bayesian test favors unit roots.

3. All numbers have been rounded off to two decimal points.

TABLE-1
FINEX DOLLAR INDEX CONTRACTS FOR MARCH 1987, 1988 AND 1989. (Std errors in the second row.)

FINEX DOLLAR INDEX	M1	M2	P11	P22	VAR1	VAR2
3/87	-0.10 0.057	-0.03 0.078	0.95 0.04	0.96 0.02	0.21 0.07	0.72 0.11
3/88	-0.009 0.033	-0.123 0.101	0.95 0.027	0.90 0.05	0.154 0.022	0.821 0.151
3/89	0.023 0.03	0.02 0.06	0.53 0.104	0.61 0.107	0.03 0.005	0.52 0.084

TABLE-2
FINEX DOLLAR INDEX CONTRACTS FOR JUNE 1986, 1987 AND 1988. (Std errors in the second row.)

FINEX DOLLAR INDEX	M1	M2	P11	P22	VAR1	VAR2
6/86	-0.09 0.069	-0.07 0.09	0.99 0.013	0.99 0.01	0.231 0.049	0.69 0.10
6/87	-0.106 0.048	0.0001 0.112	0.945 0.03	0.905 0.048	0.184 0.034	0.760 .16
6/88	0.0135 0.031	-0.204 0.138	0.94 0.028	0.822 0.084	0.147 0.018	1.10 0.23

TABLE-3
FINEX DOLLAR INDEX CONTRACTS FOR SEPTEMBER 1986, 1987 AND 1988. (Std errors in the second row.)

FINEX DOLLAR INDEX	M1	M2	P11	P22	VAR1	VAR2
9/86	-0.107 0.057	-0.069 0.087	0.99 0.014	0.98 0.017	0.317 0.057	0.69 0.118
9/87	-0.032 0.0354	-0.086 0.198	0.99 0.012	0.93 0.068	0.236 0.0274	0.949 0.290
9/88	-0.049 0.039	0.073 0.106	0.691 0.097	0.487 0.123	0.112 0.0253	0.896 0.182

TABLE-4
FINEX DOLLAR INDEX CONTRACTS FOR DECEMBER 1987, 1988 AND 1989. (Std. errors in the second rows.)

FINEX DOLLAR INDEX	M1	M2	P11	P22	VAR1	VAR2
12/86	-0.096 0.0507	-0.029 0.0807	0.951 0.035	0.937 0.042	0.249 0.050	0.64 0.12
12/87	-0.023 0.0385	-0.192 0.0934	0.989 0.009	0.9890 .013	0.237 0.027	0.700 .11
12/88	-0.005 0.034	0.121 0.218	0.95 0.035	0.72 0.191	0.205 0.032	1.27 0.47

ALL FILTER PROBABILITY FIGURES BELOW REFER TO STATE-1

FIGURE-11

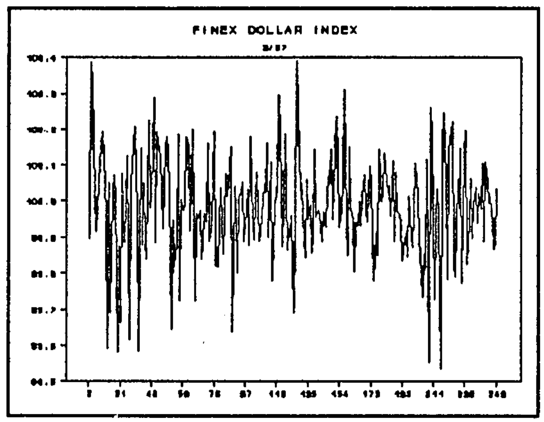


FIGURE-12

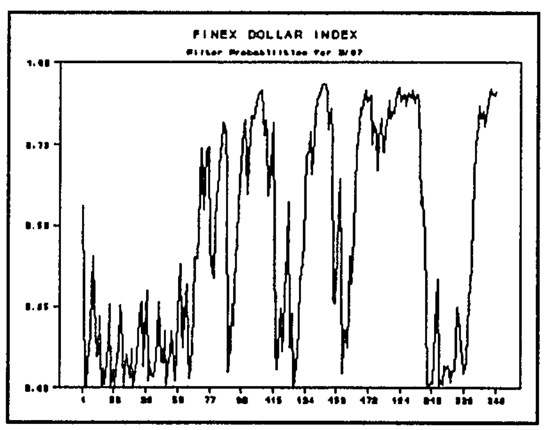


FIGURE-13

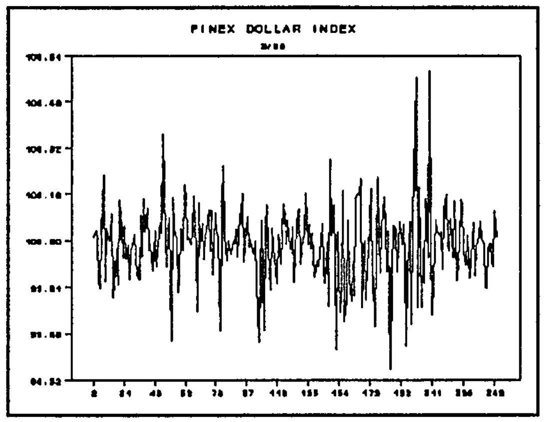


FIGURE-14

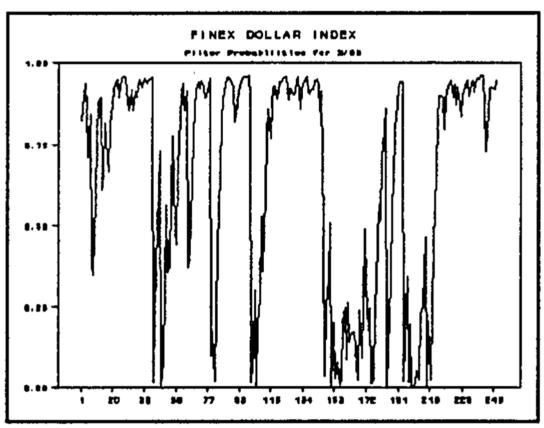


FIGURE-15

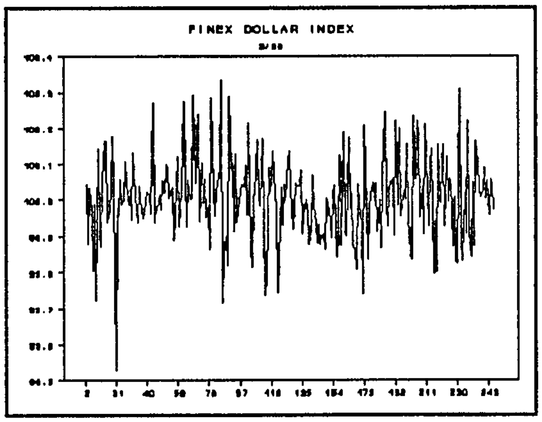


FIGURE-16

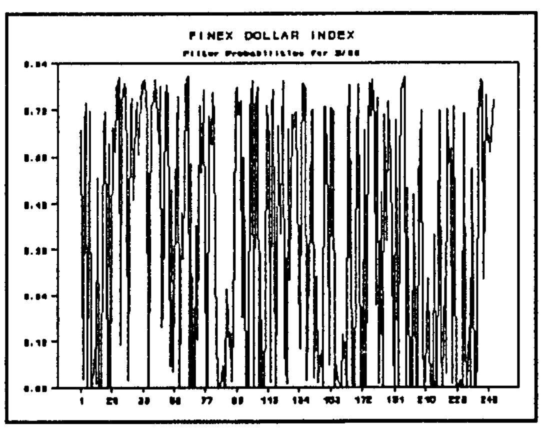


FIGURE-17

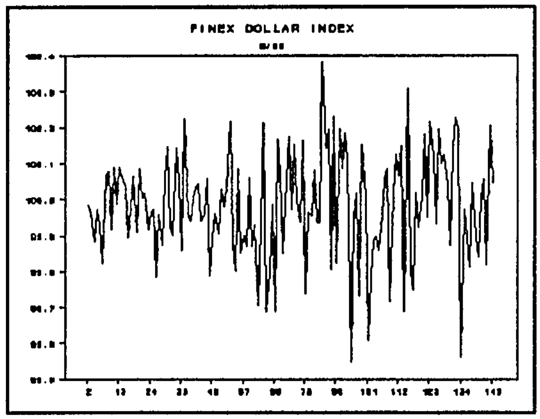


FIGURE-18

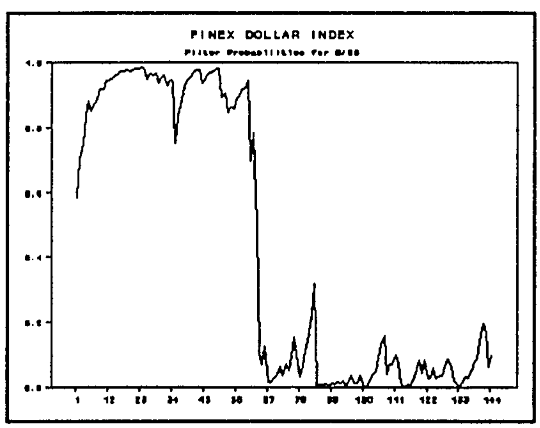


FIGURE-19

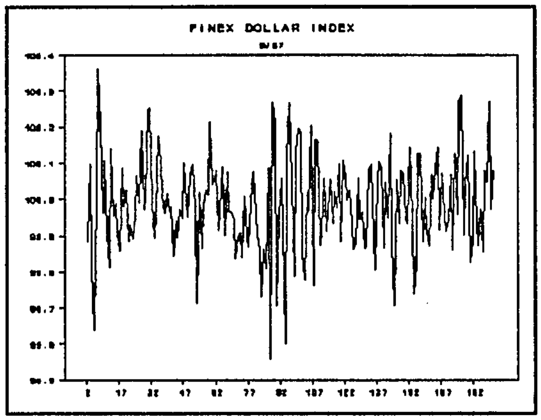


FIGURE-20

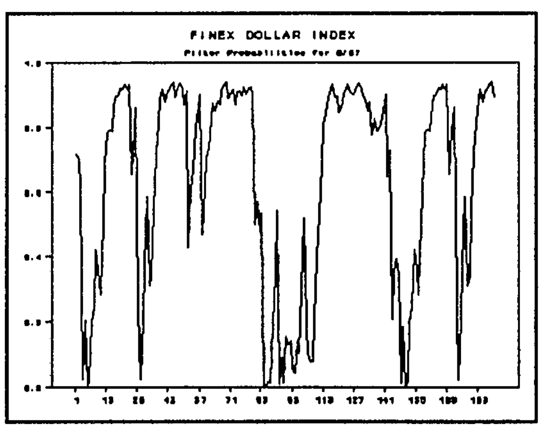
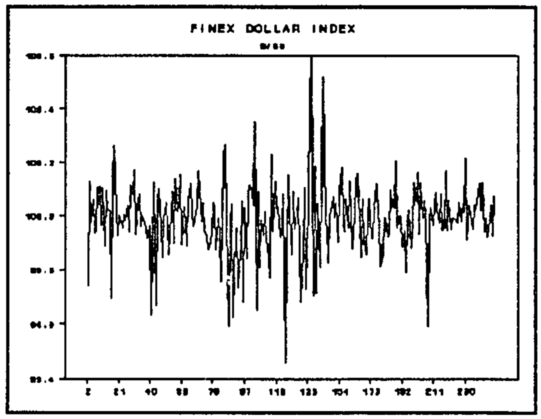


FIGURE-21



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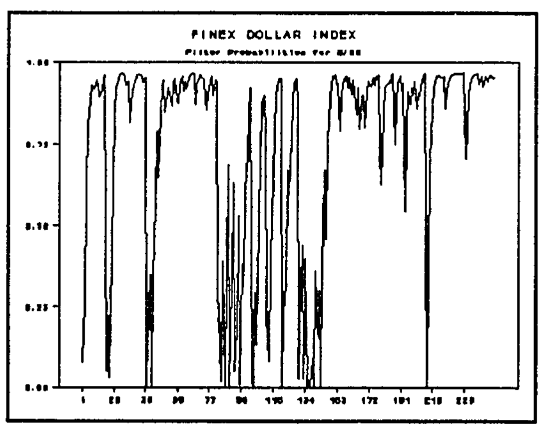


FIGURE-23

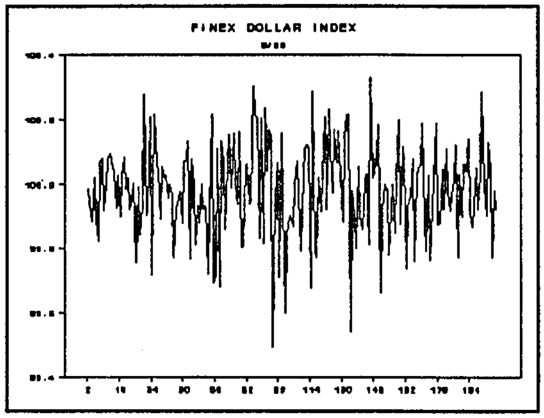


FIGURE-24

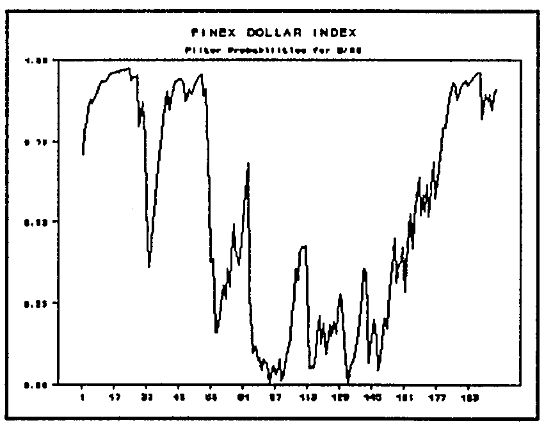


FIGURE-25

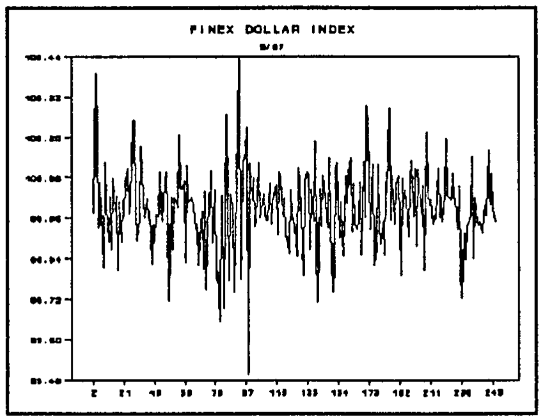
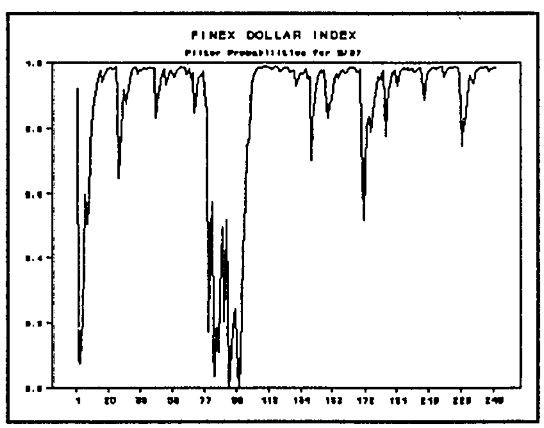


FIGURE-26



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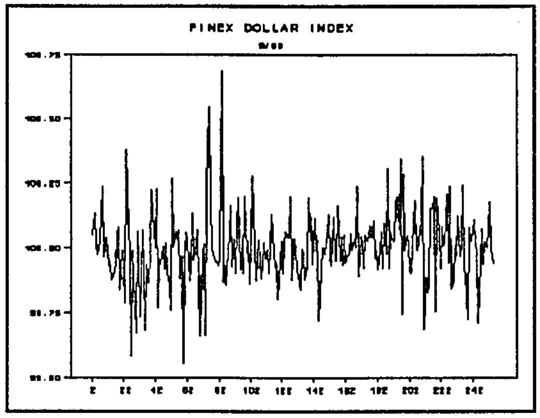

FIGURE-28

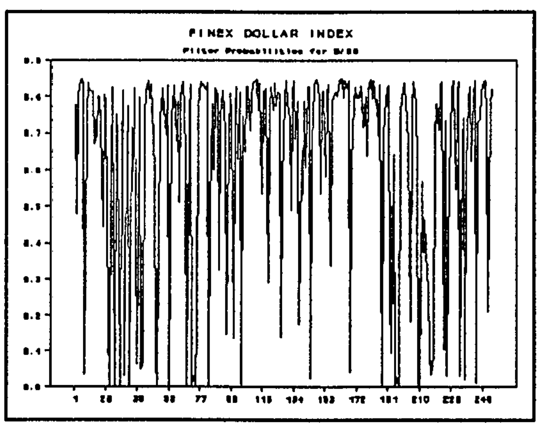

FIGURE-29

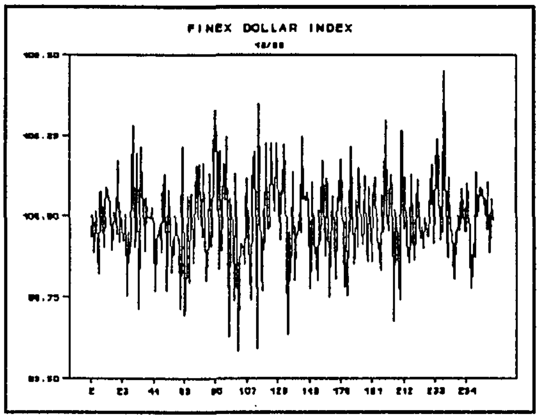

FIGURE-30

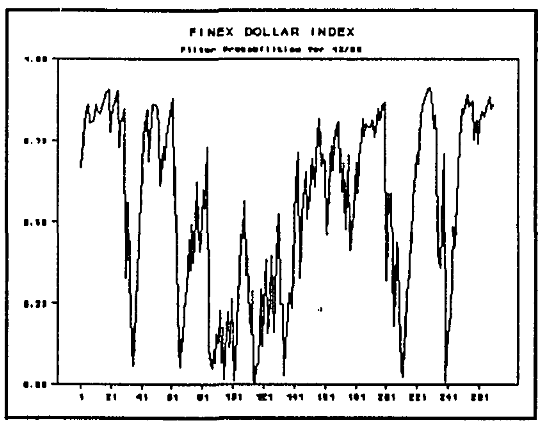

FIGURE-31

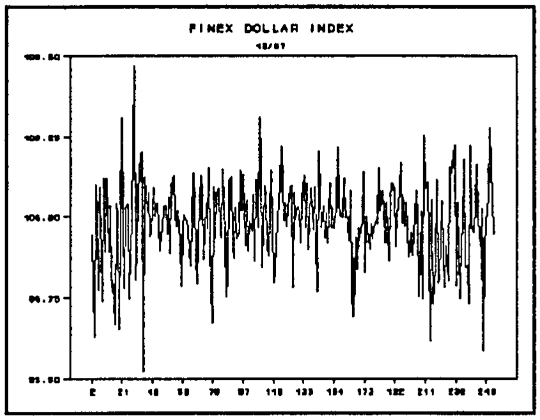

FIGURE-32

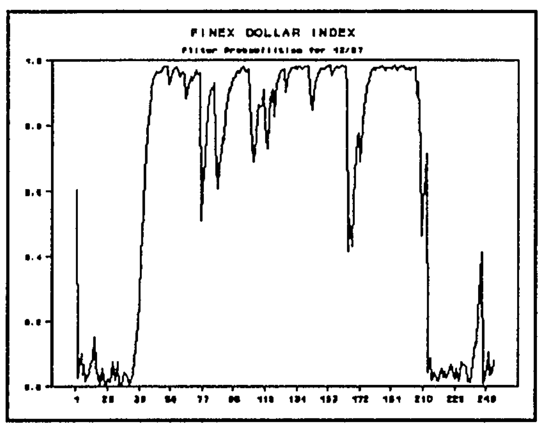

FIGURE-33

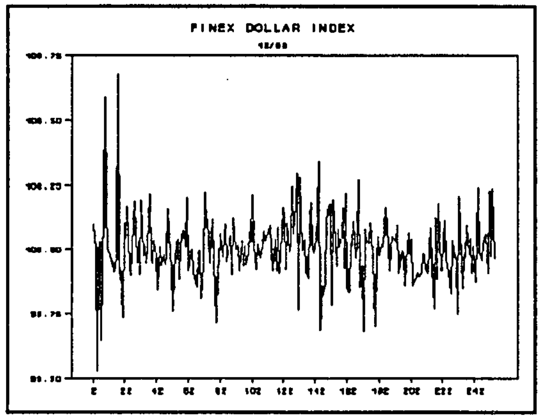

FIGURE-34

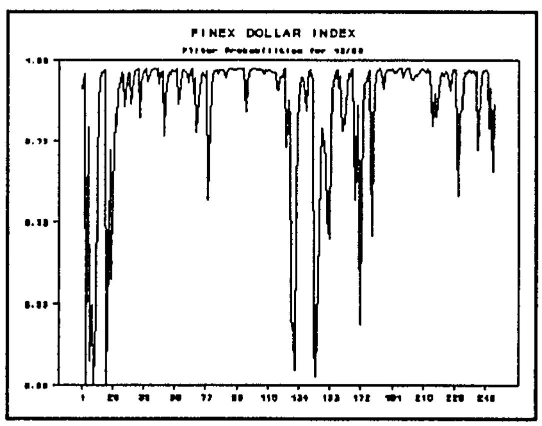

TABLE-XXIII

WALD TEST OF THE NULL HYPOTHESIS THAT THE 100 TIMES THE LOG OF FIRST DIFFERENCES OF THE OPENING PRICES FOLLOW THE MARTINGALE AGAINST THE ALTERNATIVE OF LONG SWINGS

H_o: P1 = 1-P2

Contract	Wald Test
3/87	409.95
3/88	118.44
3/89	0.58 *
6/86	2670.70
6/87	200.70
6/88	57.05
9/86	1270.58
9/87	83.03
9/88	1.09 *
12/86	176.80
12/87	2884.95
12/88	9.82

Note: An asterisk indicates the non-rejection of the null at the 5% level.
The Chi-Square (1) statistic at the 5% level is 3.84

E. CONCLUSION

Based on the evidence reported in this essay, I have to conclude that long swings are in the daily opening price data for the U.S. dollar index contracts examined here. Thus, changes in the log of these prices are forecastable with the notable exception of the March 1989 and the September 1988 contracts.

VI. APPENDIX

A. FIRST ESSAY

The first part of this appendix includes all the intermediate steps along the way to deriving the first-order condition in equation (3). This follows very closely Dornbusch (1983) but in a stochastic environment. Dornbusch's model was developed in a deterministic environment.

To derive the optimal composition of current consumption between tradeable and nontradeable goods we differentiate the expected utility function using the intertemporal budget constraint with respect to consumption of current goods:

(4) $$U_{Nt}/U_{Tt} = p_t$$

where $U_{j,t}$ is the partial derivative with respect to j = N,T arguments denoting nontraded and traded goods.

Repeating the same procedure for time (t+1) and using the equation for the optimal composition of current consumption above leads to:

$$(5) \qquad b\ E_t[U_{Nt+1}/U_{Nt}\ \ (1+r)\ \ (p_t/p_{t+1})] = 1$$

Substituting for the derivatives from the isoelastic functional form of utility with the Cobb-Douglas functional form of consumption, we get:

$$(6) \qquad C_{Nt}/C_{Tt} = [\ (1-a)/a\ ]\ /\ p_t$$

$$(7) \qquad C_{Nt} = [\ (1-a)/a/p_t\ ]^a\ C_t$$

$$(8) \qquad U_{Nt} = (1-a)\ \ C_t^{H-1}\ [\ a/(1-a)\ \ p_t\ ]^a$$

Then equation (5) because of (8) gives the Euler equation (3) in the main text which is being estimated.

The second part of this appendix describes the GMM procedure developed by Hansen (1982) and employed to estimate the Euler equation (3) in the first essay generalized for the case of P assets and N instruments. Then, an Euler equation having the form of (3) defines an error term for each asset e_{it+1} with $E_t\ (ei_{t+1}) = 0$ where $i=1,...,P$ assets and E_t defines the conditional expectation at time t. Let Z_{jt} denote a set of instruments known to all agents at time t where $j=1,...N$. Then, $E\ (e_{t+1}\ \ Z_t'\ ) = 0$ where e_{t+1} is a P dimensional vector of error terms and Z_t is a N dimensional vector of instruments and E is the unconditional expectations operator. The assumption is being made that the P disturbances and the N instruments have finite second moments. This expectational product defines a family of (PxN) orthogonality conditions.

The next step is to construct and minimize an objective quadratic function that takes the form $(g'Wg)$ where -g- is the vector that contains the elements of

$$\frac{1}{T}\sum_{t=1}^{T}(e_{t+1}\,Z_t')$$

where T is the sample size. W is a (PxN) by (PxN) symmetric, positive weighting matrix which is the inverse of a consistent estimate of the covariance matrix of these orthogonality conditions. However this matrix is not always positive definite. Newey and West (1987) suggested the use of an estimator that is robust to autocorrelated and conditionally heteroscedastic errors giving a consistent and positive definite estimate of W. The estimates of the parameters of the above objective function that minimize it are the GMM estimates. Provided that (PxN) is greater than the number of estimated parameters, the theoretical model is overidentified and the restrictions can be tested by using the minimized value of the objective quadratic function as a test statistic distributed as a chi-square with degrees of freedom equal to the difference between the number of orthogonality conditions PN and the number of parameters under the null hypothesis. Even a failure of the test recommended by Hansen should be viewed with caution as pointed out by Rotemberg (1983). This failure might be traced to two reasons. First, irrationality of expectations and model misspecification. Second, the error term might capture shocks in preferences and technology.

B. THIRD ESSAY

The third essay in this dissertation uses extreme order statistics to analyze the tail of the distribution of exchange rate returns from EMS data. The whole analysis is based on the theory of extremes which I briefly refer to in this appendix in a general form. A good reference can be found in Leadbetter, Lindgren and Rootzen (1983) as well as in Mood, Graybill and Boes (1974). Let $Y_1, \ldots Y_n$ be a stationary sequence of i.i.d. random variables whose distribution function is F. The sample size is n. Let also the maximum of these random variables be denoted by

M_n which is also the last order statistic. Then, the probability that M_n is less than some level y is given by

$$P(M_n <= y) = F^n (y)$$

and the finding of the type of the limiting distribution that M_n follows using $F_n(y)$ is the major focus of extreme-value theory.

If G(y) is another distribution function which is tail equivalent to F(y) and

$$\lim_{n\to\infty} F^n(a_n+b_n\, y) = X(y)\,, \quad \textit{for all } y$$

$$\textit{then}$$

$$\lim_{n\to\infty} G^n(a_n+b_n\, y) = X(y)\,, \quad \textit{for all } y$$

Then according to Resnick's theorem (1971), if one of them belongs to some domain of attraction so does the other and the normalizing constants are the same. In other words, we can use G(y) which might be a simple distribution instead of F(y) without changing anything else. This way we can estimate the shape parameter and the tail index which is the reciprocal of the latter. It should be pointed out that in this essay there is an implicit assumption that each period's extreme value follows one of the three limit laws exactly. This is a drawback for the the maximum-estimation approach adopted here.

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