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International Reserves Revisited: Long-Run Determinants and Short-Run Dynamics after Bretton Woods.

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2000

Submitted for the Degree of Doctor of Philosophy

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Acknowledgements

I wish to thank my supervisor Professor Ronald MacDonald for considerable help in the preparation of this thesis. For his encouragement with ideas, knowledge of theory and methods, I am very grateful. I am also appreciative of my external examiner Professor Leigh Drake and internal Dr Nicola Viegi, for a rigorous, fair and rewarding viva. I also wish to thank Professor Andrew Hughes Hallett and Professor Jim Love for help with my ESRC funding application. I am indebted to Ian Marsh, Andrew Davies, Norbert Fiess, Kevin Ryce and others at Strathclyde University Economics Department with help reading drafts, information on methods, software and computer programs. Also, I have benefited from comments from seminar participants at Strathclyde and Crieff Hydro, in particular, from Professor Anton Muscatelli.

Finally I wish to thank my family for support, prayers and encouragement. I am most grateful to my mother and father.

Abstract

This thesis examines a number of issues related to central bank international reserves holdings and foreign exchange intervention. We study the long run determinants of reserves within the context of the post Bretton Woods dirty float period. It is argued that traditional approaches fail to take account of central bank attempts to influence the real exchange rate by foreign exchange intervention. Additionally, we update previous research by employing recent developments in the non-stationary time-series and panel data literature. In particular, we utilise the Johansen VAR technology and recent innovations in panel cointegration, to assess the long-run determinants of reserves and short-run dynamics. By jointly modelling the UK reserve holdings and the monetary sector we consider the domestic economy impact of reserve changes, the stability of narrow money demand and whether monetary disequilibria effect reserves as suggested by the Monetary Approach to the Balance of Payments. The effects of daily US and German foreign exchange intervention on exchange rate volatility are also studied. We find evidence consistent with other research that US intervention reduces volatility and extend these results to bilateral rates not previously considered. Moreover, we find evidence in favour of the distinction between unilateral and concerted intervention and of the existence of policy externalities, underlining the importance of international policy coordination.

Chapter One

Introduction

The most significant change in the International Monetary System over the past three decades has been the collapse of the Bretton Woods system of fixed exchange rates. This fundamentally altered the way issues were examined in the literature on international economics. The change was precipitated by financial crisis but the move to a more widespread system of flexible exchange rates had been advocated for a number of years before the crisis (see Friedman, 1953; Johnson, 1970; and Dunn, 1983). The benefits which could be incurred by a move to flexible exchange rates were argued to include (amongst other reasons): floating rates would promote economic stability by allowing shocks to effect exchange rates and not other macroeconomic variables; and that private speculation would be stabilising.

It was suggested that, floating exchange rates would ensure balance of payments equilibrium and obviate the need for international reserve holdings. Although reserves are of some importance within a fixed exchange rate regime they were not predicted to be so with flexible rates. Under a fixed exchange rate system, an overvalued exchange rate leads to an excess supply of domestic goods, an external payments deficit and thereby an outflow of reserves. This fall in domestic reserve holdings, assuming a typical central bank balance sheet, will remove the payments deficit. It was postulated however, that a change in exchange rate regime should be associated with a change in reserve holdings. For example, Haberler (1977) argued that the adoption of a floating rates regime would lead to a fall in the demand for reserves since there is greater adjustment by changes in the exchange rate. This argument can be traced back to Friedman (1953) who proposed that reserve economies were one of the primary arguments in favour of a move to flexible exchange rates.

Despite the move to floating exchange rates after the collapse of Bretton Woods, Williamson (1976) found evidence that reserves continued to be held on a similar scale as before. Indeed Mussa et al. (1994) suggested holdings have in fact increased. Table 1.1, which contains the level of reserves as a fraction of real GDP for the UK from 1960 to 1995, is consistent with this trend.

Table 1.1: UK Foreign Exchange Reserves

Year	Reserves as a Percentage of Real GDP
1960Q1	1.125
1965Q1	0.758
1970Q1	0.789
1975Q1	1.504
1980Q1	3.839
1985Q1	2.217
1990Q1	4.858
1995Q1	4.815

Source: IMF International Financial Statistics.

Taylor (1982) is of the view that reserve holdings did not fall after the 1970s, as the central banks of the industrialised countries have engaged in foreign exchange intervention. One of the most prominent features of the recent floating exchange rate experience has been the high level of foreign exchange market intervention. This has been as high as, if not higher than, under the Bretton Woods system (see MacDonald, 1988). Despite the end of the explicitly fixed exchange rate regime a number of countries continue to engage in a range of formal and informal exchange rate arrangements. They may peg some of their rates but have others that are flexible and yet others where the monetary authorities engage in foreign exchange intervention. For example, while Britain was a member of the Exchange Rate Mechanism, pegging her currency to other members, it allowed non-ERM bilateral rates to remain flexible. Germany has been a formal member of the ERM and engaged in foreign exchange intervention against the US Dollar - particularly after the Plaza Agreement (see Figure 1) and Louvre Accord (see Edison, 1993). At the same time Germany has allowed many other Deutsche Mark bilateral rates to float freely. Viewing reserve demand within the context of foreign exchange intervention has not been extensively examined within the literature. But given that reserves and intervention are inextricably linked, we believe the literature requires re-appraisal. These arguments are a first stop in this thesis, which examines a number of issues related to international reserves and their interaction with other macroeconomic fundamentals.

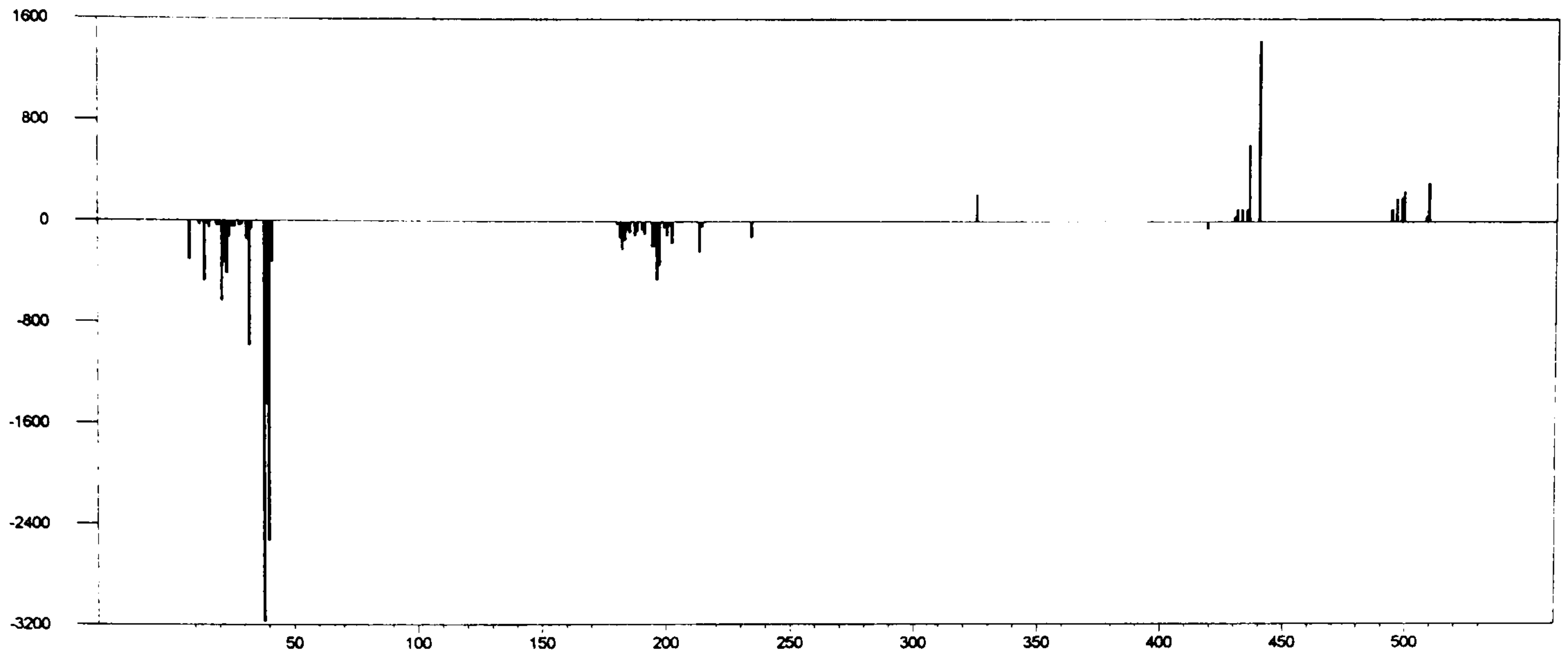

Figure 1.1 Bundesbank Daily Foreign Exchange Intervention in Millions of US Dollars, January 1985 to February 1987. Positive values represent Dollar purchases. *Source:* Deutsche Bundesbank, Frankfurt.

In particular, we examine the long run determinants of reserves and attempt to define an alternative long-run specification to traditional studies. These traditional studies have considered reserves demand within the context of the balance of payments and a fixed exchange rate. Reviewing this argument we examine a number of hypotheses related to reserves. For example, we consider the relationship between reserves and the exchange rate within a structural non-stationary Vector AutoRegression model. Also, by modeling the monetary sector we are able to assess the relationship between reserves and the wider monetary situation in an open economy context. In addition to these matters, which primarily concern themselves with the level of central bank foreign currency holdings and the level of the exchange rate, we examine some issues related to high frequency changes in reserves and the variance of the exchange rate. That is to say, we are interested in impact of daily foreign exchange intervention (see Figures 1.1 and 1.2) on nominal exchange rate volatility.

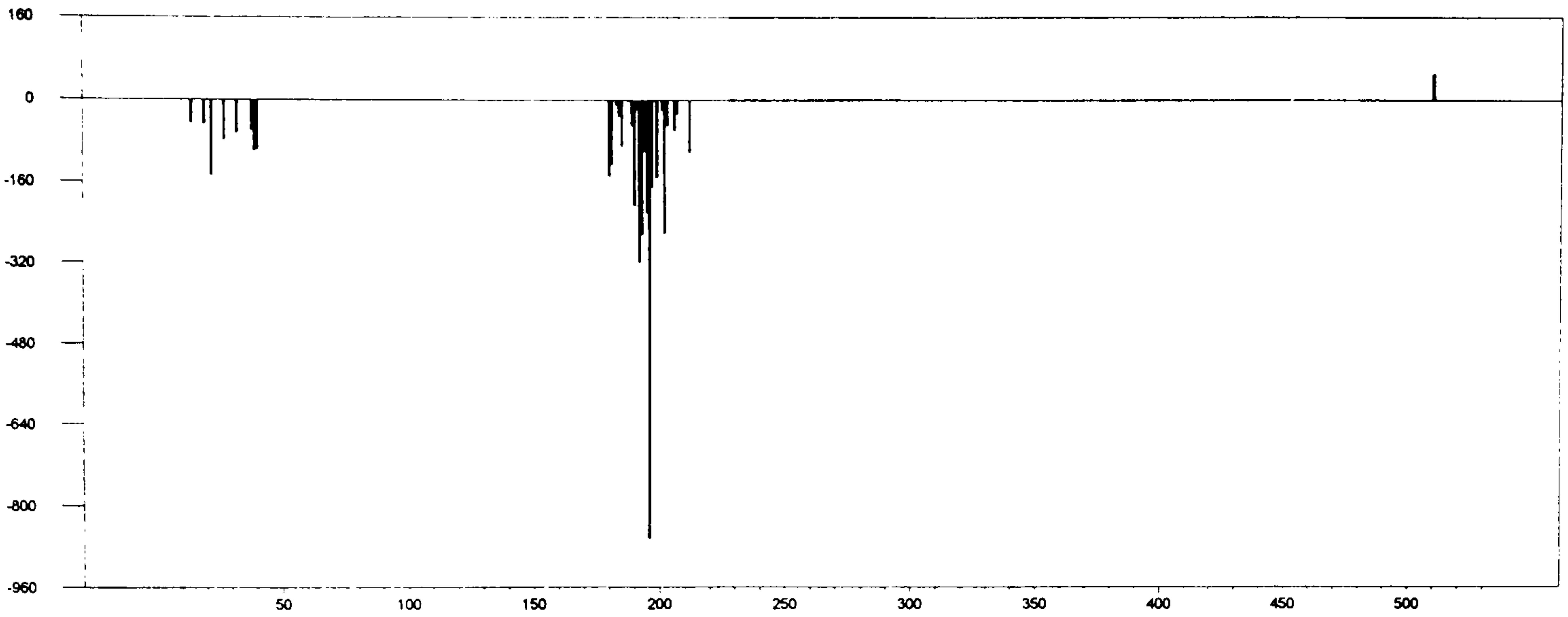

Figure 1.2 Federal Reserve Daily Foreign Exchange Intervention in Millions of US Dollars, January 1985 to February 1987. Positive values are Dollar purchases. *Source:* Federal Reserve Bank of New York.

In the literature, there are two main approaches to the analysis of reserve holdings: the descriptive approach; and the optimising approach.[1] We concentrate on the former. The first examination of the determinants of reserves within the descriptive approach was Triffin (1947). He suggested that the level of imports was a central determinant of reserve demand, with the reserve-import relationship the important measure of reserve adequacy. This approach was widely adopted by, for example, Harrod (1953) and Grubel (1965). However this simple method was criticised by Kenen and Yudin (1965) and Machlup (1966). This criticism emphasised that it was not the level of imports that mattered for reserves, but that a measure of the variability of payments was important. Since reserves were used as a buffer stock to accommodate fluctuations in external transactions, reserves will be positively associated with the extent of these

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[1] The optimising approach typically utilises a maximisation procedure. For references, see Jung (1995). Although the descriptive and optimising approaches are not entirely distinct, since they employ similar determinant variables.

fluctuations. Kenen and Yudin (1965), using the estimated variance from the residuals of a simple Markov process of the first difference of reserves on its lag, found consistent evidence that the level of reserves was related to reserve volatility. Kenen and Yudin (1965) also adopted regression techniques, which represented a move away from Triffin ratios and introduced the possibility of additional explanatory variables being included in a static reserve equation. The primary research effort focused on which are the key variables to include as determinants of reserves.

Heller (1966) suggested that the propensity to import was also a crucial determinant of reserve holdings, based on the Keynesian model of foreign trade multipliers. External disequilibrium, because of a fall in export earnings, could be corrected by a fall in output proportional to the trade multipliers. This fall in output could be avoided if the monetary authority ran down the stock of international reserves in response to external disequilibrium. As the foreign trade multiplier is inversely related to the Marginal Propensity to Import (MP), the cost of not having reserves, and hence the demand for reserves, is related to MP. Hipple (1974) took an alternative view. He suggested that a country's MP is a reflection of its degree of openness and therefore it is a measure external shocks. An increase in openness will require an increase in reserve holdings to accommodate economic shocks. This argument was widely adopted and considered as a monetarist view. The primary difference between Heller's (1966) and Hipple's (1974) approaches, is that the predicted sign of the MP's estimated coefficient in a reserve regression is different. Heller suggested it would be negative, whereas within the monetarist approach it is expected to be positive. MP is unfortunately an unobservable variable, and it is typically proxied by the Average Propensity to Import.

What of the recent empirical evidence on the long-run determinants of reserves? Frenkel (1983) used Ordinary Least Squares and a pooled data set of 22 developed countries, with country dummies, to estimate a traditional specification. He found evidence of diseconomies with respect to the estimated elasticity of income (e.g. the coefficient was greater than one). Volatility and the Average Propensity to Import (AP) were also positive and significant. Edwards (1984) found a significant role for income and lagged reserves using a Least Square Dummy Variable (LSDV) approach for a pooled sample of 23 developing countries. However he did not find evidence of a significant role for volatility or the AP with his sample of 1965-1972. These studies have been supplanted by the approaches of Elbadawi (1990) and Ford and Huang (1994) who take account of more recent developments in the single equation literature which deals with potentially non-stationary data. Elbadawi (1990) considers Sudanese demand for reserves with quarterly data over the period 1971 to 1982. Using a traditional specification of income, volatility and AP, he accepts the hypothesis of constant returns to scale and a unit long-run elasticity for the AP. He additionally includes a variable representing the remittance of Sudanese ex-patriots, given the importance of labour exporting. These swell the domestic banking system. Ford and Huang (1994) model Chinese reserves over the period 1952-89 using annual data. They find evidence that income, volatility and AP are all positively related to reserves. They also report evidence of economies of scale but note that their sample is quite small with only 38 degrees of freedom.

In addition to a scale variable (Y), a measure of volatility (VOL) and the Average Propensity to Import (AP), other researchers have included an opportunity cost measure of reserve holdings. Landell-Mills (1989) reports empirical work on the determinants of reserves for countries that borrow on international capital markets. For these countries, reserve holdings are significantly effected by the cost of their asset holdings. This was especially true for countries with difficulties servicing their debt. When the range of spreads expands so that less creditworthy countries face higher external borrowing costs these counties economise on their stocks of reserve. This refers predominantly to developing countries. Landell-Mills (1989) fails to find evidence of the importance of this variable for developed countries. This should be unsurprising given they have less difficulty servicing their debt.

Previous demand studies were primarily conducted within the context of a fixed exchange rate regime where macroeconomic adjustment to external payments disequilibria was achieved through changes in reserves. However, given the move away from a widespread system of flexible exchange rates, the question is whether reserves are held for the same reasons as before. Indeed, as mentioned above, one of the primary reasons why reserves continue to be held is to allow the monetary authority to engage in foreign exchange intervention in the post Bretton

Woods dirty float. This concurs with Frenkel's (1978) suggestion that the traditional approach is misspecified because it is framed within the context of a fixed exchange rate regime. We therefore attempt to update previous studies, by taking account of these characteristics of the recent float.

In re-appraising reserves, one approach has been to utilise the literature on the optimum degree of foreign exchange intervention (see Boyer, 1978; Chan, 1982; Kimbrough, 1983; Frenkel, 1980). It is consequently argued that reserve demand is related to the various stochastic shocks that impact upon the economy. Frenkel (1980) suggests including a variable which represents monetary and income shocks in a reserve demand equation. MacDonald (1987) implements this approach using the residuals from ARIMA models of money and income. He pools the data for 22 developed countries, and using the method of Kmenta (1971) finds evidence that these two stochastic shocks are significant.

We model reserves by paying account of the extent to which countries have engaged in foreign exchange intervention since the breakdown of the Bretton Woods agreement. In Chapter 2 we argue that this can be done by incorporating some measure of the exchange rate and therefore indirectly take account of foreign exchange intervention. The measure of the exchange rate that we adopt is the real effective rate. This is weighted to take account of the relative importance of a countries various bilateral rates, on the basis of external trade. It is equivalent to the nominal rates deflated for relative prices, since we purport that it is the real implications of exchange rate changes that the domestic country will be concerned with. A domestic monetary authority will "lean against the wind" of an appreciation of its currency (due to the deleterious effect on exporting sectors of the economy) by foreign exchange intervention and building up international reserves. The opposite will be the case with a depreciation of the real effective rate; with the monetary authority spending reserves in a effort to prevent, for example, the inflationary implications of a fall in the value of the currency. A specification of the demand for reserves which includes the real exchange rate will additionally allow us to examine the short-run interaction between our variables.

Dynamic issues have previously been considered by adopting a Partial Adjustment Model (for example Suss, 1980; Edwards, 1983 and 1984) or an Error Correction Model (ECM) used by Elbadawi (1990) and Ford and Huang (1994). Prior to the use of the ECM, Elbadawi (1990) and Ford and Huang (1994) adopt the two-step approach of Engle and Granger (1987) to examine whether reserves and their determinants are cointegrated. Elbadawi (1990) obtained a large adjustment coefficient of -0.57 using quarterly data on Sudan. Ford and Huang report a speed of adjustment of -0.81 and -0.52 to long run reserves, depending upon which monetary aggregate they include (e.g. M0 or M3). The fast speed of adjustment suggests that the authorities are quick to respond to deviations from desired reserves in the previous period and also indicates the relative importance of monetary disequilibrium based on different aggregates. However, Gonzalo (1994) argues that the Engle and Granger (1987) cointegration approach that other reserve studies adopt yields biased results. This method has been supplanted by, in particular, Johansen's (1988) Full Information Maximum Likelihood (see Gonzalo, 1994, and Hargreaves, 1994). Paying heed to recent developments in the time series econometrics literature, we attempt to incorporate a dynamic specification using two methods: a Linear Quadratic Adjustment Cost Model with a robust long-run estimator and a non-stationary VAR model estimated by Johansen technology.

The Linear Quadratic Adjustment Cost (LQAC) Model is a single equation approach based on a solvable theoretical model. It contrasts with the rather *ad hoc* partial adjustment or error correction approach, given that the latter do not explicitly provide a theoretical underpinning. An economic agent (or monetary authority) in the LQAC model minimises a multi-period quadratic cost function. There is a long-run target for reserves and agents choose an actual level to minimise the costs of being away from equilibrium and the costs of adjustment to equilibrium. Invoking rational expectations this model solves to a tightly specified dynamic equation. The second approach that we utilise is the non-stationary Vector AutoRegression (VAR) estimated by FIML. This has become a popular method of examining long-run relationships and adjustment to long-run equilibrium. Within this context it is possible to examine the short run interaction of reserves and our forcing variables. Although this approach does not have the same theoretical underpinning as the LQAC, it does have the benefit of a richer short run model and provides the basis for impulse response analysis in Chapter Three.

The basic dynamic models used in the literature have been usefully amended to incorporate other possible short run determinants of reserves. Frenkel (1978) suggests that a short run equation for reserves should incorporate domestic monetary disequilibrium. In particular, these monetary approaches predict that an excess supply of money will induce changes in reserves through balance of payments disequilibrium. Subsequently, Frenkel's idea was implemented by Frenkel (1983), Edwards (1983 and 1984), Elbadawi (1990) and Ford and Huang (1994). This represents a unification of the Monetary Approach to the Balance of Payments (MABP) and the literature on the demand for reserves. For example Frenkel (1983), Edwards (1984), Eldadawi (1990) and Ford and Huang (1994) all suggest that monetary disequilibrium amongst private agents has a significant effect on changes in reserves, at the 10% level or less. However the two papers with the most advanced econometric methods (e.g. Ford and Huang, 1994; Elbadawi, 1990) both use the Engle and Granger (1987) approach, which have been criticised as we noted above.

In Chapter Three, we examine these monetary issues within the context of a non-stationary VAR, given the broader short-run modeling possibilities within this framework. Firstly, this involves establishing a long-run equation for money demand. The existence of such a relationship had been questioned extensively in the academic literature since the 1970s (see Laidler, 1992). Recent papers have suggested these problems may be due to sample dependency (see Goodhart 1994; Muscatelli and Hurn

1996). When the volatile 1970s represents a smaller proportion of the overall sample period we may be able to produce evidence of a long-run relation based on developments in the time series literature. Establishing evidence of a money demand function allows us to consider the effect of money disequilibrium on a short-run equation for reserves. This is an indirect test of the applicability of the MABP to UK data. Setting up a VAR which includes reserves, its determinants and other monetary variables allows us to consider a number of additional issues. These include whether reserves have an impact upon the domestic monetary sector and hence whether reserve creation is sterilised. We can also examine the interaction of monetary and real variables within a small monetary model, which takes account of the external sector and an explicit role for policy makers' preferences with respect to reserves. Juselius (1996) suggests such an approach.

In Chapter Four we broaden our sample to incorporate a number of other European countries in our study of reserves. This allows us to consider whether the long-run specification for reserves we attempt to establish in earlier Chapters can be applied to other industrialised countries. We compare and contrast these estimated results with those using a traditional specification, which includes Y, VOL and AP. We utilise robust long-run estimators, tests for parameter stability and cointegration to assess the relative performance of our demand equations.

Furthermore, widening our sample allows us to utilise developments in non-stationary panel data. These are a recent innovation and have been applied *inter alia* to the literature on exchange rates and growth and convergence (see Canzoneri, Cumby and Diba, 1996; Chinn and Johnston, 1996; MacDonald, 1996; Obstfeld and Taylor, 1996; Taylor, 1996). These empirical methods are used to provide evidence of sensible long-run relationships, panel cointegration and to examine short-run dynamics or mean reversion. When examining reserves a number of researchers have pooled the times series and cross sections components of their data (including Frenkel, 1983 and Landell-Mills, 1989) and used fixed effects panel models. No studies have, to the best of our knowledge, utilised non-stationary panel data when examining the demand for reserves. Non-stationary panel methods are a means of widening the span of data available in any empirical study. Hence, by increasing degrees of freedom we can be confident our results are more robust to small sample bias when testing, in particular, for cointegration. Additionally, we also use the non-stationary panel estimator of Pesaran, Shin and Smith (1999). This allows us to compare and contrast our long-run estimated coefficients across reserve specifications and the degree of mean reversion to the long-run.

Staying within the broad area of international reserves, in Chapter Five we consider the relationship between high frequency foreign exchange intervention and the second moments of the exchange rate. That is to say, we examine whether central bank foreign exchange operations are statistically related to short run exchange rate volatility. The theoretical literature in this area emphasises that intervention can have an effect on the level of the exchange rate when assets denominated in different currencies are imperfect substitutes (see MacDonald, 1988). This is known as the Portfolio Balance Channel. If assets are not imperfect substitutes it is argued that intervention may still have an impact through the Signalling Channel. This is when the central bank conveys inside information about the course of future fundamental determinants of the exchange rate by its foreign exchange operations (see Mussa, 1981).

Much of the recent empirical interest in reserves is motivated by highly influential papers by Dominguez (1990), Dominguez and Frankel (1993a) and Catte et al. (1994), which suggest that daily foreign exchange intervention has a significant impact upon the level of the exchange rate. In particular, Dominguez (1990) and Catte et al. (1994) found statistical evidence that coordinated intervention by G-3 central banks influenced the exchange rate using a mid-1980s sample period. A number of recent studies have gone on to examine the effect of intervention on daily exchange rate volatility. These studies are summarised in Table 1.2.

Although the literature is not unanimous for the period we study between 1985 and 1987, there is reasonable evidence that intervention is significantly related to exchange rate volatility. In particular, Dominguez (1998) and Bonser-Neal and Tanner (1997) argue that daily US Federal Reserve intervention reduced nominal exchange rate volatility in a mid 1980s sample period. This suggests that foreign exchange intervention, which is predicated on the existence of reserve holdings, may be of some consequence for policy makers. We contend that the literature in this area

Table 1.2: Research on Daily Foreign Exchange Intervention and Volatility

Author	*Intervention Data Used*	*Methods Used*	*Sample Period*	*Main Results*
Baillie and Humpage (1992)	Actual Fed, Buba and BoJ Intervention Data.	GARCH	Feb1987 - Feb 1990	Volatility affected but not in a consistent manner.
Baillie and Osterberg (1997)	Actual Fed data separated into buying and selling.	GARCH	Apr 1985 - Dec 1986	No effect on volatility but some on mean.
Bonser-Neal and Tanner (1996)	Actual and reported Fed, Buba. Reported BoJ.	Implied Volatilities	Jan 1985 - Feb 1987	Reported Fed reduces DM-$ volatility.
Connolly and Taylor (1994)	Actual BoJ.	Standard Deviations	Jan 1977 - Dec1979	Intervention significantly increases volatility.
Dominguez (1998)	Fed and Buba reported and secret. Reported BoJ.	GARCH and Implied Volatilities	Jan 1985 - Feb 1987	Fed and Buba intervention reduces DM-$ volatility.
Hung (1997)	Actual Fed intervention.	Standard Deviations	Apr 1985 - Dec 1986	Fed reduces the volatility of the DM-$ rate.
Mundaca (1990)	Norwegian data.	GARCH	Jan 1977- Dec 1979	Volatility increased.

Note: Where possible we concentrate on the sample period 1985-1987, since this is our data sample in Chapter Five. Hung (1997), Bonser-Neal and Tanner (1997) and Dominguez (1998) examine additional sample periods not included.

may benefit from further research. We add to pervious volatility studies by using Generalised AutoRegressive Conditional Heteroskedasticity methods and official daily data to examine the relationship between coordinated intervention in Chapter Five. This distinction between unilateral and coordinated intervention has been emphasised when examining the effect of intervention on the level of the exchange rate. This is especially true given that previous studies, which have jointly examined simultaneous US and German intervention may suffer from a collinearity problem. We hope to circumvent this problem by defining our variables more carefully. We also examine Pound Sterling bilateral rates, which have not to the best of our knowledge been considered in the literature. This additionally allows us to examine the policy externalities of foreign exchange intervention. Chapter Six concludes this thesis on re-appraising reserves and makes further suggestions for practicable and profitable research in this area.

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Chapter Two

The Demand for International Reserves: With an Application to Foreign Exchange Intervention

2.1 Introduction

The effectiveness of foreign exchange intervention is one of the most contentious issues in international economics. Some researchers dismiss it as little more than "smoke and mirrors", with little intended effect (Obstfeld and Rogoff, 1995). Others believe that it is useful as an independent policy tool, and have found evidence consistent with this view (see for example Dominguez and Frankel, 1993). Intervention has been used extensively by a number of central banks since the collapse of the Bretton Woods system of fixed exchange rates (witness the Plaza Agreement, the Louvre Accord and the 1995 reversal of the US Dollar decline after concerted intervention by the G-7). Hence this policy instrument continues to be an important area of investigation. In this Chapter we examine issues related to intervention, within the framework of the literature on the demand for international reserves. In the past these two fields of study have been considered in isolation but, we believe, both may benefit from a more integrated approach.

Traditionally, the demand for international reserves has been studied within the context of the Balance of Payments (BoP). As a result, foreign exchange reserves are viewed as a buffer stock, held as a precaution against fluctuations in a country's international accounts. We contend that this approach requires a reassessment. This is especially in light of the development of inter-temporal theories of the BoP,[2] and the belief on the part of policy makers that fluctuations in a country's international accounts may merely reflect decisions on the part of private agents - witness the 'Lawson Doctrine' at the end of the 1980s.[3] Combining these points with the substantial growth in international capital movements, where current account deficits are temporary phenomena reflecting private sector decision that can be financed by means other than out of reserve holdings, central banks may be holding foreign assets for reasons other than as buffer stock.

Reserves have continued to be held by countries despite this reduction in concern about Balance of Payments problems. Friedman (1953) suggested that the abandonment of the fixed system of exchange rates between the currencies of the industrialised world would rule out the need for reserves. However, there is evidence to suggest that demand for reserves did not fall after the collapse of Bretton-Woods (Williamson, 1976; Gandolfo, 1995) but actually increased (Mussa et al., 1994).

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[2] For a review of inter-temporal issues related to the Balance of Payments, see Cordon (1994).

[3] See Lawson (1992).

In his analysis, Friedman (1953) referred specifically to the demand for reserves after the abandonment of a fixed regime. This meant the move to a pure float and not the dirty float that the major currencies of the industrialised world have participated in recently. Central banks have intervened extensively in the foreign exchange market since the end of Bretton Woods. They have bought and sold domestic and foreign assets, in an attempt to reduce excessive fluctuations and remove fundamental misalignment in the exchange rate. This intervention is predicated on the existence of international reserves. If we are to consider the determinants of reserves within the recent dirty float period we must pay reference to central banks involvement in the foreign exchange market.

We start by summarising some earlier empirical studies of the demand for reserves. The first such study was Triffin (1947), who considered the reserve-import ratio as a measure of reserve adequacy. More recent papers take their precedence from Kenen and Yudin (1965), who considered the level of reserves as a function of various measures of reserve instability. The latter idea has been widely imitated by those taking a static regression approach.[4] Consequently reserves (R) have been explained in terms of the following variables: a scale variable (i.e. income

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[4] Some other authors have attempted to incorporate an opportunity cost measure of the demand for reserves in their static regression (see Landell-Mills, 1989). Although this has sometimes been successful, in the sense that the included measure is significant, this seems largely dependent upon the countries debt-servicing record. Countries with less creditworthiness have higher borrowing costs and a greater tendency to economise on reserve holdings. Creditworthiness is not an important issue for the UK, hence this variable is not part of our analysis.

(Y) or imports); a measure of uncertainty in the Balance of Payments (VOL); and the degree of external exposure of the economy (AP).

$$log(R)_t = constant + \beta_1 log(Y)_t + \beta_2 log(VOL)_t + \beta_3 log(AP)_t \qquad (2.1)$$

See Chapter One and Bahmani-Oskooee (1985) for a more extensive review of static regression studies.

Frenkel (1978) makes the suggestion that this static regression approach may be misspecified, since it does not take account of the use of reserves in a dirty 0float. We can also be doubtful of traditional studies, given their extensive use of OLS estimation, which in the presence of non-stationary variables may produce spurious regression results (Granger and Newbold, 1974). Although Ford and Huang (1994) take account of non-stationarity using the Engle and Granger (1987) approach, Banerjee et al. (p214-230, 1993) show using Monte Carlo methods that their can be substantial small sample biases using this estimator. These criticisms suggest that we should include different variables in the long-run relationship and take account of the presence of non-stationarity by using an alternative estimation strategy.

The variables that we use in our long-run relationship for international reserves are the real effective exchange rate (Q), prices (P) and income (Y). The incorporation of the exchange rate more accurately reflects the actual behaviour of central banks in a managed float. It also allows us to examine whether The Bank of England has attempted to resist malevolent fluctuation in the exchange rate by "leaning against the wind". This can be defined as intervention operations that attempt to move an exchange rate in the opposite direction from its current trend. There may be some ambiguity in the definition, since it in turn depends on the current trend. Intervention in support of a currency may look as if it is leaning against last week's trend but leaning with last month's trend (when exchange rates are volatile). "Leaning with the wind" is defined as operations that are motivated by a central bank's desire to support the current exchange rate trend. This is exemplified by the 1985 Plaza Agreement between the G-5, which supported the Dollar in an already downward path.

A real - rather than a nominal - exchange rate is used in our study as this reflects domestic export competitiveness. An appreciation of the nominal rate with relative price deflation (implying a constant real exchange rate) will have less of a consequence for the domestic economy than if domestic prices were unchanged and the nominal exchange rate appreciated. Our real exchange rate measure (Q) takes account of this important distinction and its implication for real sectors of the economy. Moreover we take account of the level of trade the UK has with each particular country or bilateral UK rate, by using a trade weighted or real effective exchange rate. This variable has not, to the best of our knowledge, been used before in descriptive studies of international reserves. We believe that because it proxies exchange rate policy it is an important determinant of recent reserve holding and should be included in a static reserves regression.

Prices and income are included in the long-run relationship to assess the interaction between reserves and the government's other macroeconomic target variables. In particular, the inclusion of prices allows us to examine whether there is a long-run relationship between it and reserves - or whether the domestic monetary implications of intervention have been sterilised. When central banks intervene they buy foreign assets for domestic assets and, if they are not sterilised, these interventions will result in an increase in the domestic money base. For example, if the BoE intervenes buying Dollars, its portfolio of foreign assets increases and domestic deposits decrease. At the same time, the sterling deposits of commercial deposits at the BoE increase. As a result, the UK Money Base (commercial bank deposits at the BoE plus currency in circulation) is increased. The BoE can sterilise this increase by selling the appropriate number of Sterling-denominated assets through Open Market Operations. If interventions were not sterilised in the long-run, then we should expect prices to be related in some way with reserves through a 'monetarist channel'. The inclusion of income could be considered as a scale variable - much the same way as in traditional demand for reserves or money studies. Our alternative relationship is therefore of the form:

$$log\,(R)_t = constant + \beta_1 log(Q)_t + \beta_2 log(Y)_t + \beta_3 log(P)_t \qquad (2.2)$$

Dynamic issues are also considered in our paper. This has been attempted in papers examining reserves by Frenkel (1983) and Edwards (1983) using a Partial Adjustment framework, and Elbadawi (1990) and Huang (1995) using an Error Correction Model. These earlier approaches may suffer from a number of problems. These include: the rather *ad hoc* nature of the Partial Adjustment and Error Correction formulation; the failure, within the context of single equation estimators, to deal with weak exogeneity; and neglecting to consider whether there is more than one cointegrating vector between the variables.

To avoid these problems, we use a number of different approaches in estimating a dynamic demand relationship. First, we employ a Linear Quadratic Adjustment Cost (LQAC) Model using estimation techniques developed by Kennan (1979), Dolado, Banerjee and Galbraith (1991) and Gregory, Pagan and Smith (1993). Taking account of Pesaran's (p178, 1997) criticism of recent moves in applied work, this gives our single equation analysis a robust theoretical basis, but retains a simple solution. Second, we implement the Johansen (1988) multiple equation technology with reference to the modelling strategy of Hendry and Mizon (1993). This takes account of the absence of weak exogeneity of our explanatory variables with respect to the long-run relationships, by modelling reserves in a system of equations. We can also use this framework to examine whether weak exogeneity fails within the single equation context. Additionally, the Johansen method: addresses the possibility of multiple cointegrating vectors; facilitates the imposition of restrictions on the long-run relationships; and provides a richer framework for analyses of the short-run dynamics of our system.

The plan of Chapter Two is as follows. Section 2.2 sets out the LQAC model and considers some of the different alternatives to its estimation. We introduce the Johansen (1988) method by referring to the approach of Hendry and Mizon (1993). Section 2.3 includes an initial examination of the data. Section 2.4 presents the results from the single equation estimator of the long- and short-run demand for reserves. Section 2.5 presents the results using the Johansen technology. Section 2.6 contains a summary of the main results of Chapter 2 and concludes.

2.2 Estimation of the Demand for International Reserves

2.2.1 The Linear Quadratic Adjustment Cost Model

The Linear Quadratic Adjustment Cost (LQAC) Model is a tractable model in which an economic agent minimises a multi-period quadratic cost function (equation *(2.3)*).[5] The agent (in this case the central bank) is assumed to track a long run target for reserves (y_s^*), as given by a static equilibrium theory. The agent then chooses the actual level (y_s) so as to minimise the weighted sum of the costs of being away from equilibrium ($y_s - y_s^*$) and the costs of adjustment towards equilibrium ($y_s - y_{s-1}$).

The problem the central bank faces is therefore:

$$\min_{\{y_s\}} E_t \sum_{s=t}^{\infty} \beta^{s-t} \left[\delta (y_s - y_s^*)^2 + (y_s - y_{s-1})^2 \right] \qquad (2.3)$$

for $s \geq t$, where the expectation is taken with respect to information available to the agent at time t (Φ_t), $\beta \in (0,1)$ is a discount factor, and $\delta > 0$ is a weighting factor. The static equilibrium relationship is $y_t^* = x_t^T \theta + e_t$, where e_t is a mean 0, independently and identically distributed error with variance σ^2_e and x_t is a k*1 vector of forcing variables. We assume that e_t is in Φ_t but unknown to the investigating econometrician whose information set is only $\Im_t \subset \Phi_t$. Notice that if we assume that e_t is observable by the econometrician, then y_t is a deterministic function of the information set. We also assume serially uncorrelated errors (see Sargent 1978).

The first order condition for the optimisation problem is the Euler equation

$$\Delta y_t = \beta E_t \Delta y_{t+1} + c\left(y_t - y_t^*\right) \tag{2.4}$$

where $c = -\delta$ and E_t is the expectation taken w.r.t. Φ_t. Both roots of the characteristic equation of this second order difference equation are positive and lie either side of unity. Denote the stable roots of the quadratic $\beta z^2 - (1+\beta+\delta)z + 1 = 0$ by λ, leading to the forward solution of equation *(2.4)*:

$$y_t = \lambda y_{t-1} + (1-\lambda)(1-\beta\lambda)E_t \sum_{s=t}^{\infty} (\beta\lambda)^{s-t} y_s^* \tag{2.5}$$

In essence we have produced a tightly specified dynamic model directly from economic theory and which we will consequently apply to our data. The model has the characteristic that due to the costs associated with the central bank borrowing abroad and the unwanted impact that buying reserves on the foreign exchange market may have on the exchange rate, reserves will adjust gradually towards their optimal level. Where $\delta = -c$ is small, the movement of reserves towards the long-run

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[5] This theoretical approach to an economic problem was developed by Sargent (1978).

target is slow, and adjustment costs are high. Since the model has a forward looking aspect, the level of reserves will in part depend upon expected future levels of the forcing variables. It also has the 'plausible' property that due to the discount factor the weight on future expected values of the "forcing values", $E_t y_s^*$, declines with time. Additionally, the model proposes the inclusion of a lagged dependent variable. This is useful since time series invariably have a strong autoregressive component. In estimating the model we will assume rational expectations, i.e. agents do not make systematic forecast errors. This assumption is made so we can replace the unobservable expectations in the model by observable values.

Equations *(2.4)* and *(2.5)* form the basis of estimation techniques advanced to estimate LQAC models. We concentrate on using the Dolado et al. (1991) methodology for this particular model, since earlier approaches - in particular Kennan (1979) - failed to take account of stationarity issues. We implement Dolado using OLS, Phillips and Hansen's (1990) Fully Modified (FM-OLS) and Phillips and Loretan's (1991) Non-Linear Least Sqaures (NLLS) estimators.

2.2.2 Dolado et al. (1991) Single Equation Estimation

The Dolado et al. (1991) estimation strategy enhances the Kennan (1979) estimation procedure for Euler equations in LQAC models, to specifically deal with variables integrated of order one i.e. I(1). Dolado's strategy is rooted in the two-step approach of Engle and Granger (1987).

Firstly we check, using the normal testing procedures, that the orders of integration of y_t and x_t are consistent with the possibility of cointegration. If the equation derived from static equilibrium theory ($y_t^* = x_t'\theta + e_t$) is unbalanced, we rule out the possibility of cointegration, and the data-generating process of (y_t , x_t) can not have the characteristic of the LQAC model. The absence of a cointegrating long-run relationship, implies that the central banks optimal strategy never involves a choice of y_t such that the gap between y_t and y_t^* is asymptotically eliminated. The central bank does not incur the additional adjustment costs necessary to catch up with the target. We can not therefore characterise the behaviour of the central bank by the LQAC model, when cointegration is ruled out. If the equation is not unbalanced, we can proceed to test the null of no cointegration between y_t and x_t. If two time series y_t and x_t are both I(d), then in general any linear combination of the two time series will also be I(d). However there may exist a vector θ, such that the disturbance term from the regression $e_t = y_t - \theta x_t$ is of a lower order of integration I(d-b) where b>0, then y_t and x_t are cointegrated of order d minus b (e.g. CI(d,b) in Engle-Granger notation). For example, if y_t and x_t ~I(1) with e_t ~I(0): then y_t and x_t ~CI(1,1). We test for the existence of cointegration using the Engle and Granger (1987) Augmented Dickey Fuller test, the Johansen (1988) Trace statistic and the Hansen (1992) Lc test.

Secondly, we re-parameterise *(2.4)*, by replacing $E_t\Delta y_{t+1}$ with $(\Delta y_{t+1} + \eta_{t+1})$, where $E_t(\eta_{t+1}) = 0$. That is to say, we assume rational expectations. Therefore, equation *(2.4)* becomes:

$$\Delta y_t = \beta\Delta y_{t+1} + c(y_t - x_t'\hat{\theta}) + v_t \qquad (2.6)$$

Where $v_t = \beta\eta_{t+1} - ce_t$ and $\hat{\theta}$ is estimated by a previous static regression and exhibits Stock's (1987) super-consistency characteristic,[6] if the null of no cointegration is rejected in the long-run relationship. Equation *(2.6)* may be estimated consistently by Instrumental Variables, with instruments taken from the information set Φ_{t-1} (e.g. $\hat{e}_{t-1}\left(= y_{t-1} - x_{t-1}'\hat{\theta}\right)$ and Δy_{t-1}, instead of $\hat{e}_t$ and Δy_{t+1}).[7] This will consequently give us an estimated adjustment coefficient and discount factor. (As a useful means of comparison to the LQAC model we also use an ECM to estimate the adjustment coefficient.)[8]

Gregory et al. (1993) suggests that there are some complications to simply estimating equation *(2.6)* using IV. These include: an identification problem; potential endogeneity; autocorrelation and invalid inference on the basis of the estimated θ. Identification is a problem since there are five estimated parameters (θ_1, θ_2, θ_3, β and c) and only four variables. This first problem is resolved by the imposition of β. For example, estimate an equation of the form:

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[6] The estimate approaches its population value at a rate faster than in normal asymptotics (i.e. $O(T^{-1})$ instead of $O(T^{-1/2})$).

[7] See Dolado et al. (p927, 1991).

[8] Nickell (1985) shows that the Euler equation derived from the quadratic objective function could be reformulated as an ECM, suggesting the two approaches are isomorphic. The use of the Euler equation allows estimation of the parameters associated with the objective function, giving us an explicit theoretical specification.

$$\psi_t = c\left(y_t - \underline{x}_t'\underline{\theta}\right) + v_t \qquad (2.7)$$

with $\psi_t = \Delta y_t - \beta \Delta y_{t+1}$. Again we use a two step method in this procedure. It should be noted that although OLS estimation of the long-run parameters produces super-consistent parameter estimates under cointegration, these may be biased in small samples, especially if the model suffers from autocorrelation and endogeneity. Consequently the associated t-statistics will not be asymptotically normally distributed, and can not be used as the basis for normal statistical inference.

However, if we employ Phillips and Hansen's Fully Modified non-parametric correction to OLS, we can estimated our long-run parameters taking account of potential endogeneity and autocorrelation, by utilising information from the long-run covariance matrix.[9] The Phillips and Hansen (1990) estimator assumes y_t and x_t is generated by

$$\begin{aligned} y_t &= x_t'\theta + \xi_t \\ \Delta x_t &= \varepsilon_t \end{aligned} \qquad (2.8)$$

Two corrections are made to $\hat{\theta}$ by FMOLS. First, there is the correction for bias in $T\left(\hat{\theta} - \theta\right)$ because of the endogeneity of Δx_t. This is achieved by replacing y_t by $y_t^+ = y_t - \hat{\Omega}_{12}\hat{\Omega}_{22}^{-1}\Delta x_t$, where $\hat{\Omega}$ is an estimator of Ω, the long-run covariance matrix of $\varsigma_t' = \left[\xi_t, \Delta x_t\right]$. We then regressing y_t^+ against

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[9] See Banerjee et al. (p240, 1993) as to why FM-OLS is preferable to dynamic single equation estimation of the long-run relationship. For a contrary viewpoint see Pesaran and Shin (1995).

x_t, giving $\theta^* = \left(\sum x_t x_t'\right)^{-1} \sum x_t y_t^+$. The second correction for autocorrelation bias involves modifying θ* to $\hat{\theta}_{PH} = \theta^* - \left(\sum x_t x_t'\right)^{-1} Tg$ where

$$g = \hat{\Lambda}\begin{bmatrix} 1 \\ -\hat{\Omega}_{22}^{-1}\hat{\Omega}_{21} \end{bmatrix} \qquad (2.9)$$

and $\hat{\Lambda}$ is a consistent estimator of $\Lambda = \sum E\left(\Delta x_0, \varsigma_k'\right)$. Consequently, our t-statistics are normally distributed asymptotically and can form the basis of classical inference on the long-run parameters.

2.2.3 Multiple Equation Estimation

As mentioned above, there may be a number of problems in using single equation procedures in estimating long-run parameters. Additionally these problems will spill over into the corresponding dynamic equations. In particular the issue of weak exogeneity must be considered. According to Banerjee et al. (p163, 1993), single equation dynamic models are not optimal if weak exogeneity fails to hold. Weak exogeneity requires that there be no loss of information about the parameters of interest, say, in our long run relationship in reducing our analysis from the joint distribution to a conditional model. The absence of this property is exhibited in a bi-variate equation system by a cointegrating relationship ($\beta' x_t$) entering both the ith and jth equations. Consequently, x_{jt} can not be considered as weakly exogenous for the parameters of the ith equation, since the parameters of the two equations share the common component $\beta' x_t$ (or in other words, can not be variation free). Failure to take account of this cross equation linkage is detrimental to the validity of inference in finite samples (see Phillips and Loretan, 1991). Modelling within a system of equations framework will take account of these cross equation linkages.

A second problem with single equation models is that they fail to provide information on whether there is more that one cointegrating vector. However, if we model our time-series within a Vector AutoRegressive Model (VAR), we can test for more than one cointegrating vector. As a means of taking account of these particular issues, and also as a useful comparison to estimation of our demand for reserves using the method derived from an Euler equation, we shall adopt the Johansen (1988) multivariate cointegration framework. In doing so we pay particular attention to the Hendry and Mizon (1993) modelling strategy.

In this context, our approach begins with a general VAR in levels. Where x_t is a $(n*1)$ vector of variables[10] $x_t = \{R, Q, P, Y\}$ we assume that it has the vector autoregressive representation of the form

$$x_t = \sum_{i=1}^{p} \Pi_i x_{t-i} + \psi D_t + \varepsilon_t \qquad (2.10)$$

where ε_t is a $(n*1)$ vector of white noise disturbances, which is assumed to be zero mean, homoscedastic and serially uncorrelated, with covariance matrix Ξ. D_t contains a constant and/or trend.

We can reformulate the VAR as a Vector Error Correction Model:

$$\Delta x_t = \sum_{i=1}^{p-1} \Gamma_i \Delta x_{t-i} + \Pi x_{t-1} + \psi D_t + v_t \qquad (2.11)$$

where v_t has the standard properties. The order of the VECM is finite and the parameters Γ, ψ and Σ (the covariance matrix of v_t) are assumed constant. To calculate the number of cointegrating relations in our model, we have to find the rank of the long-run responses (Π). In other words we calculate the number of linearly independent columns of Π. Three cases can be distinguished when ascertaining the rank (r) of Π: full rank, zero rank and reduced rank.

(1) Full Rank : r = n, consequently all n variables in x_t are I(0), and hence they can only be non-stationary via D_t.

(2) Zero Rank : r = 0, therefore all n variables, in x_t are I(1), so that if there are not deterministic non-stationarities Δx_t is stationary.

(3) Reduced Rank : 0 < r < n, in which case there are (n-r) linear combinations of x_t which act as common stochastic trends, and r cointegrated linear combinations of x_t

In the third case, Π can be factored into $\alpha\beta'$ where both α and β are (n*r) matrices, of rank r, and the β contains the coefficients of the r cointegrating vectors. The cointegrating relations $\beta' x_t$ are I(0) and the loading matrix α gives the weight associated with each cointegrating vector, in all n equations of the VAR.

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[10] All variables are log transformed. See Section 2.3 for a discussion of the variables in this study.

We test for reduced rank using the Trace (*Tr(r)*) statistic. This is used to test the null hypothesis that the number of distinct cointegrating vectors is less than or equal to r, against a general alternative.

$$Tr(r) = -T\sum_{i=r+1}^{n}\ln(i-\hat{\lambda}_i) \qquad r=0,1,\ldots,n-1 \qquad (2.12)$$

$\hat{\lambda}_i$ are the characteristic roots or eigenvalues obtained from Π and T is the number of observations. The larger the eigenvalues, the larger the *Tr(r)* statistic and hence the likelihood of rejecting the null hypothesis of $r \leq i$.

Next, in our modelling strategy, we impose restrictions on the cointegrating vectors or long-run relationships. This is an attempted to produce "meaningful economic relationships", which underlie the long-run model. We also consider whether they are unique. The imposition of restrictions on the various β_i will in turn alter the estimated short-run dynamics in our VECM. This new model is considered as the Parsimonious Vector AutoRegression (PVAR).

At this point we can conduct tests on the weak exogeneity of the four variables with respect to the parameters of our cointegrating relationships. This involves examining whether the adjustment coefficients of the cointegrating relationships are significant in each of our equations. If they are not significant then we can condition on their contemporaneous differences or potentially move to a single equation or conditional model.

The following step is to model our PVAR to produce a Structural Econometric Model (SEM). This will allow interpretation of the short run dynamics. Restrictions are imposed on the individual parameters of the short-run vectors Γ_i, where the corresponding t-ratios from the PVAR are less than two. Finally, we consider whether our SEM parsimoniously encompasses the PVAR. This can be done using a test on the over-identifying restrictions, which has a χ^2 distribution.[11]

2.3 Preliminary Data Analysis

2.3.1 Data Sources

The source of all data in this study of the UK demand for Reserves was IMF *International Financial Statistics*. The sample period starts in 1975 Quarter 1 and ends in 1995 Quarter 3. Our four variables are international reserves (R), the real effective exchange rate (as a measure of competitiveness) (Q), the GDP deflator (P) and real GDP (Y). All variables are seasonally adjusted and transformed into natural logarithms. (See Appendix 2.A for further details.)

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[11] Cuthbertson and Taylor (1990) and Engsted and Haldrup (1994) estimate a LQAC model using the Johansen approach. Although these are distinct methodological approaches, Engsted and Haldrup (1994) argue that Johansen can be transformed into a comparable single equation model where Δx_t is weakly exogenous w.r.t. the long-run parameters of interest, and that Δy_t does not Granger cause the Δx_t variable. Due to the vector (y_t, x_t) including a MA(1) error term in equation *(2.6)*, the Johansen VAR method may not be theoretically appropriate (see Gregory et al. 1993, p231). This may not be such a problem if the VAR representation "soaks up" a sufficient amount of the complex error structure. Unfortunately the Granger Causality condition is not satisfied by our data (see Section 2.4), hence we use the Johansen method merely for comparative reasons not because the two approaches are isomorphic.

2.3.2 Unit Root Tests

Consistent with applied work in general, and Dolado et al. (1991) in particular, we pre-test our four variables for I(1)-ness. We utilise the tests prescribed by Fuller (1976) and Dickey and Fuller (1979) under the null hypothesis of unit root. The results, as included in Appendix 2.B, give evidence against all variables being I(0). The variables included in our study do not provide us with clear evidence of our static regression equation being unbalanced. This consequently provides us with preliminary evidence not inconsistent with the possibility of cointegration.

2.3.3 Cointegration Tests

The three tests that we use in examining cointegration in our long-run relationship are: the Engle and Granger (1987) ADF Test, the Johansen (1988) Trace Test, and the Hansen (1992) extension of FM-OLS. The augmented Dickey-Fuller (ADF) t-ratio test (see Said and Dickey 1984), recommended by Engle and Granger, is based on the residuals of the cointegrating regression. The test has a null hypothesis of no cointegration, with critical values from MacKinnon (1991). This test has notoriously low power. From above, the Johansen Trace Test is based on the rank of Π. Critical values used here are from Osterwald-Lenum (1992). Finally we use the Hansen (1992) tests for parameter stability, to consider whether there is a cointegrating relationship between our variables.

Table 2.2: Cointegration Tests on UK Reserve Demand 1975(1) - 1995(3)

Engle Granger		ADF(7)	τ_μ=-3.91*	95%C.V. = -3.85
	H_0 :	λ_{t+1}	Trace Test	95%
Johansen	$r = 0$	0.40	77.14*	53.12
	$r \leq 1$	0.23	35.98*	34.91
	$r \leq 2$	0.12	14.72	19.96
	$r \leq 3$	0.05	4.39	9.24
		Test	Estimated Statistic	p-value (≥0.20=0.20)
Hansen		L_c	0.21	0.20
		MeanF	5.09	0.20
		SupF	45.12*	0.01

Notes: Critical values for the ADF test are from MacKinnon (1991). Lag length of test in brackets. For the Johansen trace test, critical values are from Osterwald-Lunum (1992). The lag length of the VAR is three. * indicates significance at the 95% level. The Hansen (1992) Lc test has a null hypothesis of cointegration.

This utilises Phillips and Hansen's FM-OLS. The first two test statistics, MeanF and L_c, analyse the assumption that A_t follows a martingale process (i.e. $A_t = A_{t-1} + \varphi_t$): the null hypothesis is that the variance of the martingale is stationary. The SupF tests for a single unknown structural break at time t in A_t, where $y_t = A_t x_t + u_t$. Hansen's L_c can also test the null of cointegration against the alternative of no cointegration. This is useful because usually tests can only produce evidence against the null hypothesis of no cointegration. This is not the same as evidence in favour of cointegration. The other two tests have power against, but are not exactly geared towards, the alternative of random walk coefficients (see Hansen, 1992). The results from the Engle-Granger, Johansen and Hansen tests of cointegration are presented in Table 2.2.

The strike rate in favour of cointegration is high. The ADF test is significant at the 5% level. The Johansen results suggest the null of at most two cointegrating vectors can be accepted and that the null of no more than one cointegrating vector can be rejected. The Hansen (1992) tests L_c and MeanF provide evidence against a gradual change in the relationship over time, and therefore can be considered as evidence in favour of cointegration, especially the L_c test statistic. However the SupF test statistic suggests there has been an abrupt change in our equilibrium relationship. Nevertheless, the large majority of the results can reject the null of no cointegration and the L_c accepts the null of cointegration.

2.3.4 Weak Exogeneity Tests

A useful characteristic of the variables *R*, *Q*, *P* and *Y* would be weak exogeneity with respect to the long run parameters. This would allow us to use a single equation error-correction model for ΔR to estimate the adjustment coefficient to a long-run reserve relationship. Under the null of weak exogeneity this test is asymptotically distributed as $\chi^2(r)$, where r is the number of cointegrating vectors in the model.

The results in Table 2.3 from the Johansen approach suggests that with two cointegrating vectors none of the variables are in fact weakly exogenous. We should therefore be wary of using single equation estimators that do not take account of this potential endogeneity. With one cointegrating vector, only reserves appear to be weakly exogenous. This is worrying if the first vector represents our demand relationship and would

Table 2.3: Weak Exogeneity w.r.t. the Long Run Parameters of Interest

r	$\chi^2(r)$	*R*	*Q*	*P*	*Y*
1	3.84	1.91	7.30	7.61	7.32
2	5.99	9.79	7.43	9.68	7.54
3	7.81	12.76	13.50	9.69	10.63

Note: Critical values are obtained from the Econometrics Packages CATS in RATS.

suggest that reserves are not determined endogenously within this particular model. Therefore, we should remain within the system analysis. We return to these issues later.

2.4 The Single Equation Estimation

2.4.1 Long-Run Relationships

Three methods were used to estimate our long-run static equation (2.2): Ordinary Least Squares; Phillips and Loretan (1991) Non-Linear Least Squares (NLLS); and Phillips and Hansen's (1990) FM-OLS. Some mention of the NLLS estimator is necessary, by way of an introduction. NLLS is a parametric procedure for estimating the cointegrating vector in an equation where the variables are already known to be characterised by cointegration. The approach deals with simultaneity and autocorrelation in the residuals by including lagged and lead values of the changes in the regressors and lagged stationary deviations from the cointegrating relationship. Estimates are consequently asymptotically efficient and can be used as the basis for standard inference. This estimator is appropriate within the context of the LQAC model, as the first step in estimating the long-run relationship.

Table 2.4: Single Equation Estimation of the Long-Run Relationship

Estimator	**OLS**	**NLLS**	**FM-OLS**
Regressand	R	R	R
Regressors			
Constant	-116.80*	-94.53*	-96.57*
	(23.32)	(14.20)	(37.61)
Q	1.54*	0.95*	1.19*
	(0.38)	(0.23)	(0.61)
P	-0.54*	-0.20	-0.25
	(0.28)	(0.22)	(0.45)
Y	4.95*	4.21*	4.26*
	(0.82)	(0.49)	(1.31)

Notes: Standard errors of estimates are in parenthesis. * is significant at the 5% level.

Our three estimators OLS, NLLS and FM-OLS in Table 2.4 all give similarly sized and equivalently signed parameter estimates. The estimated sign of the coefficient for Q are consistent with the Bank of England intervening "against the wind" in the foreign exchange market. That is, buying foreign currency when sterling is strong and selling foreign currency when it is weak.[12] Using FM-OLS and NLLS produced similar results with respect to the significance of our three explanatory variables. Both of these estimators suggested prices were insignificant in the long-run relationship - since the t-ratios are considerably less than two. This is evidence that the domestic monetary implications of intervention are sterilised in the long-run. This was in contrast to the results for OLS which suggest that P is significant. However, we should be cautious when using

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[12] Traditional studies of foreign exchange intervention have followed the example of Wonnacott (1965) by capturing leaning against the wind in a regression of changes in reserves on changes in the exchange rate. They also include a constant and other explanatory variables. Our approach attempts to pick up leaning against the wind using the levels of the data, reflecting our combined examination of central bank intervention and the demand for reserves.

OLS estimates, since non-stationarity may bias estimates in small samples even if estimates are asymptotically efficient in the presence of cointegration. Our two robust methods suggest that the real effective exchange rate and income are both significant. For now we retain prices in our long-run relationship but return to these issue in Section 2.5.

2.4.2 Short-Run Dynamics

In this section we consider Dolado's IV method of estimating the Euler equation (*2.6*). OLS estimates of the long-run relationship were used in this two-step method, since these did not deviate substantially in estimated size or sign from the PL and FMOLS estimators. Due to the identification problem in that approach, we also execute the Gregory et al. (1993) correction and impose the discount factor, β. Following Gregory we let β equal 0.90, 0.95 and 0.99. Additionally in Table 2.5, we report the results from Engle and Granger's (1987) Error Correction Model.

The estimated adjustment coefficient from the Dolado equation is -0.14 with an estimated standard error of 0.06. Mean reversion is not large but it is statistically significant. This suggests a half-life of four quarters ($(1-0.14)^n=1/2$, hence n=4).[13] The estimated discount factor ($\hat{\beta}$) is equal to 0.15, implying a discount rate of around 85% per quarter. The estimated

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[13] The results from Suss (1980) are applicable to our UK study, and suggest that adjustment towards equilibrium is slow. Suss (1980) models the level of reserves on the change in the nominal exchange rate, imports and lagged reserves. He found that the exchange rate entered this long-run relationship significantly.

Table 2.5: Short-Run Adjustment Coefficients for UK Reserves

Estimator	**Dolado**	**Gregory (β=.90)**	**Gregory (β=.95)**	**Gregory (β=.99)**	**ECM**
Regressand	ΔR_t	ψ_t	ψ_t	ψ_t	ΔR_t
Regressors					
$(y\text{-}x\theta_{OLS})_{t-1}$	-0.14	-0.14	-0.15	-0.16	-0.12
	(0.06)	(0.16)	(0.16)	(0.17)	(0.05)
ΔQ_t					1.43
					(0.35)
ΔP_t					0.26
					(0.58)
ΔY_t					1.42
					(1.27)
ΔR_{t-1}	0.15				
	(0.11)				

Note: In estimating the Dolado approach using OLS, we included the two instruments reported, the dependent variable lagged again and lagged first differences of the forcing variables. We do the same when we impose the discount factor. This is consistent with the approach of Gregory et al. (1993). $\psi_t = \Delta y_t - \beta \Delta y_{t+1}$. Standard errors are in parenthesis.

factor is small given the values that researcher usually attribute to it.[14] We can interpreted this as evidence that the central bank is not especially forward looking, and substantially discounts future losses associated with the cost function. This myopic behaviour could be a reflection of the inherent difficulty that a central bank with a demand for reserves equation containing the exchange rate will have forecasting that variable. This result is not surprising given the difficulty in the literature since Meese and Rogoff (1983) to produce models of exchange rate determination which can outperform a random walk "out of sample".

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[14] Ilmakunnas (1989) using a LQAC model with survey expectations to study labour demand in Finnish manufacturing, obtained an estimated discount factor of 0.911 giving a 9.8% discount rate.

As described before, there is an identification problem associated with the Dolado approach. We consequently impose a discount factor consistent with past studies and find that irrespective of its magnitude, the estimated adjustment coefficients were equivalently to the other approaches. They did have smaller estimated t-statistics of around one, hence were not significant. The Error Correction Model produced a similarly sized adjustment coefficient to the Dolado IV equation. This approach also indicated a significant short run association between the exchange rate and reserves.[15]

In summary our single equation results suggest the following. We find consistent evidence that *Q* and *Y* enter the long-run relationship for reserves. There is also evidence of leaning against the wind. Reserve changes are sterilised, since prices enter insignificantly in the long-run. Across a number of techniques the estimated adjustment to disequilibrium is slow with a half live of four quarters. The results below from the Unrestricted VAR suggest that Q, P and Y were weakly exogeneous with respect to a long-run relationship for reserves. This supports the use of the single equation approach. Nevertheless as a useful means of comparison we examine the multiple equation approach.

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[15] We can directly compare the ECM to Wannacott's (1965) approach in modelling leaning against the wind.

2.5 The Multiple Equation Results

2.5.1 A Congruent VAR

Our first objective is to obtain a statistically well specified or congruent VAR. The Hannan-Quinn (HQ) and the Schwarz Information Criteria (SC) (see Hendry and Doornik, 1994) suggest that a VAR with three lags is dominated by a one lag specification. Using the latter would have proven unsatisfactory, since it removes the possibility of interpretation of short-run effects. Given that some of the variables lagged three times were significant in the individual equations and that an F-test of model reduction from three to two lags was barely acceptable, we retained three lags in our system. We also retain the longer lag because Cheung and Lai (1993) suggest that under parameterisation can result in severe distortions of the cointegration tests, whereas choosing an inappropriately long lag length does not. The diagnostic tests of the residuals, included in Table 2.6, are: AR1-5, a F-test of no autocorrelation

Table 2.6: Misspecification Tests

Multivariate Tests

	AR 1-5 F(80,175)	**Normality $\chi^2(8)$**
	1.115 (p=0.276)	78.327 (p=0.00)

Univariate Test

	ARCH(4) F(4,59)	**Normality $\chi^2(2)$**	**AR1-5 F(5,62)**
ΔR	0.811 (p=0.523)	36.104 (p=0.000)	1.507 (p=0.201)
ΔQ	0.299 (p=0.877)	14.282 (p=0.000)	0.786 (p=0.564)
ΔP	0.519 (p=0.722)	8.003 (p=0.018)	0.506 (p=0.771)
ΔY	2.082 (p=0.095)	16.216 (p=0.000)	1.320 (p=0.267)

Note: Probability values are in parenthesis.

between 1 and 5 lags; a test of no ARCH (Autoregressive Conditional Heteroscedasticity) distributed as F(4,59); and a test for normality, distributed as χ^2. For a review of these tests see Hendry and Doornik (1994).

The misspecification results in Table 2.6 indicated that there were no problems with Autocorrelation and ARCH at a univariate level and also no problem with the former test at a multivariate level. There did however appear to be a problem with normality using both multivariate and univariate tests. Cheung and Lai (1993) suggest that the trace test statistic shows more robustness to residual skewness and excess kurtosis, and hence non-normality, than the other Johansen (1988) test for cointegration, the maximum eigenvalue test. Our results are therefore based purely on the trace statistic. Normality was slightly reduced by the inclusion of dummies, to remove some observational outliers. However, as we are particularly interested in the extreme circumstance which outliers represent, we retain these observations.[16]

Table 2.7 gives the cointegrating vectors in standardised form (i.e. the leading diagonal of β equals 1) and their associated loadings. We should notice, in particular, that α_1 is significant in all the equations but the first, and α_2 is only significant in the first equation. The absences of cross equation linkages, on the basis of the adjustment coefficients,

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[16] Indeed the Johansen technology is based upon Gaussian likelihood but the asymptotic properties of the method only depends upon the assumption that the errors are identically and independently distributed (i.i.d.). Thus the normality assumption is not so crucial relative to the ARCH and AR test results.

Table 2.7: Systems Cointegration Analysis

Standardised Eigenvectors (β'_i)

	R	*Q*	*P*	*Y*	*constant*
β_1	1.000	-20.189	31.465	-53.848	1513.730
β_2	4.334	1.000	-2.360	-9.962	160.716
β_3	0.173	-1.873	1.000	-3.240	92.049
β_4	0.015	0.037	-0.402	1.000	-27.538

Standardised Adjustment Coefficients (α_i)

	α_1	α_2	α_3	α_4
ΔR	0.007 (1.914)	-0.040 (-3.568)	0.134 (1.221)	0.229 (1.112)
ΔQ	0.004 (3.391)	0.006 (1.616)	0.080 (2.200)	0.072 (1.041)
ΔP	-0.001 (-3.211)	0.001 (1.325)	0.000 (0.058)	0.024 (1.822)
ΔY	-0.001 (-3.170)	-0.000 (-0.490)	0.018 (2.097)	-0.022 (-1.351)

Note: t-statistics are in parenthesis.

implies our three explanatory variable in the single equation approach (*Q*, *P* and *Y*) are weakly exogenous with respect to the demand for reserves equation, where this is represented by the second relationship.[17]

This is evidence in favour of the validity of the single equation estimators used in Section 2.4. It is not sufficient to move to a conditional model as we firstly have to identify our long-run relationships (see Doornik and Hendry, 1994). Therefore we retain the system method when testing

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[17] This hypothesis was then tested using a likelihood ratio test under the null that the cointegrating vector only entered the first equation in our VAR. The estimated test statistic was $\chi^2(2)=2.61$ with a p-value of 0.27 suggested we could accept the null. However Engsted and Haldrup (1994) argue that to use the single equation approach we also require that ΔR should not Granger cause ΔQ, ΔP and ΔY. Granger Causality tests suggested this condition was not satisfied. The fear that reserves are also weakly exogenous to the demand for reserves relationship, suggested in Section 2.3.4 is also

structural hypothesis in the cointegrating space and to examine the short-run dynamics.

2.5.2 A Parsimonious VAR

In Table 2.2, the trace statistic confirms that there are two cointegrating vectors. This is supported by the plots our two cointegrating relationships $\hat{v}_i' X_t$ and the plots of the cointegrating relationships with the short run dynamics removed $\hat{v}_i' R_{kt}$, see Johansen and Juselius (1992, p221). The graphs look reasonably stationary (see Appendix 2.C) - although there is quite substantial difference between the two graphical representations of the first cointegrating vector.[18]

We normalise our vector due to initial suspicions that they represent a long-run real exchange rate relationship and a demand for reserves relationship. Consequently we have β_1 =(-0.05, 1.00, -1.56, 2.67, -74.98) and β_2 =(1.00, 0.23, -0.54, -2.30, 37.08). The first cointegrating relationship does not have the characteristic of a demand for reserves

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ameliorated. If anything the evidence in Table 2.7 indicates reserves are weakly exogenous with respect to the first cointegrating vector.

[18] This has two possible interpretations. It could represent an inherent tendency in the model to move towards equilibrium without ever reaching it, because of large and frequent shocks pushing it away from equilibrium (Johansen and Juselius, 1992). Alternatively it could be because the data is second order and not first order non-stationary (Hansen and Juselius, 1994). Prices are an obvious candidate for I(2)ness. The plots of the levels and first differences of prices, included in Appendix 2.C, indicate that this variable is not I(2), but the time series may suffer from a structural break when we move from the high inflation 70s to the low inflation 80s. The plots of the characteristic roots suggested that I(2) was not a problem i.e. no roots are outside the unit circle.

equation, given the evidence in Section 2.4.1. For example, the elasticity of our explanatory variables with respect to reserves are all very large and income is wrongly signed. It may however represent an equation explaining the real exchange rate. For example, an increase in domestic prices leads to an appreciation of the real exchange rate. This follows when the real exchange rate is defined as:

$$q = -s - p^* + p$$

p and p^* are the domestic and foreign price level respectively, and s is the nominal exchange rate defined in foreign currency units per unit of domestic currency.[19] Our real exchange rate equation only depends upon price and income. Reserves have only a small impact in this equation.

The second equation is slightly more appealing as a demand for reserves equation. The estimated coefficient or eigenvalue associated with *Y* is consistent with our single equation estimates and although *Q* and *P* appear 'wrongly' signed compared to estimates in Section 2.3, they are not widely different.

2.5.3 Structural Hypotheses Tests

We can examine these issues by imposing structural hypothesis on the cointegrating vector. Firstly, we conduct tests of the form $H_4=\{H\varphi_1,\psi_1\}$ i.e. they test hypothesis about a single relation in the

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[19] This is slightly different from the usual definition of the nominal exchange rate. Here an increase in the nominal exchange rate is equivalent to a domestic current appreciation. This representation was due to the method of calculating *Q* from the IFS. The real exchange rate equation has been altered accordingly.

Table 2.8: Structural Hypotheses

$H_4=\{H\varphi_1,\psi_1\}$	*R*	*Q*	*P*	*Y*	*Const*	$\chi^2(r)$	p-value
Reserves							
H1	1	0	0	0	φ	7.02(2)	0.03*
H2	1	φ	0	0	φ	3.98(1)	0.05
H3	1	0	φ	0	φ	2.09(1)	0.15
H4	1	0	0	φ	φ	0.02(1)	0.89
H5	1	0	φ	φ	φ	0.00(0)	1.00
H6	1	φ	0	φ	φ	0.00(0)	1.00
H7	1	φ	φ	0	φ	0.00(0)	1.00
Exchange rate							
H8	0	1	0	0	φ	13.25(2)	0.00*
H9	0	1	φ	0	φ	6.67(1)	0.01*
H10	0	1	0	φ	φ	9.47(1)	0.00*
H11	0	1	φ	φ	φ	0.00(0)	1.00
Prices							
H12	0	0	1	0	φ	7.01(2)	0.03*
H13	0	0	1	φ	φ	3.88(1)	0.05
Income							
H14	0	0	0	1	φ	16.19(2)	0.00*

Notes: φ represents an unrestricted parameter estimate. * indicates significance at the 5% level.

unrestricted cointegration space (see Johansen and Juseluis, 1992). In particular we examine the exchange rate equation and also the reserve equation. From the structural hypotheses in Table 2.8, we can rule out the possibility that reserves and a constant alone form a cointegrating vector (i.e. H1), at the five percent level.[20] On statistical grounds alone and at the conventional significance level we are indifferent between the hypotheses H2 to H7, since all are insignificant. H2 is barely accepted suggesting that *R*, *Q* and a constant do not form a cointegrating vector alone. They do form a stationary relationship, when combined with income in H6. This

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[20] This would be an indication of the central bank merely attempting to accumulate reserves over time, see MacDonald (1988).

would appear to be the ideal candidate representing the demand for reserves equation. We can justify this choice for three reasons. Firstly, we argued previously that the exchange rate enters the long-run demand for reserves equation. Secondly, we found in Section 2.4.1 that prices are not significant in a long-run relationship consistent with the sterilisation hypothesis. And thirdly, the reserves equation is likely to contain a scale variable. We therefore accept H6 against the other relationships as our statistically justified demand for reserves relationship.

Of the remaining hypotheses, only H11 and H13 are accepted as vectors that span the cointegrating space. H11 could possibly represent the exchange rate equation mentioned above whereas H13 can not and is barely accepted anyway. We in turn, combined the more likely cointegrating vector H11 with the first accepted hypothesis H6, and inspect the resultant estimated coefficients or eigenvalues, to see if they are compatible with the exchange rate interpretation. This results in the following restrictions on β

$$H_5\text{: } \beta = (H_1\varphi_1, H_2\varphi_2)$$

$$\text{where } \beta_1 = (0,1,\varphi,\varphi,\varphi) \text{ and } \beta_2 = (1,\varphi,0,\varphi,\varphi)$$

$$\beta = (R\,,Q\,,P\,,Y\,,\ constant)$$

We restrict reserves to equal zero in the first cointegrating vector and prices to equal zero in the second. This hypothesis was accepted easily (with a test statistic of $\chi^2(,0) = 0.000335$ and probability-value =1.000). The first cointegrating vector does indeed appear to represent a real exchange rate relationship. In Table 2.9, an increase in prices is associated with an appreciation in *Q*, consistent with our definition of the real exchange rate. A decrease in income is also associated with a real exchange rate appreciation. This is likely where a real appreciation has a deleterious effect on domestic output and income - through the export channel - and is consistent with the textbook Mundell Fleming model. Combining H11 and H7 also results in the second cointegrating vector having estimated coefficients equivalent to those for the single equation estimator. In other words, the estimated coefficient for Q in a long-run reserve relationship is now consistent with leaning against the wind.

Having imposed the long-run restrictions on our cointegrating vectors we have consequently arrived at our PVAR. The restricted β vectors and their corresponding adjustment coefficients are included in Table 2.9. Given the significance of α_2 in the short-run equation for reserves (ΔR) the second cointegrating vector only enters the first equation of the PVAR. The first cointegrating relationship on the

Table 2.9: Restricted Cointegrating Vectors and Corresponding Loadings

Eigenvectors

	R	*Q*	*P*	*Y*	*const*
β_1	0.00	1.00	-1.57 (0.20)	2.53 (0.48)	72.32 (12.99)
β_2	1.00	-0.12 (0.49)	0.00	-3.18 (0.56)	62.20 (15.59)

Corresponding Loadings

	ΔR	ΔQ	ΔP	ΔY
α_1	-0.20 [2.64]	-0.07 [2.82]	0.02 [-3.36]	0.02 [-2.86]
α_2	-0.17 [-3.36]	0.03 [1.80]	0.00 [1.07]	-0.00 [-0.70]

Notes: Standard errors are in parenthesis. t-statistics are in brackets.

same basis enters all equations (i.e. α_1 is significant in all short-run equations).[21] With the given restrictions on the long-run relationships, we do not have evidence to suggest that reserves are weakly exogenous with respect to the first cointegrating. This is because β_1 enters the short run equation for reserves. Given the long-run restrictions we can not move to a conditional model, but retain the systems method. This is stronger evidence against the single equation approach to reserve modelling than suggested by Section 2.5.1 (see Harris, 1995).

2.5.3 A Structural Econometric Model

We can now model the PVAR to obtain a Structural Econometric Model (SEM), and consequently assess whether it parsimoniously encompasses the PVAR. Removing all variables with t-ratios of less than two in the PVAR was not initially accepted by an F-test, and some of the remaining t-ratios were no longer significant. Eliminating variables when their t-ratios were insignificant we arrived at the model in Table 2.10.

Table 2.10 FIML of SEM

$$\begin{aligned}
\Delta R_t &= \underset{(0.10)}{0.37}\Delta R_{t-2} - \underset{(0.04)}{0.17}\beta_2 x_{t-1}\\
\Delta Q_t &= \underset{(0.03)}{0.06}\Delta R_{t-2}\\
\Delta P_t &= \underset{(0.10)}{0.22}\Delta P_{t-1} + \underset{(0.00)}{0.02}\beta_1 x_{t-1}\\
\Delta Y_t &= \underset{(0.12)}{-0.28}\Delta P_{t-1} + \underset{(0.00)}{0.01}\beta_1 x_{t-1}
\end{aligned}$$

$$\beta_1 x_{t-1} = Q_{t-1} - 1.57P_{t-1} + 2.53Y_{t-1} + 72.32; \qquad \beta_2 x_{t-1} = R_{t-1} - 0.12Q_{t-1} - 3.18Y_{t-1} + 62.20$$

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[21] These are distinct from the results in Table 2.7, where the second cointegrating vector only appears to enter the first equation. Of course when we impose linear restrictions upon the long-run relationship the estimated adjustment coefficients and their corresponding t-statistics will consequently change.

All t-ratios, obtained from the standard errors in parenthesis obtained from Table 2.10, are significant at the 5% level. Additionally a test of these over identifying restrictions ($\chi^2_{or}(33) = 44.28$, $p = 0.09$) is accepted, so we conclude that the SEM parsimoniously encompasses the PVAR.

The most important implications of the SEM were as follows. Since we include a lagged and differenced *R* in equation one reserves in the short run appeared to have a very strong autoregressive component. This is an argument in favour of the particular specification of the Linear Quadratic Adjustment Costs Model we used before, since this also contains an autoregressive component. Reserves also adjust to disequilibrium in the long-run relationship. The estimated adjustment is similar in size and significance to our previous single equation estimates - around 17% of any deviation from the long-run demand for reserves is corrected in any period. This gives a half-life of four periods (i.e. $(1-0.17)^4$). With reference to the short-run exchange rate equation, lagged changes in reserves have a small but significant estimated coefficient. This is consistent with the Bank of England leaning against the wind i.e. accumulating foreign currency when the nominal, and therefore real, exchange rate is appreciating. Prices and income in the short run are dependent upon lagged prices, and not extensively dependent upon the cointegrating vector.

2.6 Conclusion

In Chapter Two we have implemented the Dolado et al. (1991) approach to estimating a LQAC model for the demand for international reserves and contrasted it with the Johansen Multivariate framework. The LQAC approach involved estimating the adjustment cost parameter, using the first-order condition or Euler equation from the model and Instrumental Variables. Initial estimates of the long-run parameters were obtained taking account of non-stationarity in our data. We conditioned reserves on the real effective exchange rate, prices and income. As the variables in the model are potentially endogenously determined and there may exist serial correlation we utilised Phillips and Hansen's (1990) FM-OLS estimator. This estimator allows classical inference on the parameters of interest since they have an asymptotic normal distribution for their t-statistics.

Exploiting the LQAC approach our model produce some useful results. Our long-run equilibrium relationship satisfied a number of tests for cointegration. The size and sign of the estimated parameters for our long-run relationship, were consistent with theory. For example, an appreciation of the real effective exchange rate was associated with an increase in the level of reserves. This is illustrative of the Bank of England "leaning against the wind" of an appreciating exchange rate - for example in early 1977 and again in 1987. Reserves do not appear to have been used as an instrument to reverse trend movements, only reducing the size, of say, an appreciation. The OLS, FM-OLS and Phillips-Loretan's NLLS estimators suggest that the real effective exchange rate and income are significant in our long-run relationship. However the robust estimators suggest that prices were insignificant in this long-run relationship. This can be interpreted as evidence that reserves had a sterilised impact on the money supply, and had no price implications through a "monetarist channel" in the long-run.

The estimated adjustment coefficients from the Euler equation was not especially large but was statistically significant at around 0.14. The estimated discount factor was quite small, suggesting the behaviour of the BoE is not very forward looking when setting reserves. Given the discussion by Gregory et al. (1993) regarding a possible identification problem, we imposed the discount factor. Irrespective of its imposed size, the estimated adjustment coefficient was around 0.15. The estimated coefficient from the ECM was also equivalently sized. The weak exogeneity tests from the Unrestricted VAR suggest we do not have a problem with the single equation estimator and lend support to these estimated adjustment coefficients. This contrasts somewhat with the results from the Parsimonious VAR, which suggest that we could not move to a conditional model and that we should conduct our analysis with a systems estimator. Additionally our single equation results benefit from comparison with multiple equation estimation, as suggested by Gregory (1994). The Johansen approach indicated that there were two cointegrating vectors between the variables: a real exchange rate relationship and a demand for reserve equation. Excluding prices from the second cointegrating vector, as suggested by long-run sterilisation, resulted in estimated coefficients for Q and Y consistent with our single equation approach.

The dynamic results using Johansen also confirmed slow adjustment towards equilibrium. Lawson (1992) suggests this is a residual from the IMF crisis of 1977: the electorate equates large borrowing from abroad with economic incompetence. That is not to say that the central bank has no access to international capital markets - witness the success of issues of Floating Rate Notes (see New Palgrave, 1994). But when replenishing reserves can be avoided or delayed, the central bank will follow this course of "inaction". There is a strong autoregressive component to reserves as confirmed by the short-run reserve equation. There is also evidence that lagged reserves have a short-run impact on the exchange rate.

As for further research, it would be worth considering the impact of other monetary variables within the Johansen framework. For example we could examine the interaction between the demand for money and for reserves in a multivariate model. This would conceivably be represented by two cointegrating vectors, and would explain changes in reserves as due to deviation from the equilibrium level of reserves and money demand. We return to these issues in the next chapter. It would also be useful to compare the demand for reserves across exchange rate regimes: pre- and post-Bretton Woods.

As for reserves themselves, there may be difficulties modelling foreign exchange intervention and demand for foreign assets by the central bank, where reserves may change for reasons other than intervention. This may be suggestive of including a variable in our demand equation, actually representing intervention. However Edison's (1993) view that "reserves tell us little" seems inaccurate. A study of intervention which uses other proxies for central bank involvement in the foreign exchange market, may not necessarily be immune from flaws.[22] And our results, which provide evidence of a short- and long-run relationship between reserves and the exchange rate, could be used to assess the effectiveness of intervention.

Looking forward from our results, one of the benefits of European Monetary Union is that it allows participating countries to economise on reserve holdings (see Emerson, p183, 1992). For example, if reserve holdings are a reflection of the size of foreign transactions, a reduction in trade conducted out-with your own currency area means a reduction in demand for reserves. The other possible reason why there could be a fall in demand, is that the G3 make an agreement for explicit credit lines between the major central banks. In the absence of such an agreement and despite the fall in external trade, my approach suggests there is unlikely to be a fall in demand for foreign assets. The European Central Bank will continue to hold reserves on the basis of the income of the EU and also because of fluctuations in the external value of the euro.

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[22] See for example Klein (1992) on a discussion on inaccuracies using newspaper reports of intervention.

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Appendix 2.A: Data Description

Table 2.A.1: Description of the Data Set in Chapter Two

Variable Title	*Line in IFS*	*Variable Abbreviation*
International Reserves (US Dollars)	line 11d	*R*
Real Effective Exchange Rate	line reu	*Q*
GDP(1990=100)	line 99b.r	*Y*
GDP(market prices)	line 99b.c	
GDP Deflator	GDP(market prices) /GDP(1990=100)	*P*

Notes: All data series are from IFS, quarterly, seasonally adjusted and transformed into natural logarithms.

Appendix 2.B: Stationarity of Data

Table 2.B.1: Univariate Unit Root Results

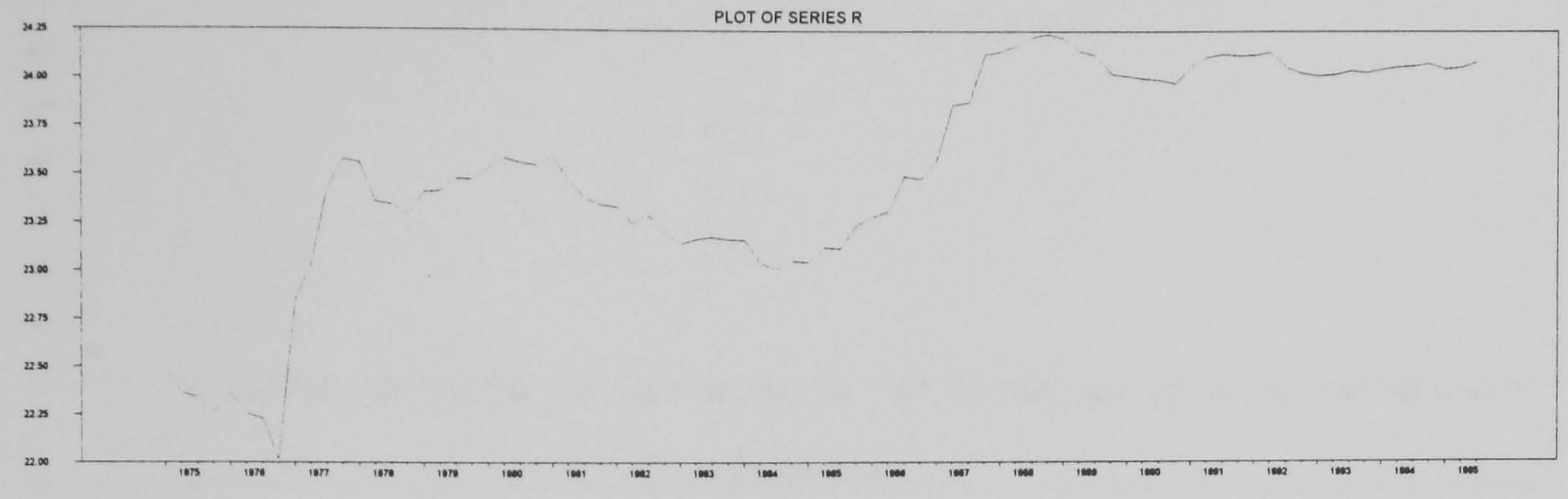
Variable	*lags*	τ_τ **test**	Φ_3	τ_μ **test**	Φ_1	τ **test**
R	2	-3.09	5.11	-2.40	3.44	0.98
Q	6	-2.40	3.78	-2.27	3.83	0.33
P	10	-2.20	3.67	-2.26	2.79	-1.52
Y	6	-2.17	4.74	-2.74	8.52	0.80
5% critical value		-3.43	6.49	-2.89	4.71	-1.95

Notes: τ_τ tests the null hypothesis of unit root with trend in mean. Φ_3 tests the joint null hypothesis of unit root and zero trend. τ_μ tests the null hypothesis of unit root and constant mean. Φ_1 tests the null hypothesis of zero constant and a unit root. Critical values are from Fuller (1976) and Dickey and Fuller (1981).

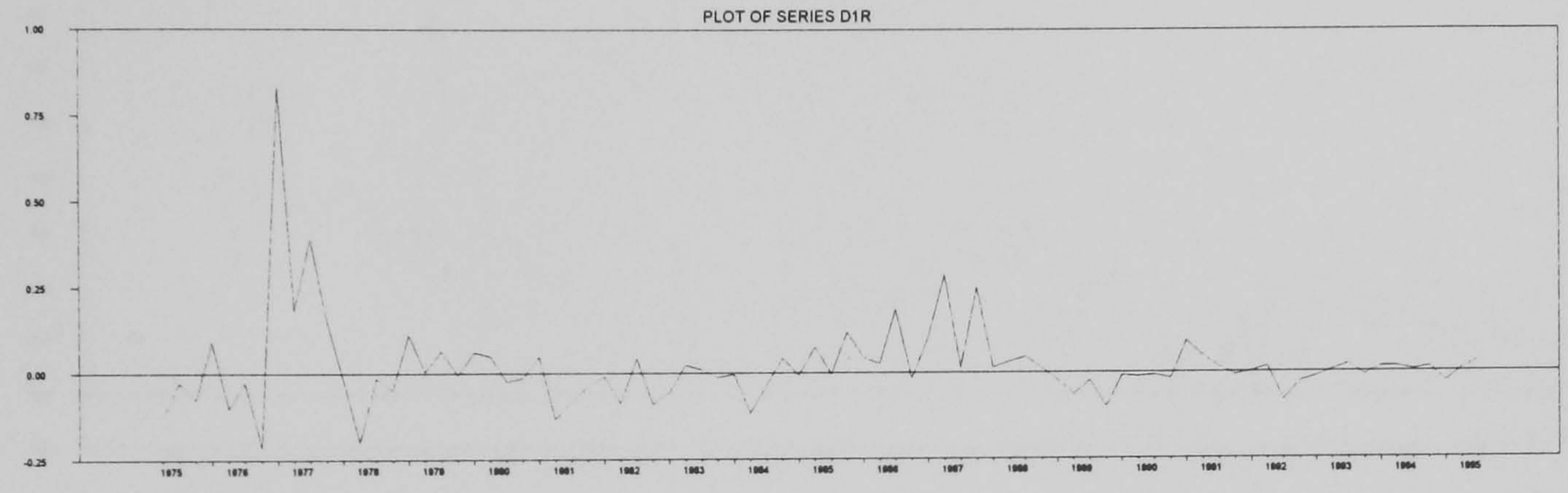
Table 2.B.2 Characteristic Roots Of The PVAR

NUMBER	ROOT	ABSOLUTE VALUE
1	(0.986, -0.000)	0.986
2	(0.869, 0.123)	0.878
3	(0.869, -0.123)	0.878
4	(0.809, -0.000)	0.809
5	(0.570, 0.000)	0.570
6	(-0.567, 0.000)	0.567
7	(0.485, -0.149)	0.507
8	(0.485, 0.149)	0.507
9	(-0.032, -0.428)	0.430
10	(-0.032, 0.428)	0.430
11	(-0.347, 0.000)	0.347
12	(-0.161, 0.000)	0.161

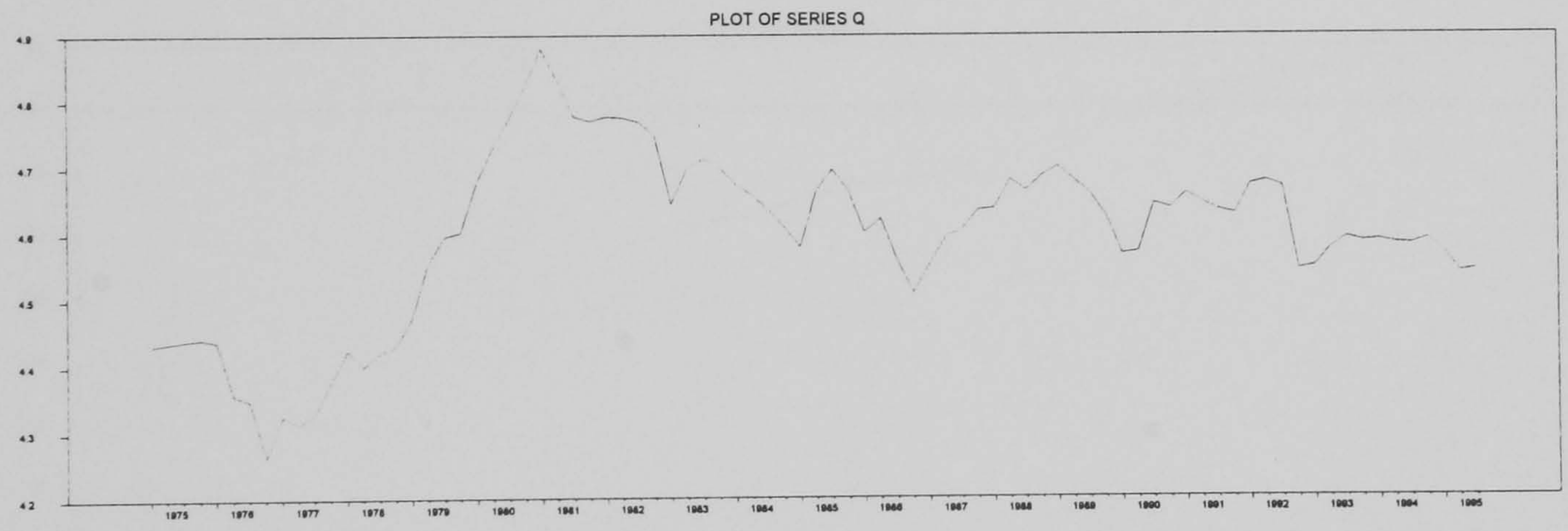
Appendix 2.C: Data Series and Cointegrating Relationships



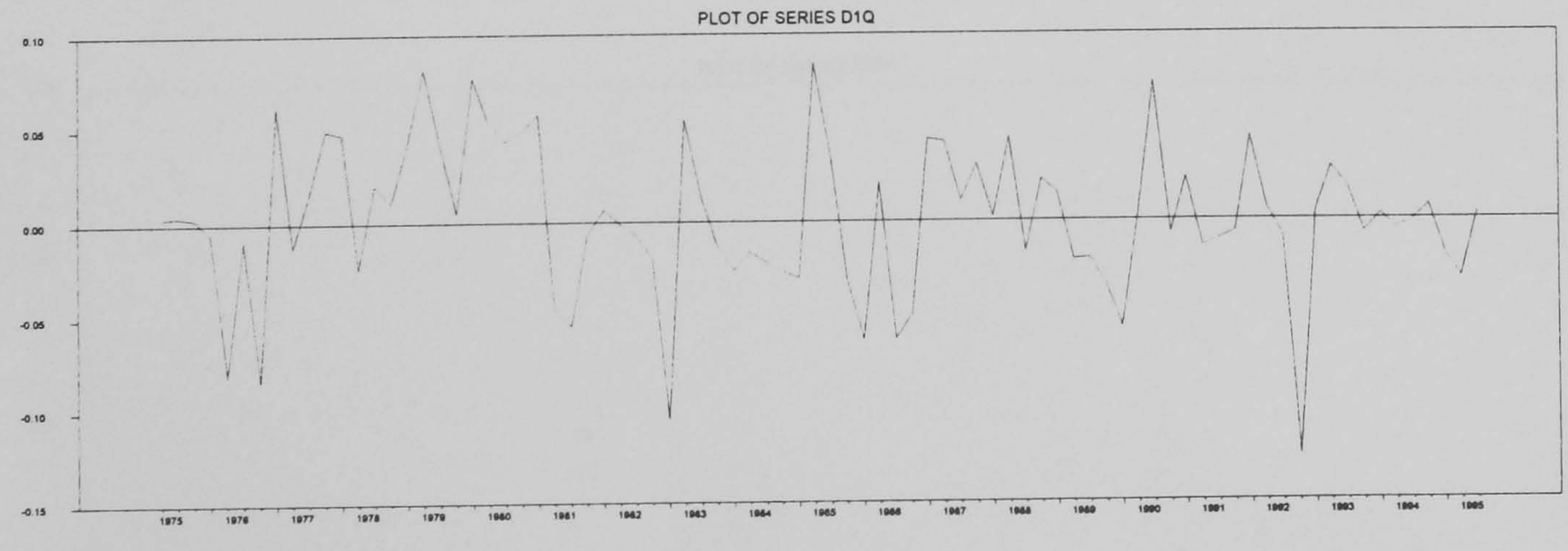
PLOT OF SERIES R



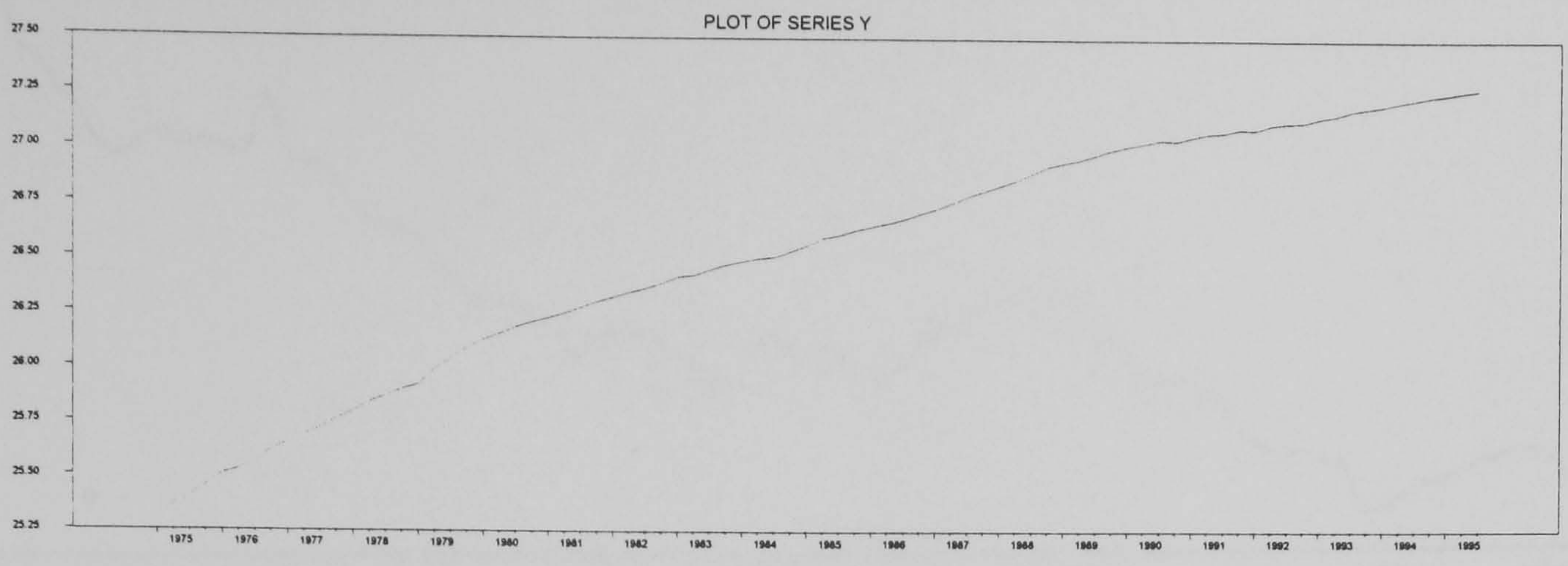
PLOT OF SERIES D1R

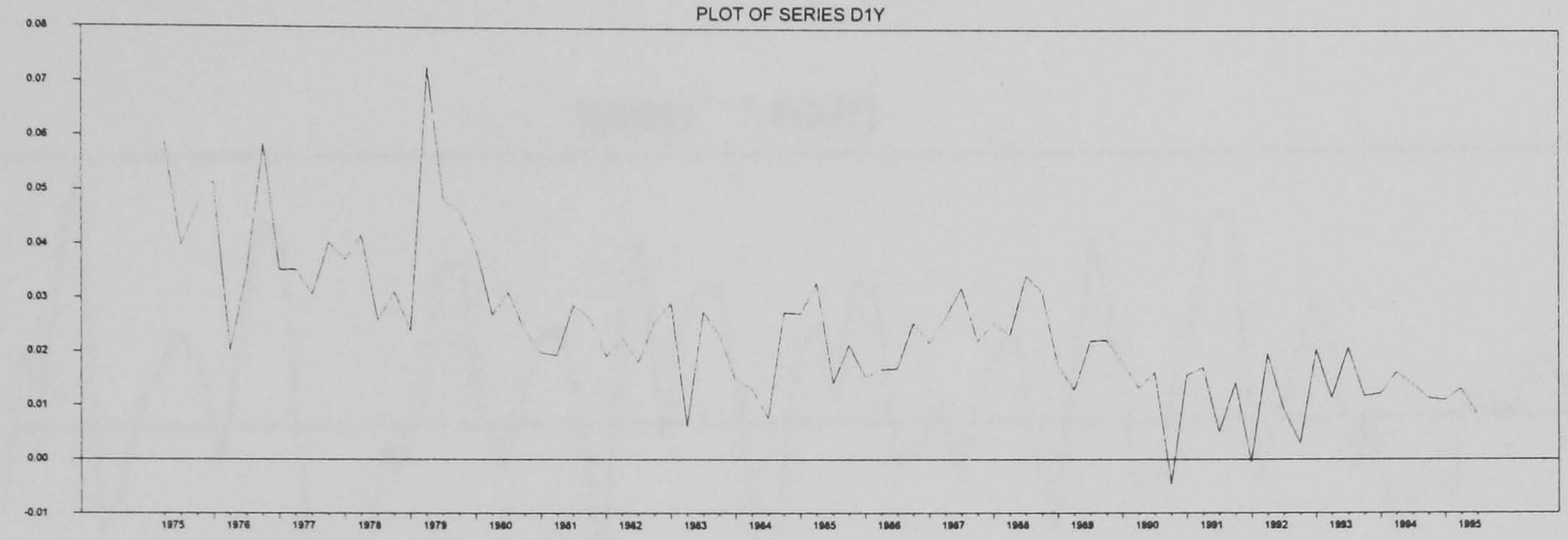


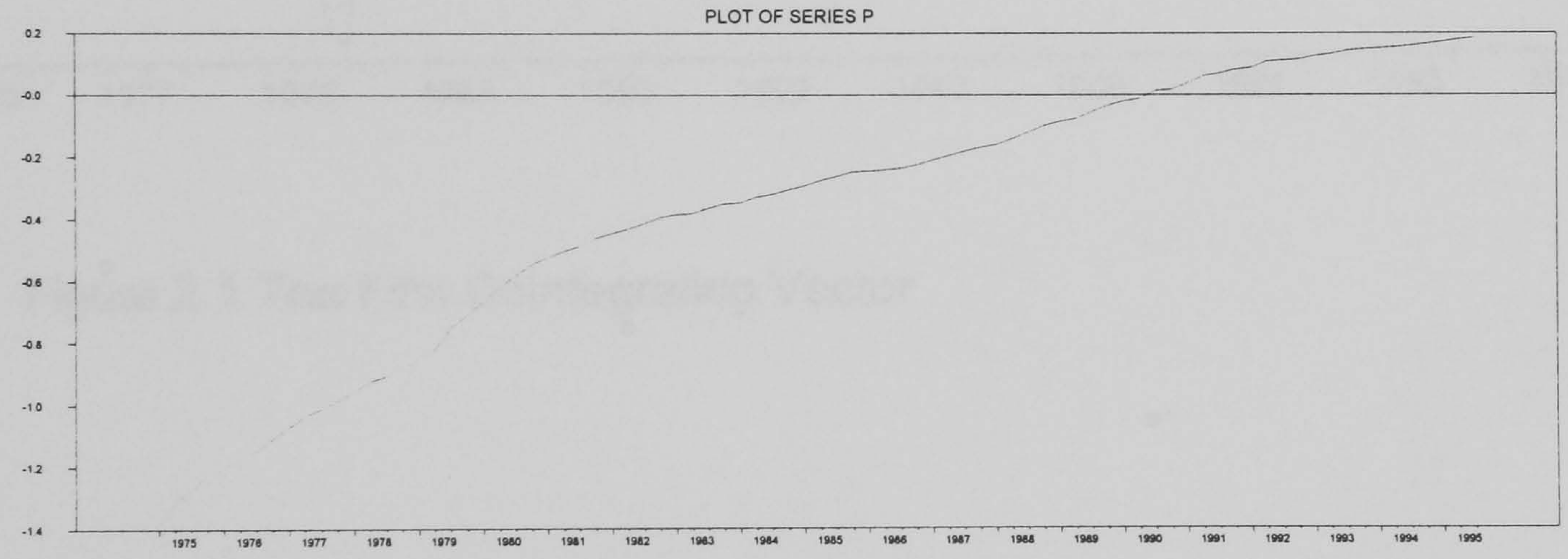
PLOT OF SERIES Q

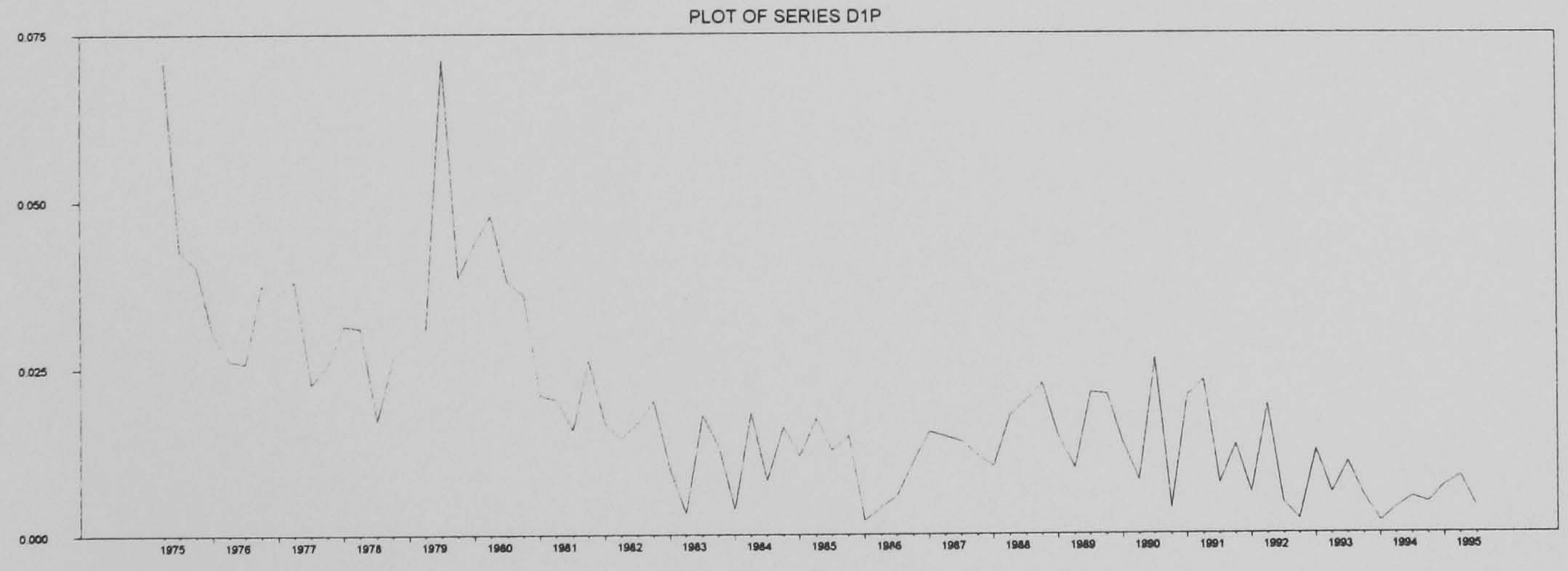


PLOT OF SERIES D1Q


PLOT OF SERIES Y
27.50
27.25
27.00
26.75
26.50
26.25
26.00
25.75
25.50
25.25
1975
1976
1977
1978
1979
1980
1981
1982
1983
1984
1985
1986
1987
1988
1989
1990
1991
1992
1993
1994
1995


PLOT OF SERIES D1Y
0.08
0.07
0.06
0.05
0.04
0.03
0.02
0.01
0.00
-0.01
1975
1976
1977
1978
1979
1980
1981
1982
1983
1984
1985
1986
1987
1988
1989
1990
1991
1992
1993
1994
1995


PLOT OF SERIES P
0.2
-0.0
-0.2
-0.4
-0.6
-0.8
-1.0
-1.2
-1.4
1975
1976
1977
1978
1979
1980
1981
1982
1983
1984
1985
1986
1987
1988
1989
1990
1991
1992
1993
1994
1995


PLOT OF SERIES D1P
0.075
0.050
0.025
0.000
1975
1976
1977
1978
1979
1980
1981
1982
1983
1984
1985
1986
1987
1988
1989
1990
1991
1992
1993
1994
1995

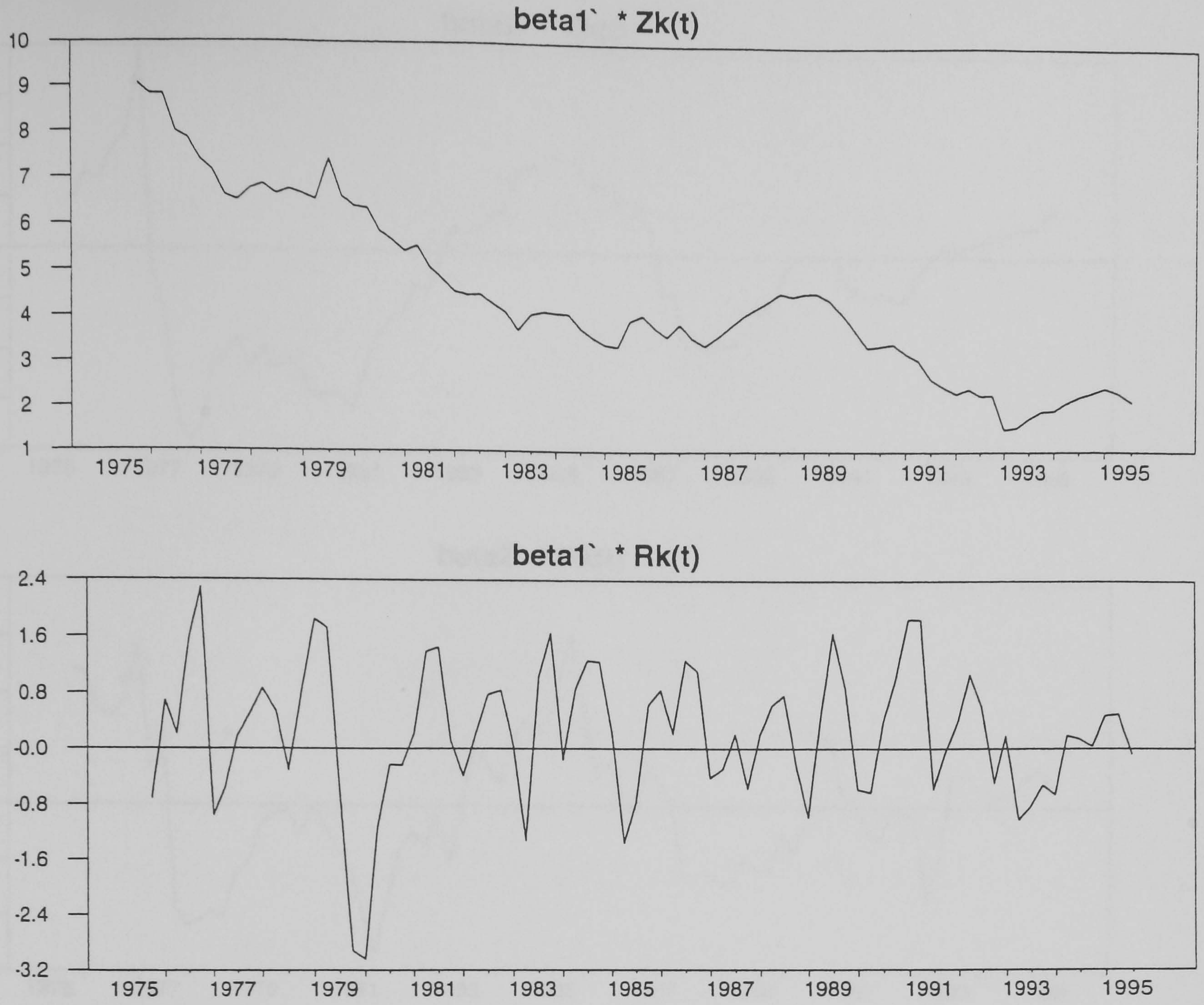


Figure 2.1 The First Cointegrating Vector

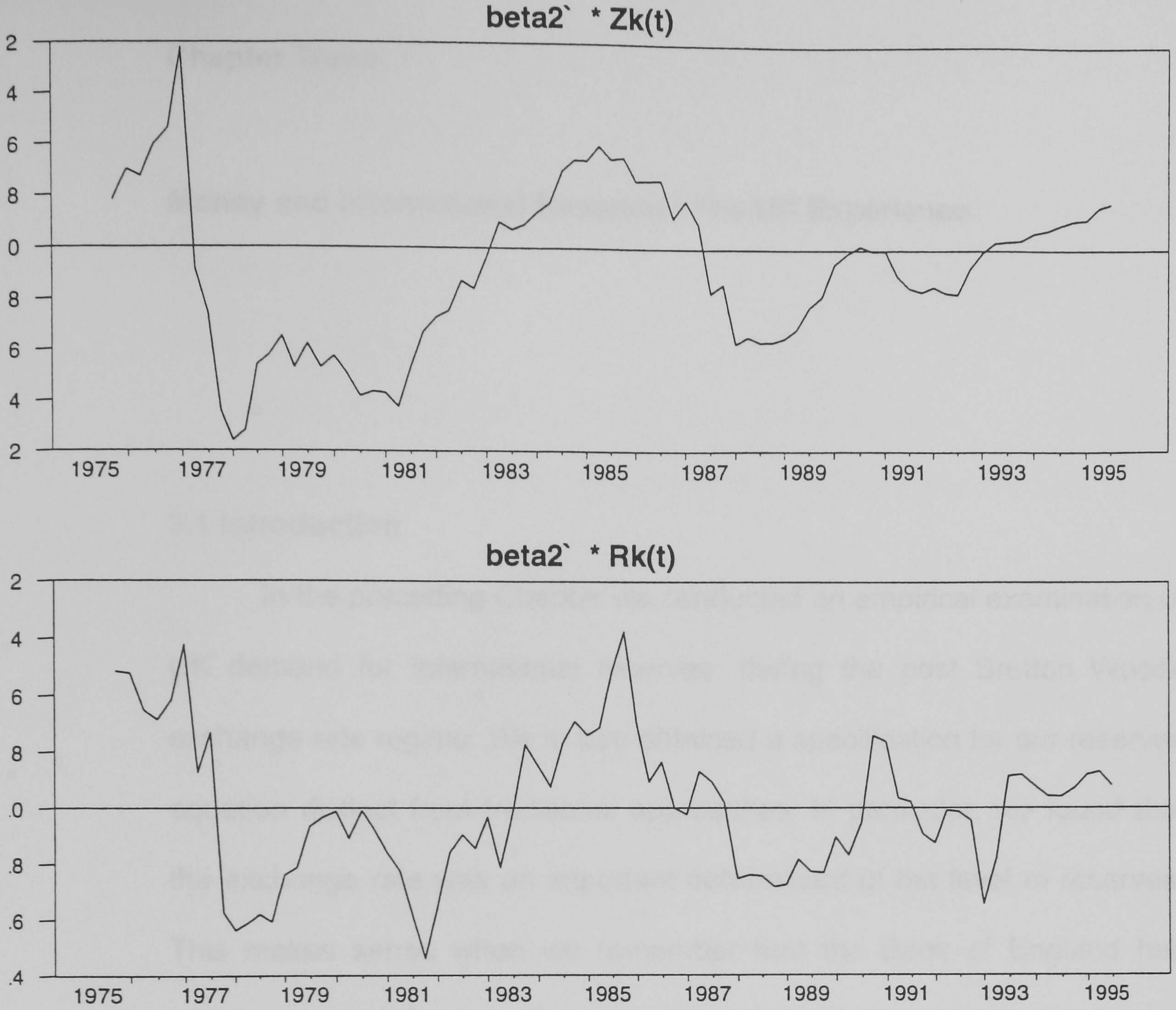


Figure 2.2: The Second Cointegrating Vector

Chapter Three

Money and International Reserves: The UK Experience.

3.1 Introduction

In the preceding Chapter we conducted an empirical examination of UK demand for international reserves, during the post Bretton Woods exchange rate regime. We in turn obtained a specification for our reserves equation distinct from traditional approaches. In particular, we found that the exchange rate was an important determinant of the level of reserves. This makes sense when we remember that the Bank of England has attempted to influence the external value of the pound by foreign exchange intervention. In addition, reserves were dependent upon income and to a lesser extend upon prices, representing the other objectives of macroeconomic policy. We also examined short-run changes in reserves using single equation and Vector Error Correction models. Reserves were dependent upon disequilibrium from the long-run reserve relationship.

However, Jacob Frenkel (1978) has noted that this disequilibrium approach for examining short-run changes in reserves may be misspecified, as it fails to take account of the domestic monetary situation. Changes in reserves, according to this view, are dependent upon the difference between money demand and money supply. This idea can be traced back to the literature on the Monetary Approach to the Balance of Payments. Reserves were central to this theory's transmission mechanism: money market disequilibria was removed by changes in reserves, which in turn, changed the money base and removed any monetary imbalance.

Although some researchers have estimated short-run equations for reserves paying attention to monetary disequilibria (see Frenkel, 1983; Edwards, 1984; Elbadawi, 1990; Ford and Huang, 1994), they generally have not used state-of-the-art econometric methods, which take account of stationarity issues within a multivariate framework.[23] We examine these issues by utilising Johansen's (1988) Full Information Maximum Likelihood estimator. Firstly, we attempt to construct long-run relationships, based on cointegration analysis, for international reserves and the demand for money. The latter is of interest in its own right, given the difficulty researchers have had finding evidence of such a long-run relationship. We then use these to produce a short-run model, with changes in reserves dependent upon disequilibria from long-run reserves and money. This is an indirect test of whether the Monetary Approach to the Balance of Payments is applicable to the UK and to whether our previous short-run equations are misspecified.

Additionally, we investigate the interaction of a number of UK macroeconomic time series using impulse response functions as

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[23] This is especially pertinent given the concern we had regarding weak exogeneity in our systems approach in the previous Chapter.

advocated by Christopher Sims (1980). Such evidence enables us to examine issues related to the direction of causality between reserves and money. This is of importance given some recent disagreement in the literature on the degree of sterilisation of foreign exchange intervention and hence whether changes in reserves are leading to changes in the domestic monetary situation.

The structure of Chapter Three is as follows: Section 3.2 introduces the economic models which we rely upon to conduct our analysis of the interaction between international reserves and money; Section 3.3 briefly reviews the statistical model which forms the basis of the Johansen method; Section 3.4 conducts an empirical investigation into a stable demand for money function; Section 3.5 contains our main empirical investigation into the UK open economy based on the Johansen approach and impulse response functions; Section 3.6 concludes.

3.2 Theoretical Issues

The first theoretical model that we examine in this Chapter is the Monetary Approach to the Balance of Payments (MABP). In part derived from David Hume's price-specie flow model, the most recent models have been developed by Meade, Johnson and Frenkel.[24] This is a useful starting point to consider the relationship between international reserves and money since it grounds our empirical analysis in a robust theoretical

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[24] See Isard (1995) for an introduction to the Monetary Approach to the Balance of Payments. See also Frenkel and Mussa (1985) for extensions to the analysis.

model. The MABP is concerned with how disequilibria in the domestic money market are ameliorated by adjustment in the balance of payments, as represented by changes in foreign exchange reserves. For instance, with demand for money greater than money supply, domestic expenditure levels will fall due to a relatively high interest rate. Where expenditure is less than the equilibrium level of income the balance of payments will be in surplus. Given a fixed exchange rate and low international capital mobility, this surplus will result in an inflow of foreign exchange and the central bank increasing its holdings of reserves. This in turn increases the stock of high powered money, assuming a typical central bank balance sheet. Money supply will rise to equal money demand.

This approach provides a clear suggestion as to the short run determinates of reserves, which can be represented by a partial adjustment equation:

$$\Delta R = \alpha(M^* - M) \qquad (3.1)$$

R is the level of central bank foreign exchange reserves, M is the domestic money stock and $M^* = \beta x$ is the demand for money. Delta (Δ) is the first difference operator. We use a partial adjustment equation since excess demand may feed through to changes in reserves with a time lag. The traditional models assume fixed exchange rates, hence all adjustment is due to changes in reserves. Interestingly, this model could be applied to a managed float, where some adjustment is through the exchange rate. This is directly comparable to studies of the Bretton Woods exchange rate system since there was still a degree of flexibility although exchange rates were constrained within bands.

The MABP also assumes a stable demand for money function. The existence of such a relationship has been the subject of much controversy and empirical work. Notable amongst work refuting the money demand equation has been Goldfeld's (1976) study, using US data. Recent attempts to revitalise the theory have utilised cointegration techniques to establish a long-run equation for money demand. As the basis of our short-run monetary disequilibrium equation and of independent interest we attempt to find a cointegrating relationship for narrow money, M0. M0 is defined as notes and coins in circulation, *plus* bank operational deposits at the Bank of England. Unlike other papers on the demand for M0 we do not introduce a cumulative interest rate measure of financial innovation, which have been criticised as *ad hoc*. Instead we model the real demand for money using income, interest rates and inflation. We focus on M0 since the MABP gives a place of central importance to the central bank balance sheet. This suggests that the domestic money base is equal to the central bank's stock of international reserves and assets denominated in domestic currency units. The accounting identity is the primary mechanism by which changes in reserves affect the money supply. A useful approximation for the money base is M0.

The short-run equation for reserves based on the MABP contrasts with the empirical literature on the demand for international reserves. The latter traditionally presumes an underlying long-run demand for reserves equation, which forms the basis of an error correction model. The demand for reserves is a function of a number of variables: including real income, the variability of reserves and the average propensity to import (see Chapter 2 for a review of this traditional approach). Frenkel (1978) suggested that these long-run equations may be misspecified, and should reflect the determinants of reserves in a floating regime. We therefore estimated an alternative specification for long-run reserves, dependent upon the real effective exchange rate, real income and inflation. Given that the central bank was unlikely to be continuously on its demand schedule, the literature utilised a partial adjustment framework, where changes in reserves represented adjustment towards a long-run target level (*R**). This can best be represented by an equation of the form:

$$\Delta R = \alpha(R^* - R) \qquad (3.2)$$

The Monetary Approach to the Balance of Payments and the empirical literature of the Demand for International Reserves were unified by the suggestion of Frenkel (1978) and the models of Frenkel (1983) and Edwards (1984). Reserve changes were due to excess demand for reserves by the central bank and by excess demand for money by the public. The partial adjustment model therein was of the form:

$$\Delta R = \alpha_1(R^*-R) + \alpha_2(M^*-M) \qquad (3.3)$$

It was argued that by excluding the term representing excess demand for money, early empirical attempts to estimate the adjustment coefficient α_1 were biased (see Frenkel 1983, p82). Indeed, there may still be specification problems with equation *(3.3)*, since the model expresses all change in reserves are due to excess demand for reserves by the central bank and excess demand for money by the public. We also consider whether there is some effect from changes in other determinants, like income or the exchange rate. Our statistical model will take account of these factors below. Nevertheless, equation *(3.3)* forms the basis of our short-run results.

The third issue that we consider is related to the literature on central bank foreign exchange intervention. In particular, this considers the extent to which foreign exchange interventions, as proxied by changes in reserves, are sterilised. That is to say the degree to which changes in the bank's portfolio of foreign assets are counteracted by changes in their portfolio of domestic assets, such that intervention has no implications for domestic monetary conditions. The assumption of the monetary model was that changes in reserves were unsterilised, whereas Neumann and von Hagen (1993), Glick and Hutchison (1995) and McKinnon (1996) find that there is a degree of - but not complete - sterilisation for Germany, Japan and the US respectively. This is also of interest since it gives us an indication of the exact nature of intervention, and hence whether we can expect it to have any impact. The impact of sterilised intervention is controversial since it does not change the money supply, whereas unsterilised intervention is generally believed to have some impact on the exchange rate (see Obstfeld and Rogoff, 1995). If we find that changes in reserves are partially unsterilised, intervention may be having quite a substantial effect.

Our empirical objectives in this Chapter are as follows: to establish a long-run equation for money demand; to combine this with a long-run equation for international reserves and obtain a short-run model; and to use this model to examine the dynamic interaction between our variables (paying particular attention to the interrelationship between money and international reserves).

3.3 The Statistical Model

Our statistical approach is based on the Johansen (1988) approach. This begins with a general VAR, in Vector Error Correction Model (VECM) formulation, to which we impose long-run restrictions based on economic theory. These represent our long-run equilibrium relationships. We then go on to investigate the short-run adjustment to these long-run relationships.

Our VECM is as follows:

$$\Delta x_t = \sum_{i=1}^{p-1} \Gamma_i \Delta x_{t-i} + \Pi x_{t-1} + \psi D_t + v_t$$

where x_t is a vector of variables, D_t can contain a constant and/or trend. ν_t is assumed to be mean zero, normally distributed, homoscedastic and serially uncorrelated. The order of the VAR is finite and the parameters Γ, ψ and Σ (the covariance matrix of ν_t) are assumed constant. The rank of the long-run responses (Π), is evidence of the number of cointegrating relations (see Johansen, 1988). The rank of Π is calculated by the trace statistic. We also examine graphical evidence to interpret the number of cointegrating relationships. The long-run responses themselves can be reparameterised as

$$\Pi = \alpha\beta'$$

which represents the cointegrating vectors (β) and associated adjustment matrix (α).

Next, in our modelling strategy, we impose restrictions on the cointegrating long-run relationships. This is an attempted to produce "meaningful economic relationships", which underlie the long-run model. We also interpret the adjustment towards these equilibrium relationships.

3.4 The Demand for Money

We are interested in the demand for money since it underlies the synthesis of the Monetary Approach to the Balance of Payments and empirical studies of the demand for international reserves. Disequilibrium in the domestic money market may have implications for international reserves. Nevertheless, the demand for money is of independent interest given the problems in obtaining a stable relationship using a traditional specification. In this section we firstly review the traditional approach, then attempt to produce a cointegrating long-run relationship for UK M0.

3.4.1 Theoretical Issues

The most popular static representation in empirical studies has the demand for real money balance as a linear function of a scale variable and an opportunity cost measure. For example, Goldfeld (1973) postulates a relationship of the form

$$M^* = P\,f(Y, i) \qquad (3.4)$$

were M^* is the demand for money, P is the domestic price, $f(.)$ is a function, Y is a scale variable and i is a measure of the opportunity cost of holding money, usually represented by a domestic interest rate. *A priori* it is unclear which particular economic times series should be used from the multitude of possible money, interest rates and income measures. The imposition of long-run price homogeneity in equation *(3.4)* is based on the assumption that the units of currency are irrelevant with respect to demand for money. What matters is real demand, which can consequently be used to purchase good and services. All variables have been transformed into natural logarithms, apart from the interest rate measure. We follow this semi-logistic form by making the implicit assumption that a 100 basis point, or one percentage point, increase in interest rates leads to the same percentage reduction in the quantity of money demanded, irrespective of whether it is added to the base rate at 5% or 10% (see p265, Friedman and Schwartz, 1982). This produces a linear equation of the following form

$$log(M^*/P) = constant + \beta_1 log\,Y + \beta_2 i + u \qquad (3.5)$$

where the operative *log* means the variables have been transformed into natural logarithms and u is a stochastic disturbance term.

However this stylised representation has met with widespread dissatisfaction. This can be traced directly to Goldfeld's (1976) modelling of M1 in the US over the period 1952-73.[25] The empirical model estimated in his paper performed badly when attempting to predict the level of money balance out of sample over the period 1974-76. This period has consequently been known as "the case of missing money". These results were also found in a comparable study of the UK by Artis and Lewis (1976).

At a general level, demand for money relationships were believed to be excessively unstable. Cuthbertson and Barlow (1991) delineate four main reasons for this instability: financial innovation; measurement problems; misspecified dynamics; and the role of money as a buffer stock. In this paper we concentrate on M0 and examine empirical attempts to produce a stable demand representation based upon financial innovation. We are interested in M0 since Frenkel (1983) suggests it is the pertinent measure of money, when examining balance of payments issues. This becomes the focus of our attention below.[26] M0 has also become of some operational significance for UK monetary policy: the Bank of England has

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[25] The macroeconomic time series used by Goldfeld included real GNP and the commercial bill rate.

[26] Specifically, Frenkel (1983) proposes that we should be interested in the Money Base in this situation. Artis and Lewis (p172, 1991) dismiss the difference between M0 and the money base as of no operational significance. However, Breedon and Fisher (p371, 1996) take a contrary viewpoint emphasising that although the difference between the two, bankers operational deposits at the Bank of England, contributes less than 1% of M0, the high volatility of bankers' balances contributes disproportionately to changes in M0.

a monitoring range for growth in M0. Additionally, there is evidence to suggest that this narrow monetary aggregate works as a leading indicator of future nominal spending[27] (see Astley and Haldane, 1995, and Henry and Pesaran, 1993).

Recent UK research on M0 has followed the approach of Hall, Henry and Wilcox (1989). Evidence of instability was based on the failure to reject the null of no cointegration between real M0, a scale and an opportunity cost measure using Engle and Granger (1987) and Johansen (1988) tests. Hall et al. (1989) suggest that "instability" is due to financial innovation. That is, the increase in the availability of cash through ATMs, the increase in the number of people's salaries being paid directly into their bank accounts, the introduction of interest rates on deposit accounts and the increased availability of credit cards as a means of payment. Once these developments are modelled using a cumulative interest rate, Hall et al., produce a cointegrating relationship.

In contrast to the line of research using a cumulative interest rate measure has been work by Muscatelli and Hurn (1994). This paper criticises the work on M4 by Hall and others, suggesting their measure of financial innovation is not necessary for a cointegrating money demand relationship. In particular, Muscatelli and Hurn draw attention to the argument by Goodhart (1989), that the failure to find a long-run demand for money equation may be due to the particular sample period chosen. If the high inflation rates of the 1970s form the bulk of the sample we are

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[27] Although this is not sufficient, to give M0 the status of an intermediate monetary target.

unlikely to obtain a stable relationship. However, if we include the 1980s in our sample period, where interest rates were high and the inflation rate was low, we are more likely to produce a long-run relationship containing interest rates in levels. This circumvents the problem of attempting to model financial innovation, to produce a significant interest rate effect. Muscatelli and Hurn (1994) go on to find evidence of a cointegrating relationship for M4. We use the 1980s in our sample period and attempt to establish a long-run relationship for M0.

We still have a number of issues to deal with before we can move to empirical estimation of the demand relation. Firstly, we have to choose which variables to represent our scale and opportunity cost measures. We resolve this issue by making a comparative study to that of Breedon and Fisher's (1996) estimation of the demand for M0, without their measure of financial innovation derived from Hall, Hendry and Wilcox (1989). The variables we include are real GDP as the scalar and the Treasury Bill Rate as a short-run nominal interest rate. Breedon and Fisher also include the inflation rate, which is the first difference of the implicit GDP deflator. Additionally, we follow their sample period from 1971Q1 until 1992Q4: which is not dominated by the 1970s.

The second issue of concern is which deterministic components to include in our long-run demand relationship. If we use the Johansen (1988) multivariate technology we can test the joint hypothesis of the number of cointegrating relationships and which deterministic components to include in our model. Following Hansen and Juselius (1995), the specification of the deterministic components in a Vector Error Correction Model is hereafter referred to as model 2, for a constant restricted to the long-run, model 3, for an unrestricted constant in the system and model 4, for a linear time trend restricted to the long-run relationship. This method of testing deterministic components is known as the Pantula principle (see Johansen, 1992). All three models are estimated from the most restrictive (no cointegration and model 2), to the least restrictive (n-1 cointegrating relationships and model 4). The procedure is to progress from the most restrictive to the least and to stop the first time the null hypothesis is not rejected (i.e. the trace statistic is less than the critical values from Osterwauld-Lenum, 1992). Other papers that have attempted to model M0 have not, to the best of our knowledge, explicitly tested for any deterministic components in the demand relation (Hall, Henry and Wilcox, 1989, and Breedon and Fisher, 1996) or have merely assumed one model over another (Westaway and Walton, 1991). The Pantula procedure is a means of testing for the inclusion of deterministic components. This may well be a factor influencing our ability to produce a cointegrating relationship for M0.

3.4.2 Empirical Results

We impose a VAR specification of two lags. Again, we do this to allow direct comparability with Breedon and Fisher's (1996) study of M0 using a cumulative interest rate. The results from residual specification tests suggested that there may be some problems with normality but this

Table 3.1: Trace Test for UK Demand for M0 1971(3) - 1992(4)

H_0:r	Model 2	95% c. v.	Model 3	95% c. v.	Model 4	95% c.v.
$r = 0$	120.84*	53.12	91.99*	47.21	99.02*	62.99
$r \leq 1$	63.74*	34.91	37.47*	29.68	**40.96**	42.44
$r \leq 2$	28.26	19.96	9.79	15.41	11.26	25.32
$r \leq 3$	6.67	9.24	2.52	3.76	3.78	12.25

Notes: Model 2 has a restricted constant, model 3 has an unrestricted constant and model 4 has an unrestricted constant and restricted drift. C.V. is critical values. Pantula procedure begins with model 2 and r=0, then model 3 and r=0, etc until model 4 and $r \leq 3$: we stop the first time the null hypothesis is not rejected. In our case this is model 4 and $r \leq 1$. 95% critical values are from Osterwauld-Lenum (1992).

may be unimportant (see below). The results from the Johansen Trace test are presented in Table 3.1 along with 95% critical values from Osterwauld-Lenum (1992).

At the 95% level we accept one cointegrating vector with model 4: that is to say we include a linear time trend in the cointegrating relationship. The estimated long-run relationship is of the form

$$M_t - P_t = 3.80Y_t - 0.09i_t - 5.75\Delta P_t - 0.03t \qquad (3.6)$$

The corresponding adjustment coefficients α_i feeding into the short run equation of $\Delta(M_t - P_t)$, ΔY_t, Δi_t and $\Delta^2 P_t$ respectively, are t-ratios in parenthesis:

$$\alpha_1 = \mathbf{-0.03\ (-5.84)}\ \alpha_2 = \mathbf{-0.01(-2.08)}\ \alpha_3 = -1.05(-1.80)\ \alpha_4 = 0.01(0.49)$$

The long-run coefficients appear to have the correct signs for a demand relation and the adjustment coefficient is error correcting in the short-run equation for real money. There does appear to be quite a substantive income elasticity of demand. This is compared to some theoretical papers which suggest that it should be around 0.5 or 1.0 (e.g. 0.5 due to Baumol and Tobin's theory of transaction demand for money and 1.0 in Friedman's quantity theory of money, see Cuthbertson and Barlow, 1991). When we restrict the long-run coefficient of income to equal one in our relationship the estimated statistic produced is χ^2 (1) = 5.57 with a probability value of 0.02. This can be accepted at the 1% significance level, but can not be viewed as particularly strong evidence in favour of the hypothesis that trend adjusted velocity is cointegrated with an interest rate measure in semi-logistic form. Breedon and Fisher (1996) also report a large income elasticity when real GDP is used as the scale variable. Testing the hypothesis that the coefficient on income equals one is, however, rejected by Breedon and Fisher's data. This leads them to argue that real GDP may not be the most appropriate scale variable to model M0. They consequently use retail sales, producing a cointegrating relationship that accepts the homogeneity of income restriction.

Nevertheless, we have found evidence consistent with a cointegrating relationship for real M0, which contains income, a nominal interest rate in levels and inflation. As such we have not had to include a cumulative interest rate as a proxy of financial innovation to produce a relationship which cointegrates, acknowledging our caveat regarding the most pertinent scale variable.

3.5 The Empirical Analysis of the Open Economy

Having found evidence of an equilibrium relationship for money demand, we now combine this with a demand for reserves equation from before. We will examine the long-run and short-run nature of money and international reserves and their interaction. Within the context of a congruent VAR, we can consider the sensitivity of changes in reserves to monetary disequilibria: and whether our short-run equation for reserves in Chapter Two is misspecified. We then go on to produce impulse response functions that can be used to assess our long-run restrictions and the dynamic interaction between the variables.

3.5.1 Data Utilised

In this UK study our sample contains quarterly data from 1975Q1 to 1995Q3. The beginning of the sample is constrained by the availability of exchange rate data. All time series are obtained from *International Financial Statistics,* except *M0* from *BoE Financial Statistics,* are converted into natural logarithms, except *i*, and are seasonally adjusted. Our vector of seven variables is $z_t = \{M\text{-}P, Y, i, \Delta P, R, Q\}$ with the variables defined as follows:

M-P = the log of M0/P
Y = the log of real GDP
i = the Treasury Bill Rate
ΔP = the first difference of the GDP deflator
R = the log of reserves (excluding gold and SDRs)
Q = the log of the real effective exchange rate.

We multiply ΔP by four to make it equivalent to a yearly inflation rate.

3.5.2 VAR Specification and Rank Determination

Before we conduct our cointegration analysis we must consider whether the residuals of our VAR are Gaussian. We therefore use multivariate tests for residual autocorrelation, a LM test of up to one and four lags (i.e. LM_1 and LM_4), and Doornik and Hendry's (1994) normality test. The univariate tests that we use are for ARCH and normality. We begin with a VAR specification with four lags. The results are contained in Table 3.2.

The tests for autocorrelation and ARCH are unproblematic but there does appear to be some evidence of non-normality. Johansen (1995) suggests that the multivariate technology used here is based on Gaussian likelihood but the asymptotic properties of the method only require an i.i.d. assumption of the errors. We can interpret the positive ARCH and LM tests results as more important than the normality problem.

Table 3.2: Misspecification Tests on VAR 1975Q1 to 1995Q3

Multivariate Tests			
Residual	LM_1 $\chi^2(36) = 38.22$	(p=0.37)	
Autocorrelation	LM_4 $\chi^2(36) = 34.07$	(p=0.56)	
Normality	$\chi^2(12) = 76.37$	(p=0.00)	
Univariate Tests	**ARCH(4)**	**Normality**	**R^2**
$\Delta(M\text{-}P)$	3.09	1.60	0.68
ΔY	3.29	32.53	0.47
Δi	2.42	9.72	0.45
$\Delta^2 P$	4.05	22.06	0.53
ΔR	2.15	3.93	0.56
ΔQ	0.76	0.61	0.65

Table 3.3: Johansen Trace Test for Reduced Rank.

Eigenvalues	$H_0:r$	*Trace*	*90%*	*95%*
0.46	r=0	149.50*	110.00	114.90
0.33	r≤1	100.36*	82.68	87.31
0.33	r≤2	68.18*	58.96	62.99
0.21	r≤3	**36.95**	39.08	42.44
0.14	r≤4	18.76	22.95	25.32
0.08	r≤5	6.72	10.56	12.25

Notes: 90% and 95% critical values are from Osterwauld-Lenum (1992). Sample period is 1975Q1 to 1995Q3.

Reducing the lag length to three was detrimental to the univariate tests for ARCH. Further reducing the lag length resulted in the failure of the multivariate tests for first order autocorrelation. We therefore retain the VAR with a lag length specification of four. The test for reduced rank is contained within Table 3.3.

At 90% and 95% significance levels the trace test suggests that we have three cointegrating relationships in our system. The plots of the cointegrating vectors, contained in the Appendix 3, also appear to be stationary. This is consistent with Section 3.4.2 with additional relations representing a reserve equation and possibly income. Next we attempt to identify our long run relations, to give them an explicit economic interpretation which are consistent with the statistical properties of our system of equations.

3.5.3 Stochastic Properties of the Data

We firstly attempt to establish what our three stochastic trends or long-run cointegrating relationships represent. We do so by imposing restrictions on the cointegrating vectors of the form

$$\beta = \{H\phi, \psi_1, \psi_2\}$$

where the first vector is restricted (with H the restrictions), and the remaining vectors are unrestricted, taking values ψ.

We are particularly interested in the possibility that

1. velocity is trend stationary,
2. there is a trend stationary demand for money relationship,
3. there exists an excess aggregate demand relationship,
4. which variables enter the long-run demand for reserves relationship.

Table 3.4 contains the results from imposing the long-run restrictions. These tests are distributed as $\chi^2(\upsilon)$, with υ degrees of freedom. Table 3.4 also includes the corresponding probability-values.

Money Relations: Real money adjusted for trend is not stationary by itself (H_1). However trend adjusted money is cointegrated with either inflation or the interest rate. Suggesting a strong long-run relationship between money and either inflation or interest rates.

Income Relations: Trend adjusted real GDP is not stationary in this sample but marginally cointegrates with inflation. The estimated coefficient for ΔP in H_6 is not exactly what is expected with a quasi-Phillips relationship (i.e. an increase in income is unlikely to be associated with a fall in inflation). A much more sensible (in terms of estimated coefficients)

Table 3.4: Long-Run Structural Hypothesis

	M-P	*Y*	*I*	*ΔP*	*R*	*Q*	*Trend*	$\chi^2(\upsilon)$	p-value
Money									
H_1	1	0	0	0	0	0	0.00	13.91(0)	0.00
H_2	1	0	0	-3.90	0	0	-0.00	0.49(2)	0.78
H_3	1	0	0.04	0	0	0	0.00	2.08(2)	0.35
Income									
H_4	0	1	0	0	0	0	-0.01	9.00(3)	0.03
H_5	0	1	0.02	0	0	0	-0.00	11.14(2)	0.00
H_6	0	1	0	2.46	0	0	-0.01	5.14(2)	0.08
H_7	0	1	-0.04	-3.85	0	0	-0.01	0.79(1)	0.37
Velocity									
H_9	1	-1	0	0	0	0	0.03	17.10(3)	0.00
H_{10}	1	-1	0.04	0	0	0	0.01	0.59(2)	0.74
H_{11}	1	-1	0	-5.54	0	0	0.01	2.82(2)	0.24
H_{12}	1	-1	0.05	0.42	0	0	0.01	0.53(1)	0.47
Inflation									
H_{13}	0	0	0	1	0	0	0	9.09(4)	0.06
H_{14}	0	0	0.01	1	0	0	0	6.35(3)	0.10
Reserves									
H_{15}	0	0	0	0	1	18.71	0	14.37(3)	0.00
H_{16}	0	-3.97	0	0	1	1.95	0	9.93(2)	0.01
H_{17}	0	-6.65	0	-35.66	1	-4.92	0	0.61(1)	0.43

Note: A probability value greater than 0.05 suggests that we should accept the null hypothesis that this relationship cointegrates at the 5% level.

aggregate demand equation is produced when interest rates are introduced into the relationship (H_7). Since the relationship is strongly cointegrated, we consider this to be a prime candidate for one of our steady state relationships.

Velocity Relations: Trend adjusted velocity does not appear stationary over our sample. This is consistent with the increases in velocity witnessed in the UK for M0 since the 1950s - see Janssen (1996). Indeed the evidence here is even stronger suggesting that even when velocity is adjusted for trend it is still not stationary.

Trend adjusted velocity does cointegrate with interest rates and inflation, replicating the results in Section 3.3. When both variables are included in H_{12} the estimated long-run coefficients have the correct sign for a demand for money relationship. Indeed on the evidence of this later sample period we obtain a cointegrating demand relationship without a measure of financial innovation. This also accepts the restriction of income homogeneity, consistent with Muscatelli and Hurn (1994) when modelling M4. We consider this relationship (H_{12}) as our second cointegrating vector.

Inflation relation: Inflation by itself is stationary but only marginally. There is more evidence that it cointegrates with the nominal interest rate, suggesting a vaguely stationary real interest rate.

Reserves: We first examine the long-run relationship between reserves and the real exchange rate (H_{15}). With a test statistic of 14.37 and a critical value with three degrees of freedom, we can safely reject the null hypothesis that there is a long-run or cointegrating relationship between reserves and the real exchange rate. This suggests that there may be other important long-run determinants of international reserves. Our preferred specification from our VAR results in Chapter 2 is $H_{16,}$ where reserves form a long-run relationship with income and the real exchange rate. This is not accepted by the data. This may be due to the different deterministic components used here - we do not have a constant restricted to lie in the cointegration space. Alternatively it may be due to the different lag length. Including inflation does however produce a third cointegrating relationship. We accept H_{17} as our third relationship.

3.5.4 Identifying Restrictions on the Long-run Structure

The next stage of our analysis is to combine our three suspected cointegrating relations for excess aggregate demand (H_7), demand for M0 (H_{12}) and demand for reserves (H_{17}) in the entire cointegration space. This will consequently allow us to examine the short-run interaction between the macroeconomic variables and cointegrating relationships. We will impose hypotheses of the form $\beta = \{H_1\phi_1, H_2\phi_2, H_3\phi_3\}$ with H_1, H_2 and H_3 representing our long-run relationships. Here all cointegrating vectors have restrictions imposed. The results are presented in Table 3.5.

These structural relationships are easily accepted with a test statistic of $\chi^2(3) = 2.25$ and a probability value of 0.52. The first relation represents aggregate demand, where income is positively correlated with prices and the interest rate. The trend in this vector represents productivity improvements. The second relation is our demand for money specification.

Table 3.5: Long-Run Cointegrating Vectors

Variable	β_1	β_2	β_3
M-P	0	1.00	0
Y	1.00	-1.00	-5.58 (0.82)
i	-0.03 (0.00)	0.05 (0.01)	0
ΔP	-3.13 (0.35)	0.55 (0.45)	-20.51 (3.52)
R	0	0	1.00
Q	0	0	-1.80 (0.53)
Trend	-0.01 (0.00)	0.01 (0.00)	0

Notes: standard errors are in parenthesis. These restrictions pass the CATS in RATS rank test for identification. β_1 is H_7, β_1 is H_{12} and β_3 is H_{17}.

Long-run homogeneity of income is accepted and interest rates enter the relationship significantly, although with a small estimated coefficient. The Treasury Bill Rate enters in levels and does not have to be transformed to a proxy for financial innovation to produce a cointegrating vector. The included linear time trend in this relationship could be considered as a proxy for technological innovation in the money markets. The third vector is our long-run demand for international reserves equation. R is positively related to Q i.e. reserves increase with an appreciation of the domestic currency. This is a reflection of leaning against the wind. They are also positively related to income and inflation. Reserves are dependent upon the other macroeconomic policy objectives.

We now turn to an examination of the short-run behaviour of our data, where the corresponding adjustment coefficients are contained in Table 3.6. Our interest centres on the importance of β_2 and β_3 (the long-run relationships for money demand and reserves, respectively) in the various short run equations. The first equation in Table 3.6 refers to short run changes in *money balances* $\Delta(M\text{-}P)$. This is significantly error correcting for β_2. This is as we would expected, where there exists a stable long-run money demand relationship, and suggests that M0 in the UK is demand determined. This is consistent with Goodhart (1994) and Juselius (1996).[28] It is also significantly error correcting in β_3. This suggests a

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[28] Juselius (1996) advises there are indeed three characteristics of a monetary aggregate being demand determined. The long-run relationship must have the characteristic of a demand equation, their must be an error correcting adjustment coefficient and the Π

Table 3.6: Short-Run Adjustment Coefficients

Equation	α_1	α_2	α_3
$\Delta(M-P)$	0.07	**-0.09**	**-0.01**
	(1.51)	**(-3.39)**	**(-2.18)**
ΔY	**0.12**	0.01	**-0.01**
	(2.65)	(0.26)	**(-2.08)**
Δi	2.23	1.80	**-1.23**
	(0.28)	(0.40)	**(-2.01)**
$\Delta^2 P$	0.26	**0.42**	0.01
	(1.71)	**(4.77)**	(0.61)
ΔR	1.13	0.62	**-0.11**
	(1.81)	(1.73)	**(-2.27)**
ΔQ	**-0.60**	-0.16	**0.05**
	(-3.25)	(-1.50)	**(3.58)**

Notes: t-statistics are in parenthesis.

significant interaction between the excess demand for reserves and changes in real M0. However the adjustment coefficient is not very large - only around 10% of the adjustment coefficient to excess real money balances. This may indicate that excess holdings of reserves by the central bank are sold for domestic money and interest bearing assets. This will result in a fall in the level of nominal and hence real money balances (i.e. prices are sticky in the short-run) and opens up the possibility that a policy of attempting to influence the exchange rate using reserves, has implications for the conduct of domestic monetary variables. This is contrary to the literature on sterilisation of foreign exchange intervention. We will return to these issues when examining impulse response functions. There is quite a large effect from excess aggregate demand on

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matrix indicates that the real money stock significant adjusts to changes in velocity and not to changes in interest rate. The third criterion is not accepted by my data.

changes in *(M-P)* but this is not significant (i.e. the t-statistic is less than two).

The second equation (i.e. for short run changes in *real income*) is not error correcting to excess aggregate demand - but it is significant. This may reflect the persistence of excess aggregate demand in the economy, consistent with a Keynesian multiplier effect. Disequilibrium in the demand for reserves appears to be significant in this equation and is associated with a small fall in income. Excess demand for money does not appear to have any real income implications in the short-run. Money shocks are not influencing output even in the short run. We contrast this with the short-run equation for inflation, momentarily.

The short-run equation for *interest rates* produces some rather curious results, although they are not strongly significant. All adjustment coefficients are greater than one in absolute value. Their corresponding t-ratios are not significant, apart from the third long-run relationship, which is negative in sign. Again this is possible where excess demand for reserves is associated with an appreciation and the domestic monetary authorities attempt to counter this appreciation by reducing the interest rate. Earlier results indicate that the interest rate was weakly exogenous with respect to the long-run parameters of interest. Hence we should not expect there to be highly significant short-run adjustment coefficients on the equation determining interest rates.

The short-run equation for *inflation* has a very strong monetarist implication. The only adjustment coefficient that is significant in this equation is that representing excess demand for money. The coefficient is large (equalling 0.42) and strongly significant, suggesting that excess money has a strong effects on prices, and within six months. Although this may appear contrary to the idea of sticky prices in the short-run, given that there is not complete adjustment, prices are still sticky to an extent. The adjustment coefficient to excess aggregate demand is only 0.26 and its corresponding t-ratio is less than two, although not by much. Excess aggregate demand may be taking longer to have an impact upon inflation. We generally find evidence that excess money rather than aggregate demand is associated with inflation. This is one of the most contentious propositions in macroeconomics. Our results on M0 are contrary to the evidence on UK M1 found by Hendry (1995) using a VAR with non-stationary data. His paper finds that excess aggregate demand rather than excess money demand is significant in a short-run equation for inflation. Hendry therefore suggests that inflation operates through capacity constraints in the real economy, whereas money aggregates are an "epiphenomenon". Johansen and Juselius (1994) who examine an IS-LM model for the Australian economy support Hendry's evidence. The money stock does not appear to influence inflation but is correlated with excess aggregate demand in the short-run (less than six-months). We purport that M0 has an impact on inflation and within six month. Excess aggregate demand may well have a more delayed impact. Consequently inflation is 'stickier' to capacity constraints than to monetary impulses.

The short-run equation for *reserves* has only one significant adjustment coefficient. This is to disequilibria from long-run reserves. This suggests that reserves are significantly error correcting, vindicating the existence of a long-run equilibrium relationship. Interestingly, excess demand for money does not appear to have a significant impact on short run changes in reserves. Although the coefficient is large, it is not negative and the t-ratio is less than two. This is contrary to the monetary approach to the balance of payments; the work of Frenkel (1983) and Edwards (1984) using partial adjustment models; and that of Elbadawi (1990) and Ford and Huang (1994) using error correction models. The data itself suggests we should not consider that domestic monetary disequilibria of private agents have an impact on reserves, as suggested by these traditional theoretical models.

3.5.5 Impulse Response Functions

Impulse Response Functions (IRFs), or dynamic responses, describe the dynamic properties of the model following certain shocks. They trace out the moving average representation of the system and graphically illustrate how one variable responds over time to a single surprise increase in itself or in any other variable. This approach has been advocated by Sims (1980), who suggested that it was an effective way to examine dynamic interdependencies and of checking Granger non-causality. It can also be used to assess the validity of our long-run restrictions (see Lutkepohl and Reimers, 1992).

Importantly, we must first transform innovations, or shocks, such that they are contemporaneously uncorrelated. With all variables endogenously determined within the system, it is necessary to disentangle the impact of one shock from another. This transformation is achieved by identification. We follow the method of Sims (1980) achieving identification by an orthogonal transformation known as a Choleski decomposition. With such an approach the residuals of the variance covariance matrix from our VECM are transformed to a lower triangular matrix with sufficient restrictions to achieve identification (see Watson, 1994). A particular recursive ordering or Wold causal chain is presumed before the transformation is made. This ordering has implications for the computed results. For example, where the two variables are ***j*** and ***i***: innovations in variable ***i*** do not contemporaneously affect variable ***j***, if ***j*** precedes ***i*** in the particular recursive ordering used; but ***j*** affects ***i*** contemporaneously.

This particular method of identification has the advantage that it reduces the investigators discretion and the scope for data mining. This could be argued to be a disadvantage, since it assumes *a priori* a particular recursive ordering which the data may not actually have. However, alternative methods to achieve identification which do not presume a recursive ordering[29] are equally susceptible to Sims originally criticism of simultaneous equation models, since they also imposed incredible identifying restrictions (see Swanson and Granger, 1997).[30]

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[29] For example, Sims (1986), Bernanke (1986) and Blanchard and Watson (1986).

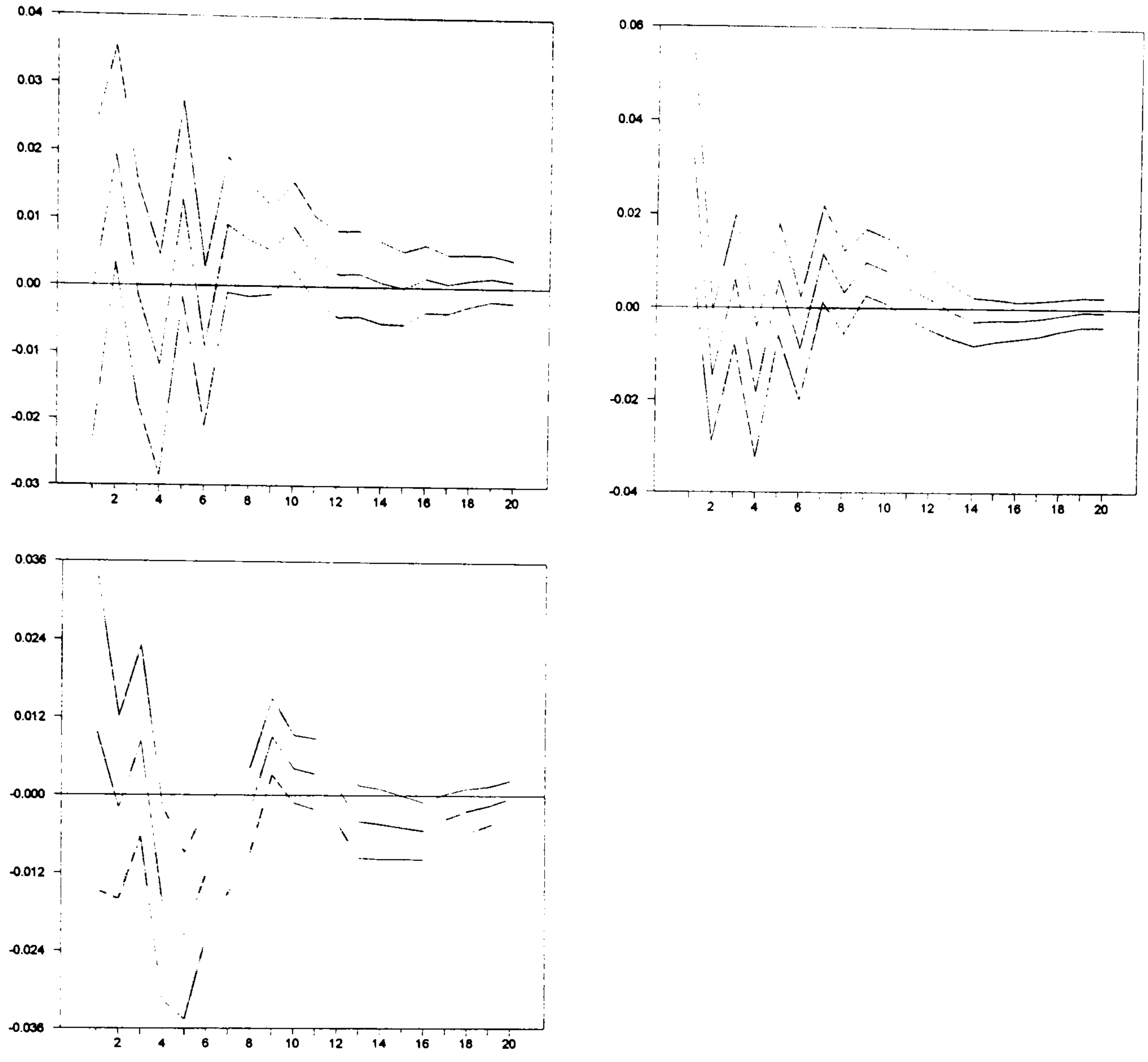
[30] This point is made by N. Gregory Mankiw in reply Leeper, Sims and Zha (1996).

The causal chain that we impose has income and inflation before interest rates. Interest rates are only influencing output and prices with a lag. We additionally assume that the real exchange rate is contemporaneously affecting reserves. Hence reserves are increasing within the quarter to an innovation in the exchange rate. One ordering which captures these relationships would be: income, inflation, exchange rate, money, reserves and interest rate.

Since the model is in VECM form, the corresponding responses will be in first difference (i.e. they are not cumulative responses). We also include estimated significance bands equal to plus and minus two standard errors. Runkle (1987) suggests the point estimates of responses themselves are not particularly meaningful.

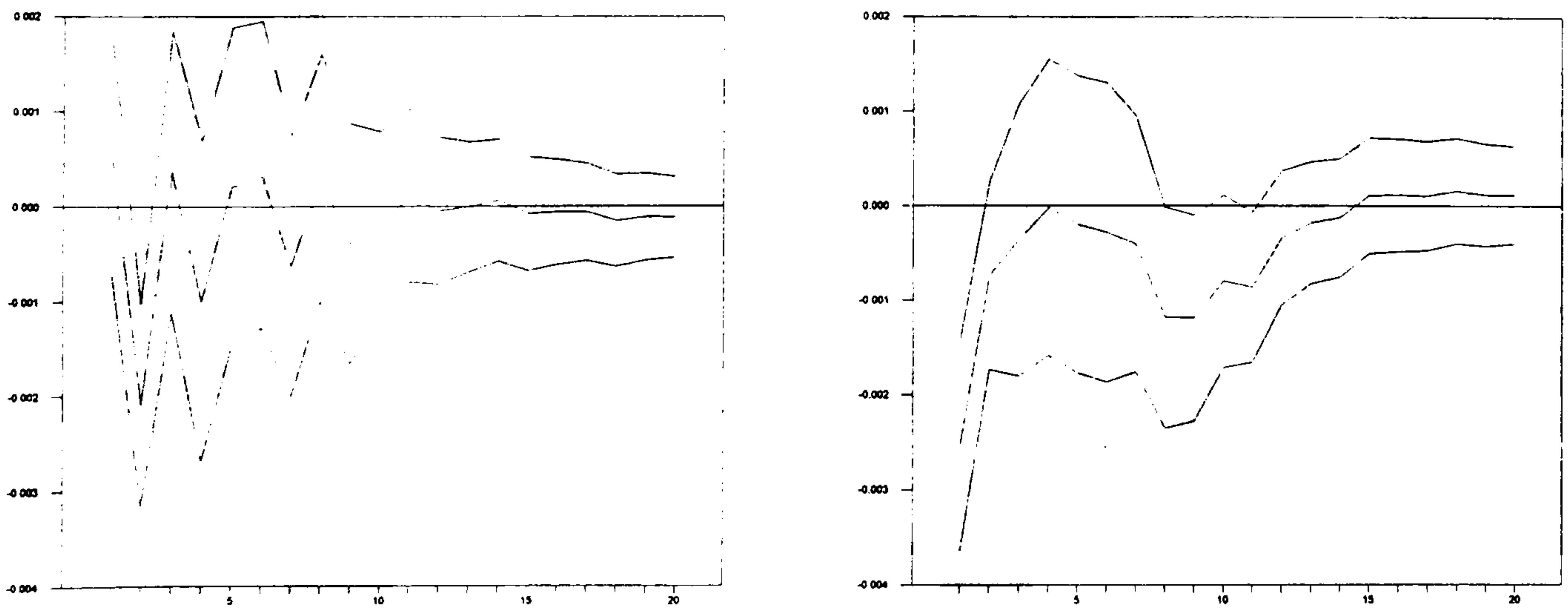
We begin by using IRF to assess the restrictions we have imposed on the cointegrating vectors (as suggested by Lutkepohl and Reimers, 1992). Focusing on the equations for money and international reserves, the computed responses are generally not inconsistent with our long-run restrictions and estimated coefficients. Firstly, reserves permanently increase in response to innovations in the real exchange rate and real income - this response is also immediate and significantly so in the case of the exchange rate (see Figure 3.1). This confirms our opinion that reserves are very responsive to the exchange rate, the Bank of England has generally reacted to an appreciation by leaning against the wind and accumulating foreign exchange reserves. The specification that we have for our reserve relationship is also affirmed by the graph indicating that

Figure 3.1 The Response of Reserves to Innovations in Income, the Exchange Rate and Inflation



Notes: Clockwise from top left the graph includes the response of reserves to innovations in income, the exchange rate and inflation. In addition to the point estimates of the responses we also include significance bands of plus and minus two standard errors.

Figure 3.2 The Response of Money to Innovations in Income and Inflation



Notes: The left-hand graphic is the response of money to an innovation in income.

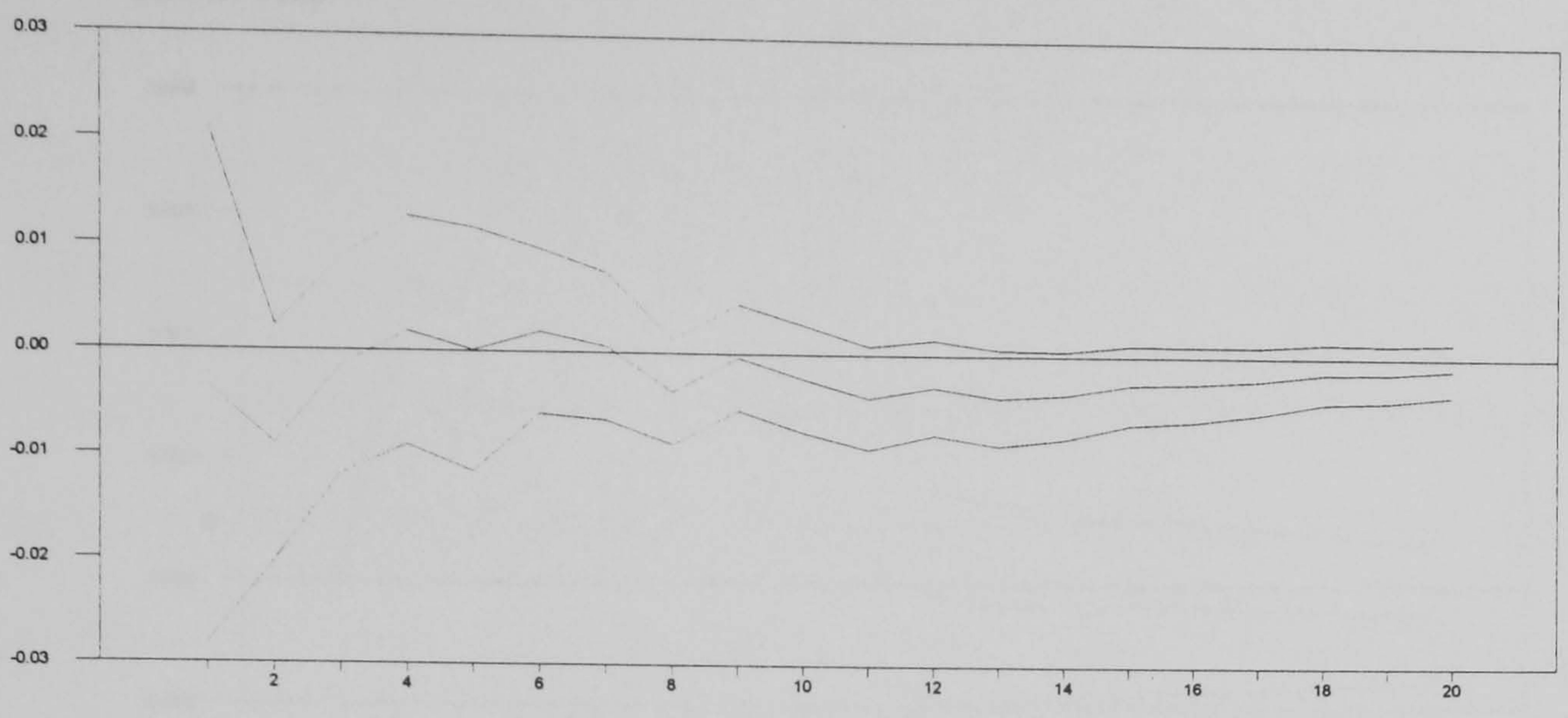
a unit increase in inflation leads to a permanent fall in reserves.

We use our impulse response functions to examine the results for our cointegrating relationship (β_2) for real money balances from Section 3.5.4. Real money balances fall permanently in response to unit innovations in interest rates (Figure 3.5) and inflation (Figure 3.2): both have negative coefficients in vector β_2. Unfortunately, money at first increases and then gradually falls in response to an innovation in income (Figure 3.2). This peculiar relationship between money and income has been found by Lutkepohl and Reimers (1992) when examining German demand for money. However, our long-run restrictions and unrestricted coefficients in these cointegrating vectors are generally consistent with the impulse response functions.

We now go on to consider some other evidence from the IRFs including the responsiveness of reserves to monetary disequilibria and the influence of reserves on domestic monetary aggregates. We also inspect the impact of interest rates on the other variables. This can be considered as a means of checking that we have successfully identified the model.

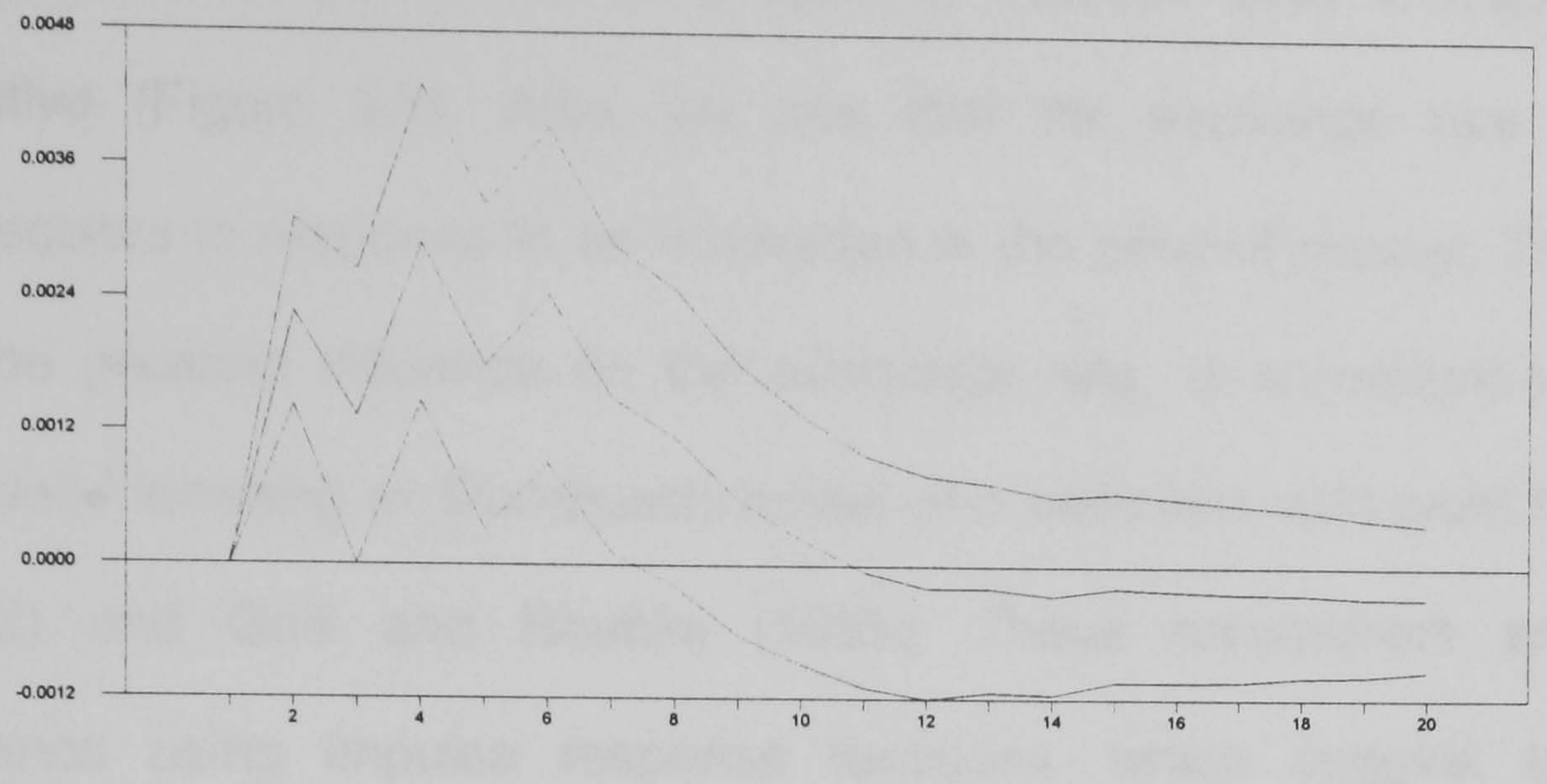
Money market disequilibrium does not appear to have an effect on reserves (see Figure 3.3). Nowhere is the effect significantly different from zero. This leads us to believe that reserves are not greatly influenced by the difference between money supply and money demand. Discounting the relevance of the MABP when modelling reserves. And confirming our short-run equation specification for reserves (i.e. changes in reserves are not dependent upon money market disequilibrium).

Figure 3.3 The Response of Reserves to Money Market Disequilibrium



The impulse response functions furnish us with information regarding the extent to which changes in reserves influence domestic monetary aggregates. This is related to the empirical literature on the degree of sterilisation of government foreign exchange intervention. Obstfeld (1983) suggests that Germany completely sterilised its foreign exchange intervention - as proxied by changes in reserves - for the period between 1975-1981. Kearney and MacDonald (1986) suggest complete sterilisation for the UK in the period 1973-1981. More recent papers offer that sterilisation may in fact not be complete. Neumann and Von Hagen (1992) propose that there is a time dimension to the degree of sterilisation for Germany. They suggest that the Bundesbank has fully sterilised international reserve changes in the short-run, but the degree of sterilisation unwinds over time. Sterilisation is not complete over the long-run (for a similar study of Japan see, Glick and Hutchinson, 1995). This time dimension was not acknowledge by Obstfeld or Kearney and MacDonald: both these papers use a single equation approach which

Figure 3.4 The Response of the Money to an Innovation in International Reserves



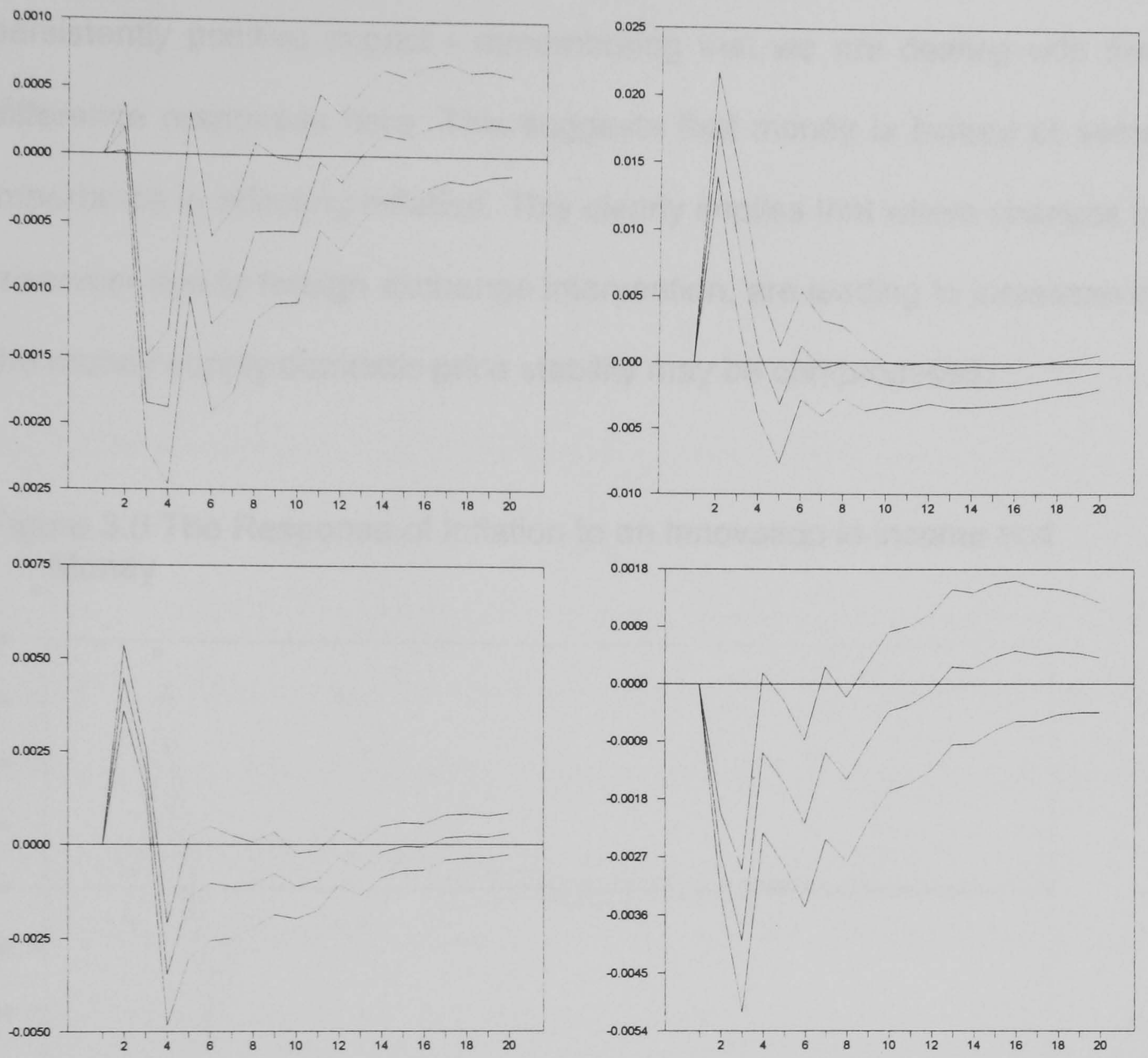
considers the contemporaneous impact of changes in reserves on the money base. Temporal issues related to sterilisation may be important for the UK given the extensive use of forward contracts and credit lines by the Bank of England. These may only influence domestic monetary conditions with a lag. If sterilisation is complete we do not expect the response of M0 to be significantly different from zero.

We find that the initial impact of an increase in reserves on the money base is negligible. Gradually the effect increases significantly and is complete after 10 quarters (see Figure 3.4). Changes in reserves have domestic monetary implications. Real M0's responsiveness to a shock in reserves is second only to the impact of a change in interest rates. This means that although cointegration analysis suggests the existence of a long-run demand equation (i.e. M0 is determined to a certain extent by private agents), M0 is sensitive to changes in policy. Internal price stability may be compromised by the pursuit of external price stability.

We now consider the impact of increases in interest rates. The response of output to an innovation in interest rates is significantly negative (Figure 3.5). Also, we see that the exchange rate sharply appreciates in response to an innovation in the price of money. This is by far the greatest influence on the exchange rate, is consistent with the Mundell-Flemming or Dornbusch model and contrasts with work by Sims (1992) and Grilli and Roubini (1995). These researchers produced evidence using impulse response functions, which suggest that the exchange rate depreciated in response to innovations in interest rates for a number of G-7 countries. This is clearly counter-intuitive, as suggested by Sims and others, and is evidence in favour of our approach of restricting our VAR by imposing long-run relationship before obtaining IRFs.

There is a slight "price puzzle" in that inflation increases in response to an increase in interest rates (Figure 3.5). This has also been widely reported in the VAR literature. Nevertheless, our response is not especially permanent, and is quickly reversed. This makes sense if we consider that prices are set as a mark-up on costs. If interest rates are a business costs, then they can be expected to lead to an initial increase in inflation. This is the rationale we propose here. There is no difficulty with the response of real money balances to an increase in rates. Money quickly falls and there is no "liquidity puzzle". This also verifies Muscatelli and Hurn's (1994) point that real money balances are significantly

Figure 3.5 The Impact of Innovations in Interest Rates on Income, the Exchange Rate, Inflation and Money.

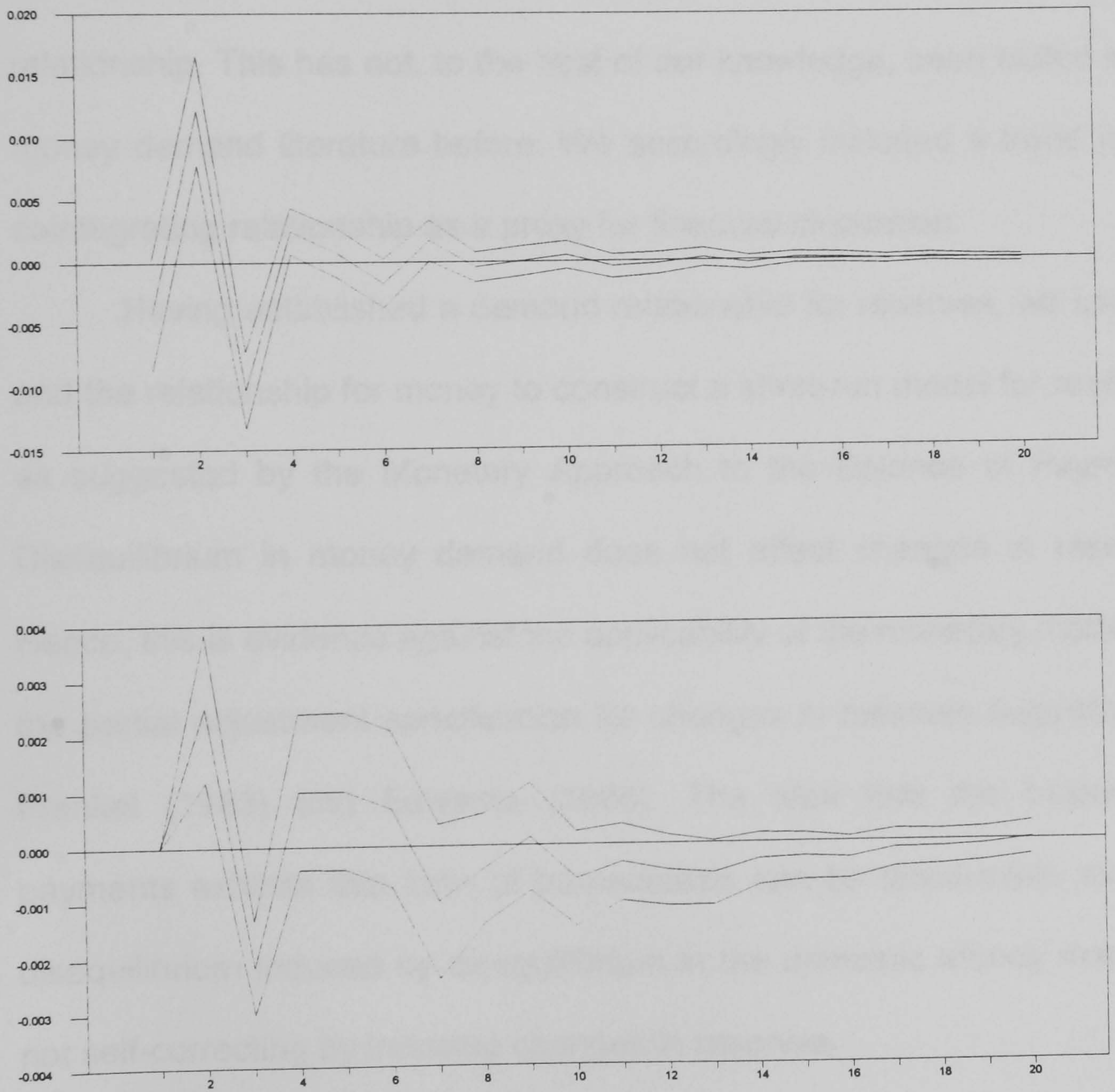


Notes: This graphic represents from top left in clockwise order the response of income, the exchange rate, inflation and real money.

dependent upon interest rates for a sample period which includes the 1980s and 1990s.

Finally, we get some interesting results from our graphs of the response of inflation to a number of shocks. This relates to our earlier discussion on inflation's dependence upon income and money. For example we find in Figure 3.6 that money is having a persistently positive impact upon inflation. Income on the other hand does not seem to have a persistently positive impact - remembering that we are dealing with first difference responses here. This suggests that money is indeed of some importance in affecting inflation. This clearly implies that where changes in reserves, due to foreign exchange intervention, are leading to increases in the money supply domestic price stability may be compromised.

Figure 3.6 The Response of Inflation to an Innovation in Income and Money



3.6. Conclusion

We have established a number of results from this Chapter on the UK monetary sector, international reserves and their interaction. We began by establishing a cointegrating relationship for M0, which only required the inclusion of an interest rate in levels and not a cumulative measure of the price of money as some authors suggest. It is argued that money balances are responsive to interest rates since our sample includes the 1980s when the rate of inflation was low and rate of interest was high. Using Johansen technology and the Pantula Principle we tested explicitly for the inclusion of deterministic components in our cointegrating relationship. This has not, to the best of our knowledge, been tested in the money demand literature before. We accordingly included a trend in this cointegrating relationship as a proxy for financial innovation.

Having established a demand relationship for reserves, we use this and the relationship for money to construct a short-run model for reserves, as suggested by the Monetary Approach to the Balance of Payments. Disequilibrium in money demand does not affect changes in reserves. Hence, this is evidence against the applicability of the monetary model and the partial adjustment specification for changes in reserves suggested by Frenkel (1983) and Edwards (1984). The idea that the balance of payments exhibits this form of homeostasis can be discounted: external disequilibrium induced by disequilibrium in the domestic money market is not self-correcting by inducing changes in reserves.

Our short-run model suggested that money was dependent upon reserves. We then verified this by using impulse response functions. To the extent that foreign exchange intervention effects the money supply we can not be sure of complete sterilisation. Given the impact of reserves on M0, the monetary authorities may be having difficulty balancing (the occasionally conflicting) objectives of external and internal price stability.

With regards to inflation we also produced some interesting results from our impulse response functions. These suggest that inflation increases in response to an increase in money. Income does not appear to be quite as important. This indicates that M0 is an initiator with regards to inflation. This is undoubtedly problematic and not least because economics is not strong on proving or refuting causal relationships. Nevertheless, it may be more useful to policy makers where it can be used as an indicator of future inflation, as suggested by Henry and Pesaran (1993), Astley and Haldane (1995) and Breedon and Fisher (1996). Notoriously, there is a strong positive association between inflation and the interest rate. This is not what one would anticipate where interest rates are the primary mechanism by which the monetary authorities attempt to implement price stability. This "price puzzle" result has also been found by Sims (1992) for major industrialised countries. However this could be a reflection of the reaction of policy makers to an increase in prices, or alternatively interest rates are raising business costs and hence prices. Happily we find that after three quarters interest rates are having a negative effect on inflation.

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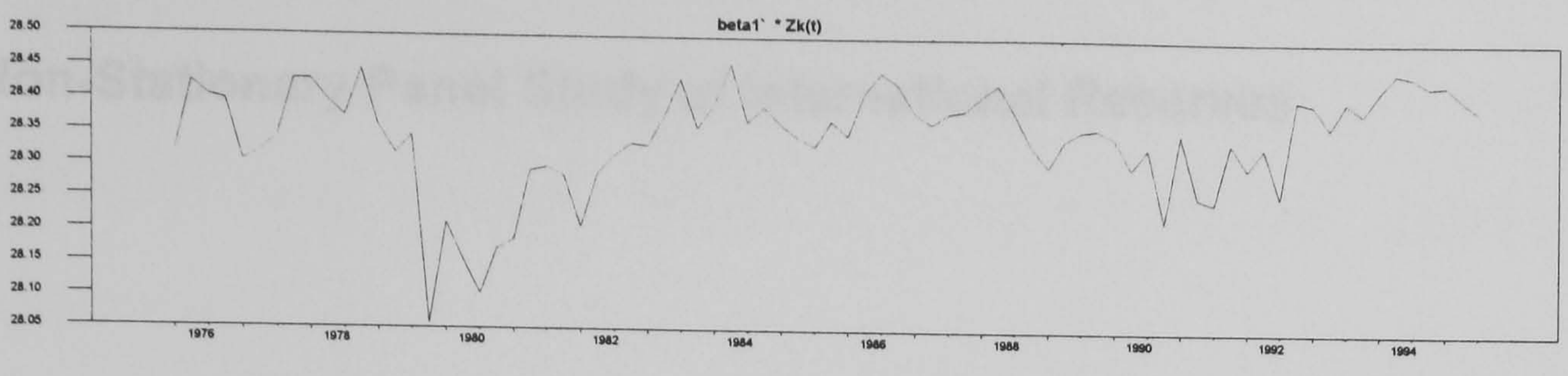
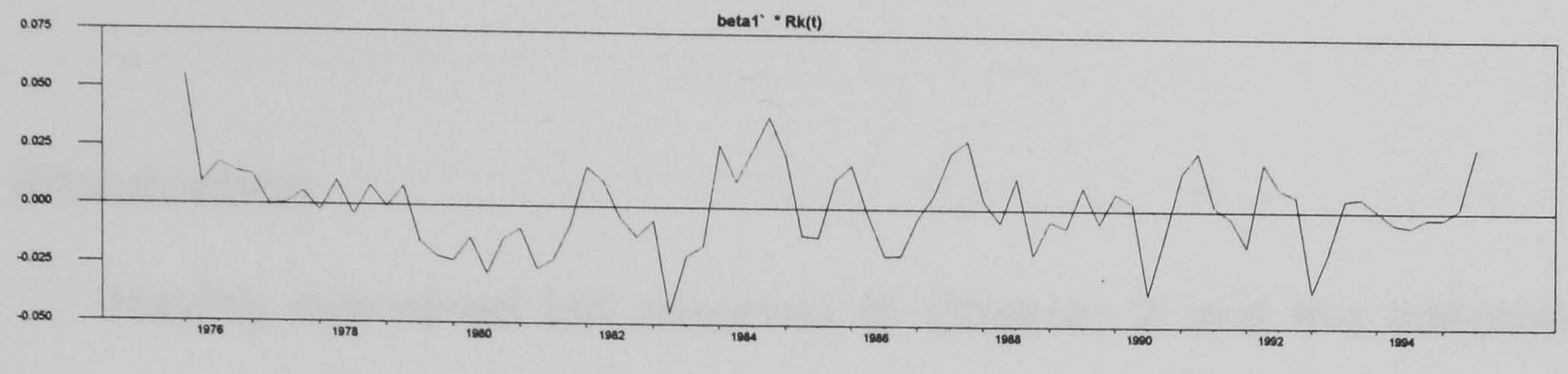
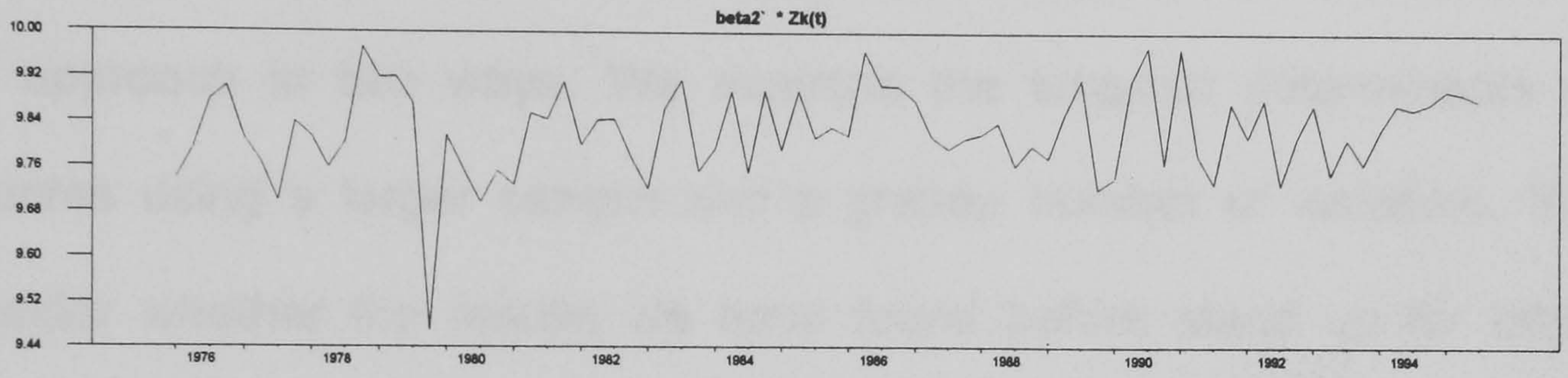
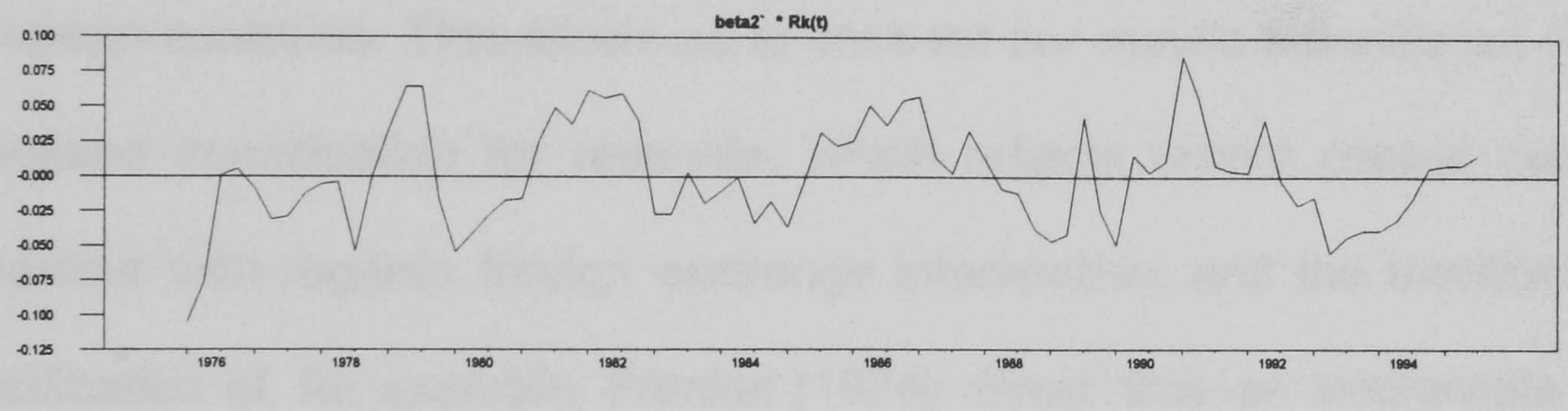
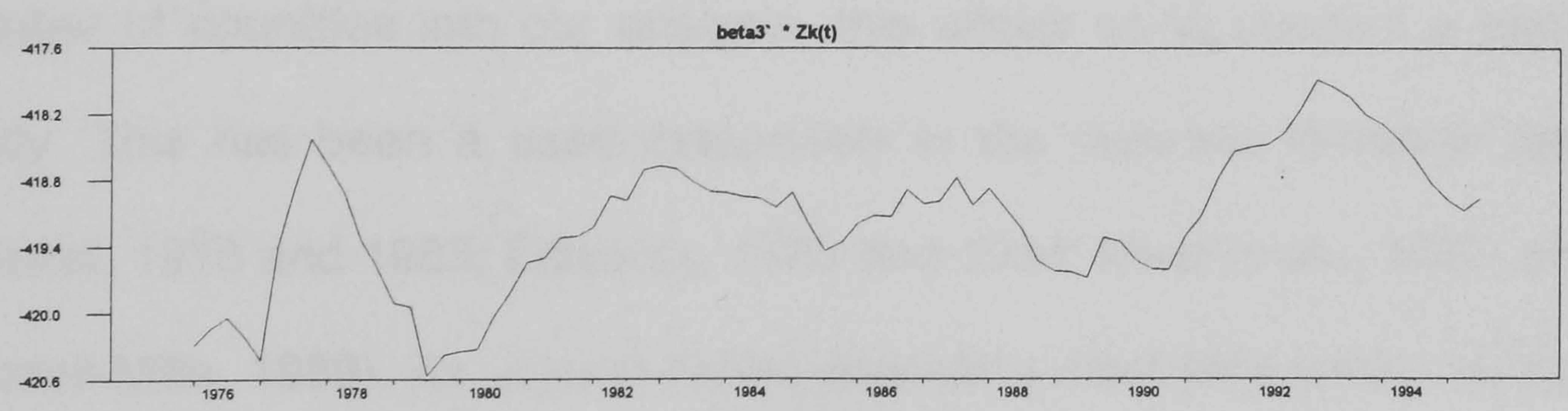
Appendix 3.A: Cointegrating Relationships

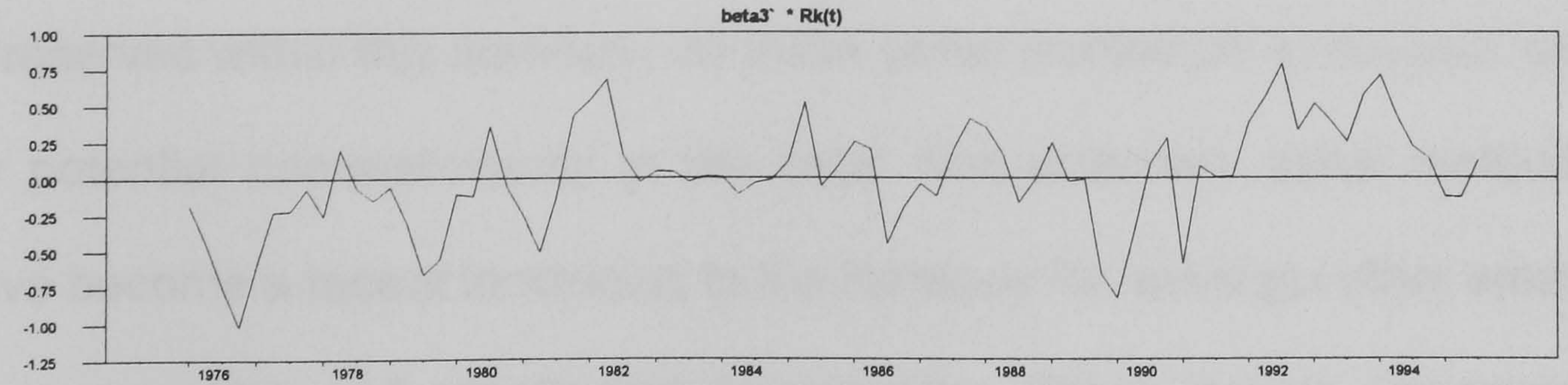












Chapter Four

A Non-Stationary Panel Study of International Reserves

4.1 Introduction

Having examined UK reserves in Chapter 2 and the interaction between reserves and the monetary sector in Chapter 3 we now broaden our approach in two ways. We examine the long-run determinants of reserves using a larger sample and a greater number of variables. We consider whether the results we have found before stand up for other European countries. This allows us to contrast our results between an re-appraised specification for reserves, which reflects recent central bank behaviour with regards foreign exchange intervention, and the traditional specification of, for example, Frenkel (1974). Given that we incorporate a number of countries into our analysis, this allows us to conduct a panel study. This has been a used extensively in the reserves literature (see Frenkel, 1978 and 1983; Edwards, 1983 and 1984; MacDonald, 1987; and Landell-Mills, 1989). As argued before regarding most time series studies of reserves within this literature, all these panel studies fail to account fully for potential non-stationarity in the data. Non-stationary panel methods have become a recent innovation in the literature for, amongst other areas, exchange rates and growth and convergence. These include Canzoneri, Cumby and Diba (1996), Chinn and Johnston (1996), Obstfeld and Taylor (1996), MacDonald (1996) and Taylor (1996). We utilise these methods which allow us to test for panel cointegration, examine the long-run specification and consider short-run adjustment to our equilibrium relationships.

In this Chapter we study nine European countries demand for reserves, for the period since the beginning of the European Monetary System until 1995. The countries included are the UK, Germany, France, Italy, Spain, Netherlands, Austria, Finland and Sweden. We compare the results across our sample of European countries. Firstly, we have the formal exchange rate targeting countries of Germany, France, Netherlands and Italy. These countries participated in the Exchange Rate Mechanism from March 1979. Austria, by pegging its exchange rate to the DM, had a *de facto* commitment to ERM bands. The UK, Spain, Finland and Sweden have had informal exchange rate policies for much of the sample but also participated in the ERM for at least a short period of time. As we see in Figure 4.1 reserve holdings have continued to be important for our nine countries. Indeed for all countries reserve holdings have increased over our sample period.

Our study contrasts with the traditional portfolio-inventory approach by incorporating a role for international reserves in intervening in the foreign exchange market. Using state-of-the-art time series methods, we compare our specification with the approach of traditional studies,

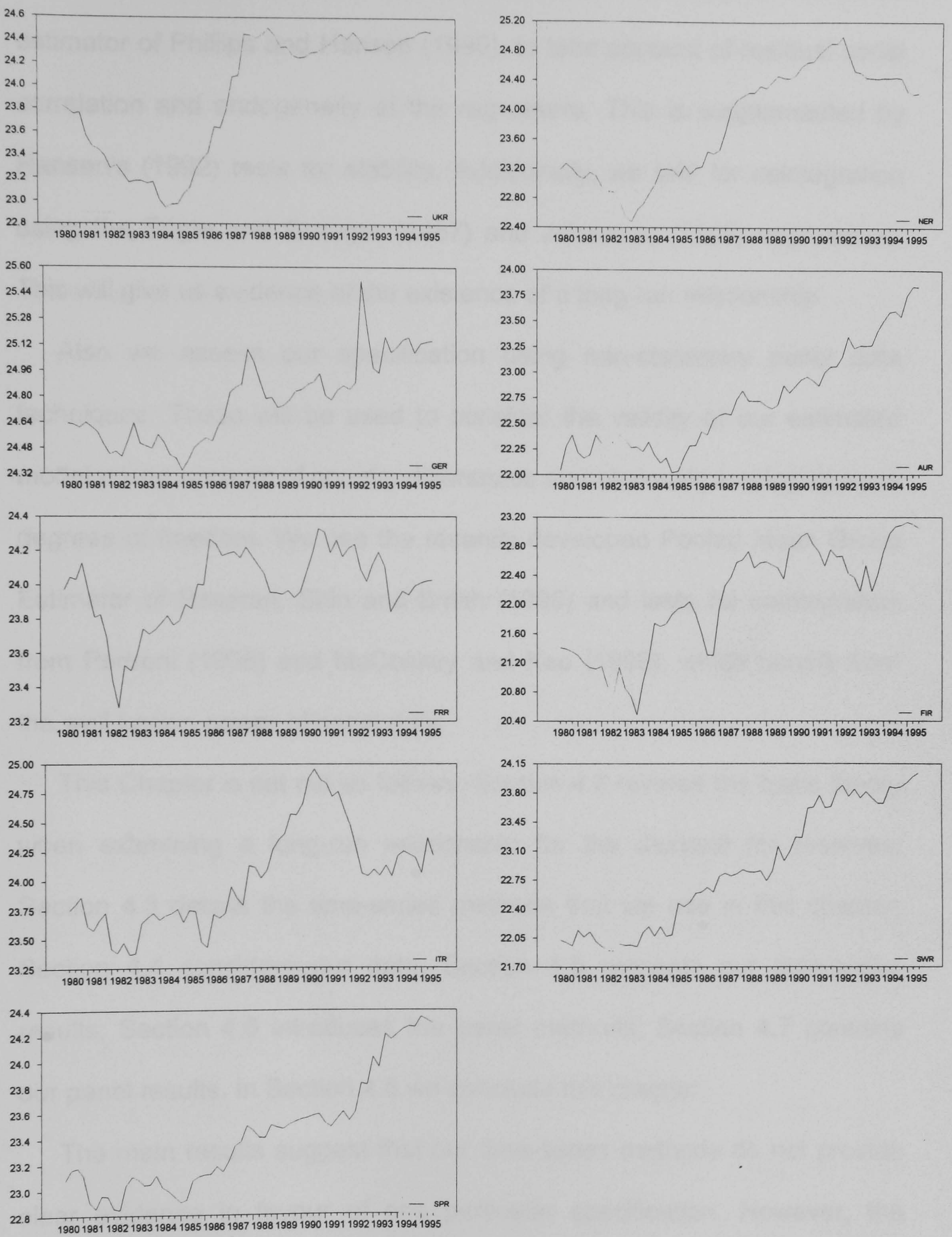


Figure 4.1 European International Reserves in Logs

which do not have an explicit role for exchange rate intervention. In particular we utilise from the time series literature the Fully Modified OLS estimator of Phillips and Hansen (1990), to take account of residual serial correlation and endogeneity of the regressors. This is supplemented by Hansen's (1992) tests for stability. Additionally, we test for cointegration using the Engle and Granger (1987) and Johansen (1988) approaches. This will give us evidence of the existence of a long-run relationship.

Also we assess our specification using non-stationary panel data techniques. These will be used to consider the validity of our estimated model taking account of country differences and obviously employing more degrees of freedom. We use the recently developed Pooled Mean Group Estimator of Pesaran, Shin and Smith (1999) and tests for cointegration from Pedroni (1998) and McCoskey and Kao (1998), which benefit from the well-known power of panel data.

This Chapter is set out as follows: Section 4.2 reviews the basic theory when examining a long-run relationship for the demand for reserves; Section 4.3 details the time-series methods that we use in this chapter; Section 4.4 considers our data; Section 4.5 presents our time-series results; Section 4.6 introduces the panel methods; Section 4.7 contains our panel results. In Section 4.8 we conclude this chapter.

The main results suggest that our time-series methods do not provide clear evidence in favour of one particular specification. However, the evidence is generally in favour of our alternative specification when we utilise the more powerful non-stationary panel methods.

4.2 Theoretical Issues

This Chapter is motivated by Frenkel's (1978) suggestion that the traditional approach to modelling international reserves is mis-specified for the post Bretton Woods period, since it does not take into account the government's attempts to influence the exchange rate through foreign exchange market operations. Traditional studies view reserves as a buffer stock against fluctuations in the Balance of Payments. Any concern with the exchange rate is only indirect. This contrast with the experience of most European countries during the 1980s and early 1990s, where the level of the exchange rate was a direct policy goal and reserves were used to implement this policy.

The traditional specification for reserves is based on the following equation.

$$log(R)_t = constant + \beta_1 log\ (Y)_t + \beta_2 log\ (VOL)_t + \beta_3 log\ (AP)_t \qquad (4.1)$$

$$0<\beta_1<1 \qquad 0<\beta_2<1 \qquad \beta_3>0$$

The level of International Reserves is based on the level of income (*Y*), the volatility of payments represented by the volatility of reserves (*VOL*) and the average propensity to import *(AP),* which is a proxy for the marginal propensity to import.[31] Income is included as a scale variable. Countries will have foreign currency to cover shortages in external payments. Therefore as the economy increases in size, external payments and potential shortfalls will also increase, requiring an increased store of reserves. This can be seen more simply, as a case of larger countries desiring greater levels of reserves. Income will have a positive estimated coefficient and the size, according to the square root law, is less than one: the demand for international liquidity is characterised by economies of scale. We do not necessarily anticipate that the coefficient will be less than one in our study since we regress nominal reserves on real income (following Landell-Mills, 1989).

The inclusion of a measure of volatility was based on the buffer stock view of reserves. As external payments become more volatile, the need for reserves will become greater – resulting in a positive coefficient for β_2. The average propensity to import was also incorporated to represent, under a monetarist perspective, the degree of openness of the economy. The more open a country, the more reserves will be required to counter fluctuation in the external accounts.[32]

The regression that we consider as a possible alternative to the traditional approach has the level of reserves regressed on the real effective exchange rate, the level of real income and also the level of prices.

$$\log(R)_t = constant + \beta_1 \log(Q)_t + \beta_2 \log(Y)_t + \beta_3 \log(P)_t \qquad (4.2)$$

$$\beta_1>0 \qquad \beta_2>0 \qquad \beta_3>0$$

The real effective exchange rate (Q) is included since we presume that the monetary authority attempts to influence the level of Q by intervening in

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[31] All variables are in logs.

[32] The theory itself suggests that the Marginal Propensity to Import should be used but this is generally proxied by Average Propensity.

the foreign exchange market. This is an effective exchange rate and not simply a bilateral rate. Hence, the domestic country resists changes in its bilateral exchange rates with foreign countries, to the extent that it trades with those foreign countries. This model will take account of leaning-against-the-wind of a trend appreciation or depreciation, which policy makers believe is unwelcome and undeserved. The latter will be the case if the change in Q is not based on macroeconomic factors. The estimated coefficient of β_1 in equation *(4.2)* is likely to be positive e.g. as *Q* appreciates the government will increase international reserves holdings, by selling domestic currency in an attempt to curtail the increase in Q.[33]

Income has traditionally been included as a scale variable, and we also give it this interpretation in equation (*4.2*). Changes in reserves are also likely to affect the general price level (*P*). In Chapter Three we found that increases in reserves are likely to be associated with positive monetary impulses. Hence, the monetary authority is unlikely to increase reserves when prices are increasing, if there is an aversion to inflation. This suggests a negative value for the coefficient β_3 in equation (*4.2)*. Price increases are associated with falls in reserves. Alternatively it may be important to include prices to the extent that the monetary authorities will replenish the level of reserves when their value is undermined by inflation. The estimated coefficient in this case is likely to be positive. Hopefully the estimated results in the long-run relationship will resolve this

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[33] The real exchange rate is defined in foreign currency units per unit of domestic currency. This means that an increase in Q is equivalent to an appreciation of the real

question. This can also be considered as an examination of whether we should be using real or nominal reserves. We also take account of the possibility that prices may be I(2),[34] by introducing inflation into our analysis. Given that our sample does not incorporate the inflationary 1970s, it is unlikely that there will be a problem with prices.

We conduct our analysis utilising recently developed time-series methods. Most research in this area does not take account of non-stationary issues in time series estimation. And none that we know of have done so within a non-stationary panel framework. We elucidate further upon these methods in the next section.

4.3 Time Series Econometric Methods

In this section we review the econometric methods that we utilise in this study. These include estimation of our long-run relationship to check that parameter coefficients are as expected. We also consider tests for cointegration, which can be used to assess whether there exists a long-run relationship between our variables or whether the estimated coefficients are spurious. We then go on to conduct similar tests in a non-stationary panel framework.

Given that our data is trending through time we are required to take account of non-stationary data issues. One possible approach is to use the Fully Modified OLS estimator advocated by Phillips and Hansen

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exchange rate.

[34] See Larsson et al. (1998).

(1990). This estimator utilises a non-parametric correction to takes account of the impact of auto-correlation on the residual term. It also corrects for potential endogeneity of the supposedly independent variables. For instance, an endogeneity correction is important where reserves may be having an impact on the exchange rate in specification *(4.2)*.

We also utilise the Hansen (1991) Lc, MeanF and SupF tests for the stability of the linear regression model. These tests have a null hypothesis of stability and have probability values estimated automatically, which improve exposition.[35]

The tests that we use for cointegration are the single equation Engle and Granger (1987) Augmented Dickey Fuller (ADF) test and the multivariate maximum likelihood methods of Johansen (1988). The strategy for choosing lag specification for the ADF test is as suggested by Perron and Ng (1996). Beginning with a generous lag structure, we delete the last lagged first differenced term systematically until it is significant. For this two-step test the critical values are from MacKinnon (1991). The Johansen analysis is conducted by firstly imposing a lag structure of two in our Vector AutoRegression: with only 62 observation and 32 estimated coefficients for a lag length of two, we are concerned with degrees of freedom. Critical values are provided by Osterwauld-Lenum (1992). We find the rank of the Vector Error Correction Model or the number of

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[35] These methods have been examined in greater detail in previous Chapters. See, for example, Section 2.2.

cointegrating vectors using the Johansen Trace Test, since it is relatively robust to the presence of non-normality.

4.4 Data

All data was obtained from the IMF International Financial Statistics CD-ROM. The data is quarterly, transformed into natural logarithms and seasonally adjusted by the IFS (apart from the data for Austria, Finland and Sweden – these were seasonally adjusted using the ESMOOTH procedure in RATS).

The variables include international reserves minus gold (R), following Landell-Mills (1989). Also following the last author, R is in nominal terms. This makes sense since we are particularly interested in fleshing out the relationship between reserves and the real effective exchange rate. The real effective exchange rate (Q) is trade-weighted and is constructed by IFS. Real GDP (Y) and the corresponding deflator (P) are also used. The latter was transformed into a measure of annual inflation by taking first differences and multiplying by four. The variables that we used from the traditional approach included a measure of volatility (VOL) and the Average Propensity to Import (AP). The volatility measure is based on Frenkel (1974). AP was constructed by dividing the level of imports by the level of income for each country and then transforming the variable into logs.

4.5 Time-Series Results

Our objective is to consider the comparative performance of various specifications for reserves over our cross section of nine countries. Previous Chapters, which examined UK reserves, suggested the inclusion of income, the real exchange rate and some measure of prices (see equation *4.2*). We also investigate the performance of a more traditional specification (see equation *4.1*). The methods that we use are FM-OLS estimation, Hansen's (1991) tests for structural breaks, and the Engle and Granger (1987) and Johansen (1988) tests for cointegration.

Firstly, we consider the stationarity properties of the time series using Augmented Dickey Fuller (ADF) tests. The results are contained in Appendix 4.A. These provide consistent evidence that inflation is I(0) (e.g. UK, France, Italy, Spain, Netherlands, Finland and Sweden). Although slightly contentious, this result has been found by others (see Larsson et al., 1998). However there is also evidence that prices (for four countries) are stationary which is difficult to reconcile with previous studies. A visual inspection of the data and examination of correlograms indicates substantial persistence and tend to contradict the ADF results. There is very sporadic evidence of stationarity in the other variables, which contradicts the use of some of our time-series estimators. But given the number of ADF tests that we use here (189 in total) this should not be too surprising.

We now summarise the main results from Tables 4.1 to 4.9 for our nine countries. Firstly we have the UK in Table 4.1. Based on the criteria of significant long run parameters with signs consistent with theory, evidence of stability and cointegration, specification three with income and the real exchange rate is the most appropriate. For Germany, on the basis of significant coefficients consistent with expected signs, single equation cointegration and two of the three stability tests accepting at the 5% significance level, we prefer specification one. For France, The Hansen and Engle-Granger results do not emphasis any specification in particular. The Johansen results indicate either equation one or six: having cointegration, mean reversion and beta coefficients consistent with FM-OLS. Since the signs of the coefficients for equation six are not consistent with the traditional approach, we prefer equation one noting the problem of insignificance of the parameters. This suggests some consistency of results for a sensible long-run reserve relationship for our first three countries; including the exchange rate, income and possibly prices.

Table 4.1: Demand for UK Reserves: Time-Series Results

FM-OLS				***Hansen Stability Tests***			***Cointegration***	
Const	**Regressors**			**Lc**	**MeanF**	**SupF**	**ADF Test**	**Rank of VECM**
-214.10 (55.55)**	6.85 **Q** (1.21)**	7.66 **Y** (1.99)**	0.62 **P** (0.95)	0.39 [0.20]	10.24 [0.01]**	72.65 [0.01]**	-2.89 (0 lags)	2
-223.84 (31.99)**	5.40 **Q** (1.37)**	8.26 **Y** (1.00)**	2.05 **INF** (2.58)	0.18 [0.20]	2.53 [0.20]	6.28 [0.20]	-2.92 (4 lags)	3
-249.54 (29.78)**	7.16 **Q** (1.15)**	8.92 **Y** (0.94)**		0.27 [0.20]	3.63 [0.20]	9.00 [0.20]	-3.10 (0 lags)	2
-234.96 (36.23)**	6.52 **Q** (1.55)**	8.49 **Y** (1.10)**	0.01 **VOL** (0.23)	0.62 [0.06]*	8.64 [0.03]**	19.45 [0.03]**	-2.47 (4 lags)	1
-100.14 (23.62)**	0.64 **VOL** (0.26)**	4.68 **Y** (0.88)**		0.49 [0.10]	5.67 [0.07]*	14.29 [0.07]*	-2.90 (4 lags)	0
-129.86 (11.17)**	0.39 **VOL** (0.12) **	5.30 **Y** (0.39)**	-4.20 **AP** (0.63)**	0.22 [0.20]	2.41 [0.20]	4.40 [0.20]	-2.17 (4 lags)	1

Notes: FM standard errors are in parentheses. The kernel used is a Quadratic Spectral. P-values are in brackets for the Hansen stability test: p-values less than 0.05 reject the null of stability (0.20 means greater than or equal to 0.20). Critical values for the Engle-Granger test are from MacKinnon (1991). The 5% critical value for three regressors is –3.88 and for two regressors –3.44. Large negative values reject the null of no cointegration. The lag specification for the ADF test is as suggested by Perron and Ng (1996). The rank of the VECM is equal to the number of cointegrating vectors in the Johansen Trace Test at the 95% level. More detailed results are in Appendix 4.B.

Table 4.2: Demand for German Reserves: Time-Series Results.

FM-OLS				***Hansen Stability Tests***			***Cointegration***	
Const	**Regressors**			**Lc**	**MeanF**	**SupF**	**ADF Test**	**Rank of VECM**
-24.80 (15.23)	2.69 **Q** (0.53)**	1.31 Y (0.52)**	-2.13 **P** (0.79)**	0.67 [0.04]**	4.97 [0.20]	16.56 [0.08]*	-4.71** (0 lags)	3
13.36 (9.68)	1.99 **Q** (0.54)**	0.08 Y (0.42)	0.51 **INF** (1.43)	0.46 [0.14]	4.66 [0.20]	37.19 [0.01]**	-4.47** (0 lags)	2
12.30 (9.05)	1.94 **Q** (0.50)**	0.13 Y (0.39)		0.41 [0.16]	2.62 [0.20]	9.49 [0.20]	-4.41** (0 lags)	1
16.56 (5.64)**	2.15 **Q** (0.34) **	-0.06 Y (0.25)	-0.04 **VOL** (0.04)	0.37 [0.20]	7.14 [0.07]*	24.43 [0.01]**	-4.31** (0 lags)	2
-9.73 (6.58)	0.09 **VOL** (0.08)	1.22 Y (0.23)**		0.16 [0.20]	6.94 [0.03]**	39.80 [0.01]**	-3.33 (4 lags)	2
-0.57 (3.86)	-0.05 **VOL** (0.04)	0.74 Y (0.15)**	-1.46 **AP** (0.24)**	0.17 [0.20]	4.66 [0.20]	18.03 [0.05]*	-3.75 (4 lags)	1

Table 4.3: Demand for French Reserves: Time-Series Results

FM-OLS				***Hansen Stability Tests***			***Cointegration***	
Const	**Regressors**			**Lc**	**MeanF**	**SupF**	**ADF Test Statistic**	**Rank of VECM**
-46.55 (65.40)	5.63 **Q** (3.36)	1.51 Y (1.96)	0.11 **P** (0.81)	0.21 [0.20]	2.93 [0.20]	16.48 [0.08]*	-3.34 (3 lags)	4
-29.42 (57.16)	5.11 **Q** (3.41)	1.02 Y (1.56)	-5.18 **INF** (2.70)	0.21 [0.20]	2.25 [0.20]	5.65 [0.20]	-2.66 (0 lags)	1
-79.37 (60.88)	4.89 **Q** (4.07)	2.75 Y (1.53)		0.10 [0.20]	0.88 [0.20]	2.16 [0.20]	-3.11 (3 lags)	1
40.46 (56.36)	0.80 **Q** (2.86)	-0.75 Y (1.57)	-0.68 **VOL** (0.25)**	0.26 [0.20]	2.76 [0.20]	6.21 [0.20]	-2.98 (0 lags)	0
53.29 (29.36)	-0.70 **VOL** (0.23)**	-1.06 Y (1.01)		0.19 [0.20]	2.04 [0.20]	5.60 [0.20]	-2.97 (0 lags)	1
77.87 (22.33)**	-0.72 **VOL** (0.16)**	-2.05 Y (0.80)**	-1.51 **AP** (0.48)**	0.34 [0.20]	2.44 [0.20]	5.27 [0.20]	-3.46 (0 lags)	1

Table 4.4: Demand for Italian Reserves: Time-Series Results

FM-OLS				***Hansen Stability Tests***			***Cointegration***	
Const	**Regressors**			**Lc**	**MeanF**	**SupF**	**ADF Test Statistic**	**Rank of VECM**
-321.21 (51.29)**	1.75 **Q** (0.41)**	9.70 Y (1.48)**	-1.63 **P** (0.41)**	0.64 [0.05]*	27.28 [0.01]**	44.88 [0.01]**	-3.91** (0 lags)	1
-229.06 (17.98)**	2.07 **Q** (0.31)**	7.01 Y (0.51)**	7.84 **INF** (1.08)**	0.59 [0.07]*	16.64 [0.01]**	26.53 [0.01]**	-4.68** (0 lags)	1
-134.23 (20.23)**	1.90 **Q** (0.60)**	4.31 Y (0.50)**		0.61 [0.05]*	16.54 [0.01]**	39.75 [0.01]**	-3.00 (0 lags)	0
-139.26 (20.25)**	1.59 **Q** (0.59)**	4.48 Y (0.57)**	-0.31 **VOL** (0.24)	0.54 [0.09]*	16.65 [0.01]**	26.71 [0.01]**	-3.17 (0 lags)	0
-132.27 (35.20)**	-0.75 **VOL** (0.41)	4.45 Y (1.00)**		0.11 [0.20]	1.31 [0.20]	2.63 [0.20]	-2.34 (0 lags)	0
-133.67 (44.56)**	-0.71 **VOL** (0.42)	4.51 Y (1.34)**	0.15 **AP** (0.90)	0.18 [0.20]	2.04 [0.20]	6.17 [0.20]	-2.30 (0 lags)	0

For Italy, the estimated coefficients look consistent with the UK results with some additional importance for prices or inflation. These variables are significant and their inclusion suggests the existence of a long-run relationship. Evidence of cointegration at the 95% level is restricted to equations one and two in Table 4.4. Also, given that the Engle-Granger and Johansen approaches reinforce one another and that the FM coefficients are all significant we prefer equations one and two. In the case of Spain (Table 4.5), given the significance of all estimated coefficients and evidence of cointegration using both tests, we prefer an equation that incorporates *Q*, *Y* and volatility. That said, the more traditional approach does not do too badly either. Nevertheless, the consistent evidence produced regarding the reasonable sign and significance of *Q*, give results consistent with the previous countries. For the Netherlands, the question becomes weighing up the preference for coefficients with expected sign and significance, with evidence of stability: and hence between the newer and older specifications. On the basis of estimated coefficients we prefer the newer specification of equation four, which implies a role for *Y*, *Q* and *VOL*.

So far the evidence across countries suggests a specification which includes income – consistently positive and significant - and the real exchange rate – consistently positive for all countries and significant, apart from France. They have often included a measure of prices or volatility. This consistency of results can be expected from the first six countries that have participated in ERM for the greatest period of time.

Table 4.5: Demand for Spanish Reserves: Time-Series Results.

FM-OLS				***Hansen Stability Tests***			***Cointegration***	
Const	**Regressors**			**Lc**	**MeanF**	**SupF**	**ADF Test Statistic**	**Rank of VECM**
-141.21 (31.70)**	4.47 **Q** (1.10)**	5.35 **Y** (1.18)**	-0.12 **P** (0.43)	0.95 [0.01]**	9.64 [0.02]**	13.22 [0.20]	-2.73 (4 lags)	2
-133.41 (16.19)**	4.84 **Q** (1.07)**	5.00 **Y** (0.46)**	1.13 **INF** (1.41)	0.76 [0.03]**	5.68 [0.16]	8.19 [0.20]	-3.14 (0 lags)	3
-132.80 (15.48)**	4.66 **Q** (1.02)**	5.00 **Y** (0.44)**		0.80 [0.02]**	5.92 [0.06]*	8.67 [0.20]	-2.71 (4 lags)	2
-99.67 (7.35)**	2.84 **Q** (0.45)**	4.11 **Y** (0.21)**	0.35 **VOL** (0.05)**	0.49 [0.12]	8.08 [0.04]**	13.46 [0.20]	-4.72** (0 lags)	3
-60.17 (7.72)**	0.47 **VOL** (0.10)**	3.14 **Y** (0.28)**		0.36 [0.20]	4.88 [0.20]	8.23 [0.20]	-2.51 (4 lags)	1
-53.47 (6.36)**	0.45 **VOL** (0.08)**	2.96 **Y** (0.23)**	0.56 **AP** (0.23)**	0.61 [0.06]*	6.97 [0.08]*	14.93 [0.14]	-3.68 (0 lags)	2

Table 4.6: Demand for Dutch Reserves: Time-Series Results

FM-OLS				***Hansen Stability Tests***			***Cointegration***	
Const	**Regressors**			**Lc**	**MeanF**	**SupF**	**ADF Test Statistic**	**Rank Test**
-160.55 (32.03)**	1.55 **Q** (0.69)**	5.64 **Y** (1.08)**	-2.57 **P** (1.73)	0.25 [0.20]	11.24 [0.01]**	30.29 [0.01]**	-4.06** (4 lags)	2
-111.96 (12.03)**	1.85 **Q** (0.56)**	4.06 **Y** (0.42)**	-0.65 **INF** (0.73)	2.50 [0.01]**	40.17 [0.01]**	121.21 [0.01]**	-2.17 (4 lags)	2
-115.87 (11.29)**	2.02 **Q** (0.57)**	4.15 **Y** (0.39)**		0.32 [0.20]	10.75 [0.01]**	29.12 [0.01]**	-2.33 (4 lags)	1
-115.95 (8.97)**	1.74 **Q** (0.45)**	4.18 **Y** (0.31)**	-0.18 **VOL** (0.09)**	0.35 [0.20]	7.90 [0.05]*	17.86 [0.05]*	-4.59** (4 lags)	1
-126.73 (19.72)**	-0.45 **VOL** (0.20)**	4.75 **Y** (0.63)**		0.10 [0.20]	1.90 [0.20]	3.98 [0.20]	-4.20** (4 lags)	0
-126.27 (22.32)**	-0.50 **VOL** (0.16)**	4.72 **Y** (0.75)**	0.08 **AP** (0.82)	0.18 [0.20]	3.93 [0.20]	7.88 [0.20]	-4.35** (4 lags)	1

Table 4.7: Austrian Reserve Demand: Time-Series Results

FM-OLS				***Hansen Stability Tests***			***Cointegration***	
Const	**Regressors**			**Lc**	**MeanF**	**SupF**	**ADF Test Statistic**	**Rank of VECM**
-76.75 (110.71)	-0.38 **Q** (3.07)	3.79 **Y** (3.76)	0.36 **P** (2.03)	0.29 [0.20]	2.59 [0.20]	7.28 [0.20]	-2.43 (3 lags)	2
-116.47 (59.47)**	0.57 **Q** (2.99)	5.10 **Y** (1.73)**	0.96 **INF** (3.07)	0.47 [0.14]	4.30 [0.20]	12.61 [0.20]	-2.44 (5 lags)	2
-110.97 (54.34)**	0.32 **Q** (2.79)	4.94 **Y** (1.57)**		0.16 [0.20]	1.89 [0.20]	7.37 [0.20]	-2.38 (5 lags)	1
-196.92 (41.83)**	4.48 **Q** (1.95)**	7.50 **Y** (1.25)**	0.64 **VOL** (0.18)**	0.46 [0.14]	6.17 [0.12]	13.53 [0.20]	-3.11 (0 lags)	1
-102.56 (12.09)**	0.40 **VOL** (0.17)**	4.72 **Y** (0.46)**		0.26 [0.20]	2.75 [0.20]	9.08 [0.20]	-2.65 (0 lags)	1
-96.42 (9.23)**	0.36 **VOL** (0.13)**	4.42 **Y** (0.36)**	-1.54 **AP** (0.50)**	0.49 [0.12]	5.37 [0.19]	9.64 [0.20]	-2.20 (4 lags)	1

Table 4.8: Demand for Finnish Reserves: Time-Series Results

FM-OLS				*Hansen Stability Tests*			*Cointegration*	
Const	**Regressors**			**Lc**	**MeanF**	**SupF**	**ADF Test Statistic**	**Rank of VECM**
-137.32 (88.70)	-1.37 **Q** (1.05)	6.50 **Y** (3.61)	-0.42 **P** (1.62)	0.43 [0.16]	4.42 [0.20]	8.96 [0.20]	-3.98** (2 lags)	3
-123.30 (29.66)**	-1.38 **Q** (0.66)**	5.97 **Y** (1.14)**	0.14 **INF** (1.94)	0.23 [0.20]	2.96 [0.20]	6.68 [0.20]	-3.93** (2 lags)	1
-121.48 (28.69)**	-1.38 **Q** (0.63)**	5.88 **Y** (1.09)**		0.10 [0.20]	1.12 [0.20]	4.80 [0.20]	-3.93** (2 lags)	0
-76.27 (41.98)	-0.53 **Q** (0.85)	4.00 **Y** (1.70)**	0.50 **VOL** (0.43)	0.39 [0.20]	3.36 [0.20]	17.04 [0.07]*	-4.13** (2 lags)	1
-86.40 (42.45)**	0.58 **VOL** (0.32)	4.31 **Y** (1.65)**		0.19 [0.20]	2.30 [0.20]	7.70 [0.20]	-4.08** (2 lags)	0
-83.80 (36.67)**	0.78 **VOL** (0.31)**	4.27 **Y** (1.44)**	0.74 **AP** (0.80)	0.23 [0.20]	2.53 [0.20]	8.31 [0.20]	-3.15 (0 lags)	2

Table 4.9: Demand for Swedish Reserves: Time-Series Results

FM-OLS				*Hansen Stability Tests*			*Cointegration*	
Const	**Regressors**			**Lc**	**MeanF**	**SupF**	**ADF Test Statistic**	**Rank of VECM**
150.93 (38.38)**	1.09 **Q** (0.40)**	-4.98 **Y** (1.48)**	4.38 **P** (0.49)**	1.13 [0.01]**	10.29 [0.01]**	18.65 [0.04]	-2.25 (5 lags)	4
-165.23 (26.17)**	-1.99 **Q** (0.61)**	7.45 **Y** (0.97)**	-0.25 **INF** (0.71)	1.08 [0.01]**	8.65 [0.03]**	20.40 [0.02]**	-2.03 (4 lags)	4
-173.24 (26.24)**	-1.99 **Q** (0.62)**	7.75 **Y** (0.97)**		0.65 [0.04]**	7.12 [0.03]**	19.52 [0.01]**	-1.99 (0 lags)	3
-88.80 (19.54)**	0.23 **Q** (0.39)	4.25 **Y** (0.74)**	0.75**VOL** (0.10)**	1.17 [0.01]**	8.95 [0.03]**	12.44 [0.20]	-1.78 (1 lags)	1
-89.89 (19.43)**	0.72**VOL** (0.09)**	4.33 **Y** (0.73)**		0.49 [0.10]*	5.53 [0.08]*	6.87 [0.20]	-1.84 (1 lags)	0
-90.83 (20.73)**	0.70**VOL** (0.10)**	4.34 **Y** (0.77)**	-0.36 **AP** (0.44)	0.81 [0.02]**	5.97 [0.14]	8.58 [0.20]	-1.82 (1 lags)	0

In the case of Austria (Table 4.7), the best combination of estimated coefficients and evidence of Johansen cointegration is for a specification including volatility and income in the long-run results. However, while the Johansen results give good estimated betas or long-run coefficients (see Appendix 4.B.5) there is no evidence of mean reversion. This is only found for the alternative specifications one to three. Given that they produce not entirely plausible estimated beta coefficients this is not very strong evidence in favour of these specifications.

The preferred specification is based mostly on the estimated coefficients, suggesting equation four or five. *Q* is significant in the former, with a sign consistent with other countries. Austria does not display much consistency across specifications, with respect to *Q*'s estimated coefficients.

The results for Finland do not appear to produce a consistent tale. But given the evidence of stability, cointegration, correctly signed coefficients and only one coefficient insignificant, the results for Finland would appear to prefer the traditional specification.

Table 4.9 indicates estimated coefficients are consistently positive and significant for income and volatility in the case of Sweden. The only difficulty with the income coefficient is when prices are included in the regression: *Y* becomes significant and negative. There is little apparent consistency with the real exchange rate. The fourth specification would appear to suggest that Sweden has often been leaning-against-the-wind of an appreciating exchange rate using reserves. The presence of *Q* seems to ensure the existence of a cointegrating relationship. And on that basis we could rule out equations five and six. Given volatility is consistently significant we prefer specification four.

In summary we have found consistent evidence of the importance of income. We also found evidence of a positive and significant role for the real exchange rate for the UK, Germany, Italy, Spain, Netherlands and marginally in the case of France, Austria and Sweden. For each country we discovered some variation in other variables contained in a long-run relationship. Some included prices or inflation (Germany, France and Italy) others include the more traditional variable, volatility (e.g. Spain, Austria and Sweden). This would appear to suggest the existence of a preferred specification for reserves. We proceed to conduct a panel study with a view to shedding more light on the most appropriate specification for our sample.

4.6 A Panel Study of International Reserves

In the exchange rate literature, sensible long relationships can be found by increasing the span of the data, from the post Bretton Woods period, to a period of one hundred years of annual data (for a survey see Froot and Rogoff, 1995 and MacDonald, 1995). However a significant problem with increasing the data span on a historical basis is that we introduce the possibility of regime shifts. Another approach to increasing the span of the data involves utilising panel sets (MacDonald, 1988, 1996; Chinn and Johnston, 1996; Frankel and Rose, 1996). Examining purchasing power parity and the time series properties of the real exchange rate, these researchers have found sensible long-run cointegrating relationships, with reasonably signed and sized estimated coefficients and evidence of mean reversion.

Although we have found evidence of sensible reserve relationships using time-series estimation, these will certainly benefit from corroboration using more robust methods. For instance, we have a small span of quarterly data from 1980Q2 to 1995Q3, giving 62 observations, while Toda (1994) suggests that Johansen (1988) tests benefit from a sample of 300. In the context of our study, panel tests have a clear statistical advantage because they have greater power to reject the null of no cointegration when it is in fact false. It should be noted that extending the span by including a greater number of countries may introduce cross sectional heterogeneity. The methods we introduce below take account of individual differences across our sample of countries. In this Section we firstly introduce the Pedroni and McCoskey and Kao panel cointegration tests and then consider the Pesaran, Shin and Smith (1999) Pooled Mean Group Estimator.

4.6.1 Panel Cointegration Tests

4.6.1.1 Pedroni's Tests for Panel Cointegration

Pedroni (1995, 1997 and 1998) has developed a suite of seven tests based on the null of no panel cointegration. These tests allow for individual specific fixed effects and time trends. The associated cointegrating vectors and the dynamics of the underlying error process are permitted to display considerable heterogeneity over cross sections. All seven of the tests that we use here are based on the multivariate regression *(4.3)*.

$$y_{it} = \alpha_i + \delta_i t + \beta_{1i} X_{1it} + \beta_{2i} X_{2it} + \ldots + \beta_{Mi} X_{Mit} + e_{it} \qquad (4.3)$$
$$for \qquad t = 1,\ldots,T; \quad i = 1,\ldots,N; \quad m = 1,\ldots,M$$

T refers to the number of observations over time, N refers to the number of cross sections in the panel, and M refers to the number of regressors. The slope coefficients are allowed to vary by cross section i.e. this is a heterogeneous slopes model. The parameter α_i is the member specific intercept, this also varies across cross sections - this is Pedroni's 'fixed effect' parameter. We may also wish to include deterministic time trends $\delta_i t$, which are specific to each individual N.

It may also be of interested to include a common set of time dummies. These are intended to capture shocks that are shared across the different members of the panel i.e. to remove a potential common factor problem. The consequence of this adjustment is to make the disturbances independent across individual members. We can achieve this for each variable by demeaning the data over the cross section dimension, or, in other words, subtracting out cross section averages $\bar{y}_t = N^{-1}\sum_{i=1}^{N} y_{it}$. The unit root asymptotics are unaffected by removing cross section means. However, it should be noted that including time dummies is not equivalent to demeaning the data across cross sections when we have heterogeneous coefficients. This could introduce data dependencies into the estimated residuals so that asymptotic distributions are no longer nuisance parameter free. Estimating time dummies may be preferable in small samples. We remove time averages and note this caveat.

The first four Pedroni tests are based on the traditional panel within estimator (see Hsiao, 1986) and are known as the *Panel Estimators*. Pedroni's *Panel* statistics are produced by summing both the numerator and denominator terms separately for each N (see Table 4.10 for the construction of these test statistics). This effectively pools the autoregressive coefficients across different members of the unit root process of the estimated residuals. The null hypothesis for these test is $H_o:\rho_i = 1$ for all i against the alternative H_a: $\rho_i = \rho < 1$ for all i. Therefore this test presumes a common value for ρ under the alternative. Amongst the *Panel* tests are a variance ratio test (v-statistic), Phillips and Perron (1988) ρ-statistic and t-statistic (non-parametric), and an ADF t-statistic (parametric).

An additional three statistics are based on pooling along the between dimension and these are known as *Group Mean Panel Estimators*. To construct these statistics we divide the numerator by the denominator prior to summing over N (again see Table 4.10). Hence these statistics can be seen as simple averages of the estimated coefficients for each of the N members. The null hypothesis is of the form H_o: $\rho_i = 1$ for all i where the alternative has H_a: $\rho_i < 1$. In contrast to the *Panel Estimators*, this method does not presume that there is a common value for rho under

Table 4.10: Pedroni's (1998) Panel Cointegration Statistics

1. Panel v-Statistic $T^2 N^{3/2} Z_{\hat{v}_{NT}} \equiv T^2 N^{3/2} \left(\sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^2 \right)^{-1}$

2. Panel ρ-Statistic $T\sqrt{N} Z_{\hat{\rho}_{NT}-1} \equiv T\sqrt{N} \left(\sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^2 \right)^{-1} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \left(\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i \right)$

3. Panel PP t-Statistic $Z_{t_{NT}} \equiv \left(\tilde{\sigma}_{NT}^2 \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^2 \right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \left(\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i \right)$

   (nonparameteric)

4. Panel t-Statistic $Z_{t_{NT}}^* \equiv \left(\tilde{s}_{NT}^{*2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{*2} \right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^{*} \Delta \hat{e}_{it}^{*}$

   (parameteric)

5. Group ρ-Statistic $TN^{-1/2} \tilde{Z}_{\hat{\rho}_{NT}-1} \equiv TN^{-1/2} \sum_{i=1}^{N} \left(\sum_{t=1}^{T} \hat{e}_{it-1}^2 \right)^{-1} \sum_{t=1}^{T} \left(\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i \right)$

6. Group PP t-Statistic $N^{-1/2} \tilde{Z}_{t_{NT}} \equiv N^{-1/2} \sum \left(\hat{\sigma}_i^2 \sum_{t=1}^{T} \hat{e}_{it-1}^2 \right)^{-1/2} \sum_{t=1}^{T} \left(\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i \right)$

   (nonparameteric)

7. Group t-Statistic $N^{-1/2} \tilde{Z}_{t_{NT}}^* \equiv N^{-1/2} \sum_{i=1}^{N} \left(\sum_{t=1}^{T} \hat{s}_i^{*2} \hat{e}_{it-1}^{*2} \right)^{-1/2} \sum \hat{e}_{it-1}^{*} \Delta \hat{e}_{it}^{*}$

   (parameteric)

where $\hat{\lambda}_i = \frac{1}{T} \sum_{s=1}^{k_i} \left(1 - \frac{s}{k_i + 1} \right) \sum_{t=s+1}^{T} \hat{\mu}_{it} \hat{\mu}_{it-s}$, $\hat{s}_i^2 \equiv \frac{1}{T} \sum_{t=1}^{T} \hat{\mu}_{it}^2$, $\hat{\sigma}_i^2 = \hat{s}_i^2 + 2\hat{\lambda}_i$,

$$\tilde{\sigma}_{NT}^2 \equiv \frac{1}{T} \sum_{i=1}^{N} \hat{L}_{11i}^{-2} \hat{\sigma}_i^2 , \quad \hat{s}_i^{*2} \equiv \frac{1}{T} \sum_{t=1}^{T} \hat{\mu}_{it}^{*2} , \quad \tilde{s}_{NT}^{*2} \equiv \frac{1}{N} \sum_{i=1}^{N} \hat{s}_i^{*2} ,$$

$$\hat{L}_{11i}^2 = \frac{1}{T} \sum_{t=1}^{T} \hat{\eta}_{it}^2 + \frac{2}{T} \sum_{s=1}^{k_i} \left(1 - \frac{s}{k_i + 1} \right) \sum_{t=s+1}^{T} \hat{\eta}_{it} \hat{\eta}_{it-s}$$

The residuals $\hat{\mu}_{it}$, $\hat{\mu}_{it}^*$ and $\hat{\eta}_{it}$ are obtained from the following regressions

$\hat{e}_{it} = \hat{\rho}_i \hat{e}_{it-1} + \hat{\mu}_{it}$, $\hat{e}_{it} = \hat{\rho}_i \hat{e}_{it-1} + \sum_{k=}^{K_i} \hat{\gamma}_{ik} \Delta \hat{e}_{it-k} + \hat{\mu}_{it}^*$ and $\Delta y_{it} = \sum_{m=1}^{M} \hat{b}_{mi} \Delta X_{mit} + \hat{\eta}_{it}$

the alternative. The three Group Mean Estimators take the form of Phillips and Perron (1988) ρ-statistic and t-statistics, and a parametric t-statistic.

All seven tests are distributed as standard normal asymptotically. This requires a standardisation, which is based on the moments of the underlying Brownian motion functionals. The Panel variance ratio test will reject the null of cointegration for large positive values (i.e. greater than 1.96 at the 5% level) whereas the other six will reject the null of no cointegration with large negative values.

The Monte Carlo experiments of Pedroni (1997) suggest that the different statistics have varying comparative advantages, depending upon the situation under examination. The panel v-statistic generally performs the poorest. Among the seven tests the panel ρ-statistic suffered the smallest size distortion, the group ADF generally exhibited the largest and the group ρ-statistic exhibited empirical sizes that were too low in many cases. The group Phillips and Perron (PP), panel PP and the panel ADF test generally fell somewhere in between the group ADF and the panel ρ-statistic. All statistics, apart from the occasional exception of the panel variance statistic, performed well with respect to power.[36] For panels of T less than 100 - our panel data set -the group ADF generally did best with respect to power, followed by the panel ADF and the panel ρ-statistic. Overall, trading off size and power, the panel ρ-statistic appears to be the

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[36] The power of a test is equal to the probability of rejecting a false null hypothesis (i.e. 1 minus the probability of a type II error). This should preferably be close to one.

most consistently reliable statistic, particularly in situations with somewhat larger values for T.

4.6.1.2 McCoskey and Kao's LM Test for Panel Cointegration

The McCoskey and Kao (1998) LM Test is a residual-based test for the null hypothesis of cointegration in panel data. The test statistic is constructed using Lagrange Multipliers and is based on FM-OLS or Dynamic-OLS estimation: see Stock and Watson (1993) and Saikkonen (1991) regarding DOLS. The test is analogous to the Locally Best Unbiased Invariant (LBUI) for a moving average unit root (see Shin, 1994, and Harris and Inder, 1994). Again our panel variables are I(1) processes for all cross sections and can be written as the following varying coefficients regression

$$
\begin{aligned}
y_{it} &= \alpha_i + x'_{it}\beta_i + v_{it}, \quad i = 1,\dots,N, t = 1,\dots,T && (4.4)\\
x_{it} &= x_{it-1} + \varepsilon_{it} && (4.5)\\
v_{it} &= \gamma_{it} + u_{it} && (4.6)\\
\gamma_{it} &= \gamma_{it-1} + \theta u_{it}
\end{aligned}
$$

where u_{it} are iidN(0,σ_u^2). The null hypothesis is H$_0$:θ =0 against the alternative H$_a$: $|\theta| \neq 0$.[37] By backward substitution of *(4.6)* we write model *(4.4)* as

$$
\begin{aligned}
y_{it} &= \alpha_i + x'_{it}\beta_i + \theta\sum_{j=1}^{t} u_{ij} + u_{it} && (4.7)\\
&= \alpha_i + x'_{it}\beta_i + e_{it} && (4.7')
\end{aligned}
$$

From Tanaka (1996) we can construct a LBUI statistic and, as in *(4.8)*, a LM statistic

$$LM = \frac{\sum_{i=1}^{N}\sum_{t=1}^{T} S_{it}^2}{s^2} \tag{4.8}$$

S_{it} is partial sum process of the residuals,

$$S_{it} = \sum_{j=1}^{t} \hat{e}_{ij}$$

s^2 is a consistent estimator of σ_u^2 under H$_0$, where e_{it} is derived from equation (4.7')

$$s^2 = \frac{1}{NT}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{e}_{it}^2$$

The statistic in *(4.8)* is a LM statistic even if u_{it} is non-normal and serially correlated.

Based on Monte Carlo simulations of the means and variances from the functionals of Brownian motion of the limiting distribution, McCoskey and Kao emphasise DOLS over FM estimation when constructing our test statistics. This result is suggested for a small number of T, but does not present significant problems for our estimated results. Based on their examination of the empirical size of the two tests there does appear to be a slight tendency for LM-FM to over reject the null and LM-DOLS to slightly under-reject the null.

One issue regarding the null should be highlighted in the construction of the LM-DOLS and FM statistics. The LM tests presume that all the cross sections are equivalently cointegrated, unlike Pedroni's Group Mean statistics which allows there to be heterogeneity across cross sections in the degree of cointegration. Acceptance of the null is therefore

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[37] The null is cointegration against the alternative of no cointegration.

quite a strong test and rejection may well be due to at least one of the cross sections having a spurious long-run relationship.

4.6.2 Long-Run Panel Estimation

Pesaran, Shin and Smith's (1999) Pooled Mean Group Estimator (PMGE) was developed for panel data when both T and N are relatively large and of the same order. They emphasis that there are two traditional methods when using panel estimation: averaging or pooling. Averaging involves running N separate regressions and calculating coefficient means (see for example the Mean Group Estimator of Pesaran and Smith, 1995). This approach does not account for the fact that certain parameters are equal across cross sections. Alternatively we can pool the data and assume that the slope coefficients and error variance are identical. However, although there may be reasons to presume that the long-run coefficients are homogenous the reasons for the short-run dynamics and error variances being homogeneous are less compelling. Pesaran, Shin and Smith (1999) go on to propose an intermediate case between averaging and pooling, which involves aspects of both. For their PMGE the long-run coefficients are constrained to be equal, whereas short-run coefficients and error variance differ across groups. We can therefore obtain pooled long-run coefficients, and averaged short run dynamics as an indication of mean reversion.

The PMGE is based on an AutoRegressive Distributive Lag (p,q,…,q) model

$$y_{it} = \sum_{j=1}^{p} \lambda_{ij} y_{i,t-j} + \sum_{j=0}^{q} \delta'_{ij} \mathbf{x}_{i,t-j} + \mu_i + \varepsilon_{it} \qquad (4.9)$$

where $\mathbf{x}_{it}$ (k*1) is the vector of explanatory variables for group i, μ_i represents the fixed effects, the coefficients of the lagged dependent variables (λ_{ij}) are scalers and δ_{ij} are (k*1) coefficient vectors. T must be large enough such that the model can be estimated for each cross section.

Equation *(4.9)* is then re-parameterised as

$$\Delta y_{it} = \phi_i y_{i,t-1} + \beta'_i \mathbf{x}_{it} + \sum_{j=1}^{p-1} \lambda^*_{ij} \Delta y_{i,t-j} + \sum_{j=0}^{q-1} \delta^{*\prime}_{ij} \Delta \mathbf{x}_{i,t-j} + \mu_i + \varepsilon_{it} \qquad (4.10)$$

where $\phi_i = -\left(1 - \sum_{j=1}^{p} \lambda_{ij}\right)$, $\beta_i = \sum_{j=0}^{q} \delta_{ij}$, $\lambda^*_{ij} = -\sum_{m=j+1}^{p} \lambda_{im}$ and $\delta^*_{ij} = -\sum_{m=j+1}^{q} \delta_{im}$

They also make the following assumptions. Firstly, the residuals in *(4.10)* are iid with zero mean, variance greater than zero and finite fourth moments. Secondly, the roots of equation *(4.10)* must lie outside the unit circle. The latter assumption ensures that $\phi_i<0$, and hence that there exists a long-run relationship between y_{it} and x_{it} defined by

$$y_{it} = -\left(\beta'_i / \phi_i\right) \mathbf{x}_{it} + \eta_{it}$$

The long-run homogeneous coefficient is equal to $\theta = \theta_i = -\left(\beta'_i / \phi_i\right)$, which is the same across groups. The PMGE uses a maximum likelihood approach to estimate the model and a Newton-Raphson algorithm. The lag length for the model can be determined using, for instance, the Akaike Information Criteria.

A couple of points are worth emphasising. The estimated coefficients in the model are not dependent upon whether the variables are I(1) or I(0). Hence this approach is more general than the cointegration panel approach. The key feature of the PMGE is it makes the long-run relationships homogenous while allowing for the heterogeneous dynamics and error variances.

4.7 Panel Results

In this section we examine our non-stationary panel data results. We pre-test the data for stationarity, test for the existence of panel cointegration and estimated the pooled long run coefficients for our different specifications using the PMGE.

4.7.1 Panel Unit Root Tests

It is useful to consider initially the panel stationarity properties of each variable in our analysis. We do this using the Levin and Lin (1993) panel ADF tests. Reserves, the real effective exchange rate, income, volatility and the average propensity to import all appear panel unit root (the results are in Appendix 4.C). The results for inflation and prices on the other hand, mostly indicate stationarity. This reflects the results from the univariate test for the series, which suggested evidence of stationarity. On that basis it may be an idea to exclude these variables at this stage of the panel analysis. However, given the quite substantive debate in the literature (see Larsson et al., 1998) as to whether inflation is stationary or non-stationary, we can be reasonably sure that prices are unlikely to be I(0).[38]

4.7.2 Panel Cointegration Tests

The first Pedroni (1998) test that we examine is the variance ratio test (v-statistic). This test is included for completeness but, from the above discussion, we know that the small sample properties are consistently poor. The second test is a panel version of a non-parametric statistic similar to the Phillips-Perron (PP) ρ-statistic. The third is analogous to the PP t-statistic. The fourth panel cointegration statistic is similar to an ADF test. The other three statistics from Pedroni are based on the Group Mean approach. These are akin to the PP ρ-statistic and t-statistic, and the ADF test. All seven tests are based on the null of no cointegration. Large positive values reject the null for the panel v-statistic, while large negative values reject the null for the other tests. For the parametric tests a maximum lag length of five is allowed for each cross section.

The McCoskey and Kao (1998) multivariate panel cointegration tests are based on a LM approach utilising DOLS and FM-OLS. The null hypothesis for these tests is cointegration. Statistics greater than the critical values reject the null. In the case of LM-DOLS a common lead

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[38] The problem with some of the univariate tests (e.g. France) is apparently being directly transferred to the Levin and Lin (1993) test, which is unsurprising. Indeed if we utilise the t-bar tests of Im, Pesaran and Shin (1997) we have estimated statistics for the case without and with trend of 1.44 and 5.50, respectively. Where we require our estimates to

Table 4.11: Panel Cointegration Tests

	Q Y P	*Q Y INF*	*Q Y*	*Q Y VOL*	*VOL Y*	*VOL Y AP*
Pedroni's Panel Statistics						
v-stat	1.12	0.40	1.16	2.36**	3.09**	2.39**
PP ρ-stat	-1.27	-0.16	-1.15	-1.47	-1.87	-1.69
PP t-stat	-2.11**	-0.71	-1.61	-2.05**	-1.99**	-2.39**
ADF t-stat	-3.32**	-1.64	-2.26**	3.62**	-3.41**	-4.57**
Pedroni's Group Mean Statistics						
PP ρ-stat	-0.86	-0.24	-0.75	-1.13	-0.97	-1.08
PP t-stat	-2.16**	-0.96	-1.50	-2.17**	-1.73	-2.42**
ADF t-stat	-5.21**	-3.55**	-3.18**	-5.24**	-3.82**	-6.50**
McCoskey and Kao's LM Tests						
LM DOLS(2,2)	-0.38	-0.11	1.45	0.16	1.35	-1.02
LM-DOLS(1,1)	1.09	2.06**	3.43**	1.71**	3.06**	0.72
LM-FM (3)	2.27**	5.82**	6.82**	5.30**	5.26**	5.23**
LM-FM (5)	1.68**	3.64**	4.63**	4.18**	4.40**	4.78**

Notes: At the 5% level the critical value is –1.96 for all Pedroni's tests, with large negative values rejecting the null of no cointegration. (That is excluding the panel variance ratio test, which has a critical value of 1.96 and the null hypothesis of no cointegration is rejected for large positive test statistics). Maximum number of lags in Pedroni's parametric tests is 5.

McCoskey and Kao's LM test is based on the null of cointegration, hence it rejects the null for large positive values (i.e. values greater than 1.645). The number of cross sections is 9 and number of time periods is 62 for all statistics.

and lag length was chosen. For LM-FM a common lag length of Barlett window was used. For both of the McCoskey and Kao tests the results were dependent upon the choice of lag length: for fullness we present all our results.

All specifications exhibit some evidence of cointegration, but consistency in results is not always present. We see from McCoskey and Kao's LM test statistic that the null of cointegration is accepted at the 5% level for all equations, that is to say that estimated statistics are less than 1.645. There is some evidence of cointegration for the second and third specifications but this is not particularly strong for either. The specification including the real exchange rate, income and prices gives a count of 6 out of 11 tests supporting cointegration at the 5% level. This result was the

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be less than -1.96 to reject the null of unit root, prices are not even close to being I(0). This again discounts any problem with prices.

same when we incorporated volatility instead of prices. The more traditional specification, which included volatility, income and AP, was the best with 7 out of 11 tests suggesting cointegration. However, given that Pedroni (1997) places little emphasis on the usefulness of the variance ratio test there is little to differentiate between our newer specification and the more traditional set up.

An important part of the analysis is to take account of time effects that impart on cross sections equally. We do this by removing cross section time means for each variable. The results taking account of time effects in Table 4.12 are quite different in nature to those with raw data. This difference suggests that there may be important time effects at work in our panel. Specification four (Q, Y and VOL) and six (VOL, Y and AP) both have counts of 3 out of 11. However, regression one is quite emphatically supported by the LM statistics and gives a total count of six out of eleven when we also examine the Pedroni results. The results with time means removed emphasise equation one.[39,40]

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[39] It should be noted from Section 4.6.1 that the technique of removing time means may introduce data dependencies into the results.

[40] We did not incorporate McCoskey and Kao (1998) t-bar test into our analysis since the results consistently rejected all our specifications. This contrast with the results when we used Johansen (1988) approach for cointegration in a panel context following Larsson et al. (1998) which produced consistent results of cointegration.

Table 4.12: Panel Cointegration Test Removing Time Effects

	Q Y P	*Q Y INF*	*Q Y*	*Q Y VOL*	*VOL Y*	*VOL Y AP*
Pedroni's Panel Statistics						
v-stat	1.08	0.32	0.48	1.17	1.32	0.55
PP ρ-stat	-0.88	0.37	-0.05	-0.77	-0.77	0.03
PP t-stat	-1.40	0.06	-0.27	-1.33	-1.03	-0.53
ADF t-stat	-2.91**	0.59	-0.10	-2.00**	-1.45	-1.33
Pedroni's Group Mean Statistics						
PP ρ-stat	0.19	0.94	0.52	-0.78	-0.61	-0.21
PP t-stat	-0.87	0.35	-0.05	-1.60	-1.08	-0.85
ADF t-stat	-2.58**	0.50	-0.16	-4.31**	-2.74**	-3.51**
McCoskey and Kao's LM Test						
LM-DOLS (2,2)	-1.96	0.11	1.77**	-0.47	2.01**	-0.05
LM-DOLS (1,1)	1.09	2.06**	3.43**	1.71**	3.06**	0.72
LM-FM (3)	1.40	5.87**	10.45**	5.42**	6.09**	6.45**
LM-FM(5)	1.30	4.57**	8.18**	3.42**	4.00**	4.03**

Notes: See Table 4.11.

4.7.3 Long and Short-Run Panel Estimation

In addition to the evidence provided by cointegration methods, we also consider the long run estimated coefficients in a panel environment. Furthermore, we consider the speed of mean reversion to the long-run. Our results for long-run panel estimation using the Pooled Mean Group Estimation are included in Table 4.13.

Table 4.13: Pooled Mean Group Estimation

Regressors	*Eqtn. 1*	*Eqtn. 2*	*Eqtn. 3*	*Eqtn. 4*	*Eqtn. 5*	*Eqtn. 6*
Q	1.86	1.75	1.71	0.33		
	(8.64)**	(8.68)**	(8.59)**	(1.41)		
Y	0.06	0.05	0.05	0.06	0.04	0.06
	(4.49)**	(4.30)**	(4.62)**	(4.76)**	(2.38)**	(7.10)**
P	-1.19					
	(-2.23)**					
INF		0.02				
		(0.05)				
VOL				0.10	0.14	-0.39
				(1.62)	(1.75)	(-7.05)**
AP						-0.65
						(-22.22)**
Mean Adjust. Coeff.	-0.64	-0.59	-0.58	-0.61	-0.62	-1.05
	(-8.03)**	(-8.53)**	(-8.90)**	(-10.71)**	(-10.29)**	(-19.39)**

Notes: The lag length is determined using AIC with a maximum of one for p and q in the ARDL(p,q,...,q) specification. T-statistics are in parenthesis. Variables significant at the 5% level are indicated by **.

These results suggest that specifications one and six have coefficients that are all significant at the 5% level. Income itself is consistently significant and positive, although not with a particularly large coefficient. The exchange rate coefficient is always positive and typically significant, apart from equation four where we include our measure of volatility. A more traditional specification does not perform quite so well in this case. Although all coefficients are significant in specification six the signs of the estimated coefficients are not exactly as expected. The mean adjustment coefficient is greater than one, which is also rather curious.

The Pooled Mean Group estimator of Pesaran, Shin and Smith (1999) assumes that cross sections of our panel are independent. This is similar to the assumption used with panel cointegration tests. This is important, because there is likely to be a common factor problem, which involves common shocks impacting upon our countries. This could take the form of an important political effect that influences the widespread confidence of European institutions. Or alternatively this may be an external change, which all countries experience equally, like a change in US bilateral rates. We deal with this problem, following Pesaran et al. (1999), by taking account of time effects. The results are given in Table 4.14.

Since the results in Table 4.14 are robust to non-independence of our cross sections, we are inclined to place more importance upon them in the final analysis. Nevertheless, once we take account of the common

Table 4.14: Pooled Mean Group Estimation Removing Time Effects

Regressors	*Eqtn. 1*	*Eqtn. 2*	*Eqtn. 3*	*Eqtn. 4*	*Eqtn. 5*	*Eqtn. 6*
Q	1.35 (2.88)**	0.09 (0.14)	-0.22 (-0.34)	-0.54 (-0.87)		
Y	0.05 (3.00)**	0.01 (0.50)	0.02 (1.43)	0.03 (1.96)**	0.04 (1.30)	0.07 (9.38)**
P	-1.68 (-3.28)**					
INF		-0.84 (-1.27)				
VOL				0.16 (2.36)**	0.10 (0.81)	-0.51 (-10.66)**
AP						-0.68 (-26.93)**
Mean Adjst. Coeff.	-0.54 (-9.20)**	-0.48 (-13.22)**	-0.49 (-13.75)**	-0.55 (-7.75)**	-0.53 (-13.83)**	-1.06 (-18.43)**

Notes: The lag length is determined using AIC with a maximum of one for p and q in the ARDL(p,q,...,q) specification. Variables significant at the 5% level are indicated by ** t-statistics in parenthesis.

factor problem our results generally replicate those using raw data. Equation 1 has equivalently signed, sized and significant coefficients to results in Table 4.13. Also this equation has an average short-run adjustment parameter which is error correcting and significant. This is much larger than for UK results in previous Chapters. This suggests that European countries may be slightly more preoccupied about their long-run target of reserves. This is consistent with their more formal exchange rate arrangements and greater certainty regarding policy. The more traditional specification of equation 6 has all coefficients significant but the signs are again of a curious nature and certainly not consistent with theory. The adjustment coefficient suggests dynamic instability.

In summary, specification one which includes Q, Y and P dominates the more traditional specifications when we use PMGE, irrespective of whether time effects are removed or not. And the new specification is also preferred with the panel cointegration tests after removing time effects.

4.8 Conclusion

In our Forth Chapter we have examined the demand for international reserves for nine European countries for the period 1980-95. In an attempt to discriminate between different specifications for reserves we have made use of time-series and panel data methods. The broad range of estimator that we have used, are a means of corroborating the previous results that we have found.

From the time series results, and in terms of significant and correctly signed coefficients, stability and evidence of cointegration, the traditional specification, which included income, volatility and the average propensity to import, was only acceptable for Spain. This questioned the usefulness of this approach for countries where intervention in the foreign exchange market is a primary reason for holding international reserves. We have therefore examined the possibility of constructing an alternative specification. However, using purely time series methods did not allow us to accept an alternative specification, which passed all the tests, for all the countries. Nevertheless we did find recurring results for each of the countries and, more often than not, the newer specification did dominate.

The time-series results suggested that income was consistently important as a long-run determinant of reserves for all countries, apart from France. We also found considerable evidence that the real effective exchange rate was important. This was the case for the UK, Germany, Italy, Spain, and Holland. And marginally so in the case of France, Austria and Sweden. However there was a degree of variation in which other

4.8 Conclusion

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The time-series results suggested that income was consistently important as a long-run determinant of reserves for all countries, apart from France. We also found considerable evidence that the real effective exchange rate was important. This was the case for the UK, Germany, Italy, Spain, and Holland. And marginally so in the case of France, Austria and Sweden. However there was a degree of variation in which other variables combined with Q and Y to form a preferred long-run relationship for each of the countries. Some included prices or inflation (Germany, France and Italy) others include the more traditional measure of volatility (Spain, Austria and Sweden).

We then turned to non-stationary panel methods to provide more powerful tools of analysis. The Pooled Mean Group Estimator of Pesaran et al. (1999) suggested an alternative specification of prices, income and the real effective exchange rate as a sensible long-run relationship for reserves. This was not inconsistent with the panel cointegration results of Pedroni and McCoskey and Kao. Although univariate and panel ADF tests often suggested that prices were stationary, the evidence in Larsson *et al.* (1998) refutes this possibility. Indeed, if we further our analysis by utilising the more recent panel unit root (t-bar) test of Im, Pesaran and Shin (1997) there is no problem. Given this evidence and the increase in power of the panel methods we are persuaded by the alternative specification.

Additionally we found it important to remove any common factor problem. This is a means of correcting for common shocks impacting upon the countries sampled. This is obviously important within a European context due to political and business cycle pressures and external fluctuations against the Dollar. Taking account of this problem also supported an alternative specification including Q, Y and P. Interestingly the short-run dynamic adjustment term provided evidence of a large amount of disequilibrium in the long-run reserves relationship being error corrected. This contrasts to a certain extent with what has gone before but is substantial evidence in favour of a reserve equation in a European context.

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Appendix 4.A: Univariate Augmented Dickey Fuller Tests

Table 4.A.1: Univariate ADF Tests for Reserves

	Model 1		*Model 2*		*Model 3*	
	Lags	*Test*	*Lags*	*Test*	*Lags*	*Test*
UK	3	0.3932	3	-1.6455	3	-3.4401
Germany	2	0.8288	2	-0.8524	0	-3.6958**
France	0	0.0530	0	-1.7763	0	-2.0999
Italy	3	0.4541	3	-1.5558	3	-1.8325
Spain	0	1.8893	0	0.4116	0	-2.2973
Netherlands	5	0.7717	5	-1.0114	4	-1.9053
Austria	0	2.4730	0	0.3931	0	-1.6080
Finland	0	0.9574	0	-0.9131	0	-2.4490
Sweden	1	2.8015	1	0.0802	0	-3.1020

Notes: 5% critical value for model one is –1.95, for model two –2.93 and for model three -3.50.
Values significant at the 5% level (i.e. large negative values) marked by *.
Critical values are due to Fuller (1976).

Table 4.A.2: Univariate ADF Tests for Real Effective Exchange Rate

	Model 1		*Model 2*		*Model 3*	
	Lags	*Test*	*Lags*	*Test*	*Lags*	*Test*
UK	0	-0.8839	0	-1.3962	0	-2.8560
Germany	0	2.0981	0	1.0063	0	-2.6952
France	2	-0.8429	2	-1.9859	0	-3.0221
Italy	0	-1.0553	0	0.0805	0	-0.3504
Spain	0	-1.3203	0	-3.3727*	0	-2.9383
Netherlands	0	-0.3698	0	-1.4893	0	-1.8394
Austria	0	-1.6753	0	-0.9234	0	-2.6948
Finland	5	-0.4343	4	-3.3627*	4	-3.9603*
Sweden	0	-1.3126	1	-1.4788	1	-1.6973

Notes: See Table 4.A.1.

Table 4.A.3: Univariate ADF Tests of Income

	Model 1		*Model 2*		*Model 3*	
	Lags	*Test*	*Lags*	*Test*	*Lags*	*Test*
UK	1	2.8955	1	-0.8209	1	-1.3898
Germany	0	3.2873	0	0.4885	4	-2.6320
France	2	2.3461	2	-0.5441	2	-2.0145
Italy	2	2.7401	2	-0.7656	2	-1.6631
Spain	0	4.4954	0	0.7098	0	-2.9702
Netherlands	1	1.8325	1	-1.1602	1	-2.2707
Austria	2	4.6690	2	0.6292	2	-2.3260
Finland	0	2.8046	0	-2.0012	5	-3.1452
Sweden	4	1.5387	4	-1.5318	4	-2.3521

Notes: See Table 4.A.1.

Table 4.A.4: Univariate ADF Tests for Inflation

	Model 1		***Model 2***		***Model 3***	
	Lags	*Test*	*Lags*	*Test*	*Lags*	*Test*
UK	2	-2.7797*	1	-4.3414*	0	-6.1595*
Germany	3	-0.9883	3	-2.0328	3	-2.0154
France	5	-2.7540*	5	-1.7854	5	-1.2388
Italy	1	-1.9250	1	-2.1310	0	-3.6475*
Spain	3	-2.0786*	3	-2.7062	3	-2.5182
Netherlands	4	-1.0689	0	-8.1055*	0	-9.7611*
Austria	4	-1.0020	4	-1.9731	4	-2.2995
Finland	4	-1.7517	4	-2.1553	0	-7.5824*
Sweden	4	-2.2095*	4	-3.6819*	4	-4.2031*

Notes: 5% critical value for model one is –1.95, for model two –2.93 and for model three -3.50. See Table 4.A.1.

Table 4.A.5: Univariate ADF Tests for Prices

	Model 1		***Model 2***		***Model 3***	
	Lags	*Test*	*Lags*	*Test*	*Lags*	*Test*
UK	3	-0.6763	0	-4.5259*	0	-1.8139
Germany	4	-0.5291	4	-0.2453	4	-2.7479
France	2	-0.3867	0	-13.0773*	0	-5.1202*
Italy	2	-0.5958	1	-3.9126*	1	-4.1991*
Spain	4	-1.4597	0	-3.5479*	0	-1.8095
Netherlands	4	-0.8479	4	-0.8635	4	-3.3237
Austria	5	-0.3624	0	-2.5500	5	-3.3368
Finland	5	-1.3108	0	-4.9961*	0	-1.4832
Sweden	5	-1.5819	5	-2.1108	5	-0.8885

Note: See Table 4.A.1.

Table 4.A.6: Univariate ADF Tests for Volatility

	Model 1		***Model 2***		***Model 3***	
	Lags	*Test*	*Lags*	*Test*	*Lags*	*Test*
UK	0	1.0170	0	-1.2004	0	-1.3431
Germany	0	-0.7308	0	-1.6913	0	-2.0256
France	4	0.5605	4	-1.6394	4	-2.9953
Italy	4	-0.0964	4	-2.4499	4	-2.5392
Spain	5	-0.5535	4	-3.0435*	4	-3.4882
Netherlands	4	-0.1670	4	-2.6742	4	-2.4471
Austria	0	-0.4097	0	-1.2062	0	-1.1183
Finland	1	-1.8625	1	-1.9138	1	-2.0759
Sweden	0	-1.3184	0	-1.3535	4	-2.1431

Notes: See table 4.A.1.

Table 4.A.7: Univariate ADF Tests for Average Propensity to Import

	Model 1		***Model 2***		***Model 3***	
	Lags	*Test*	*Lags*	*Test*	*Lags*	*Test*
UK	5	-0.6177	5	-1.8096	5	-1.8058
Germany	4	0.5255	4	-1.9123	4	-3.4474
France	4	0.2190	4	-2.5917	4	-3.8930*
Italy	4	0.1220	4	-2.6825	4	-2.8416
Spain	4	-0.4946	4	-2.1923	4	-2.4073
Netherlands	5	0.5763	4	-2.4138	4	-3.8127*
Austria	4	0.2390	4	-3.2229*	4	-3.5734*
Finland	3	0.7505	3	-1.8831	3	-0.8470
Sweden	4	-0.9758	4	-2.1235	4	-2.2653

Notes: 5% critical value for model one is –1.95, for model two –2.93 and for model three -3.50. See table 4.A.1.

Appendix 4.B: Johansen Results

Table 4.B.1: Johansen Results for β_i =(R, Q, Y, P, constant)

	Trace Statistic				***Long-Run Coefficients*** ***β_i= (R, Q, Y, P, constant)***	***Mean Reversion***
H_o:	r=0	r≤1	r≤2	r≤3		
95% c.v.	53.12	34.91	19.96	9.24		
90% c.v.	49.65	32.00	17.85	7.52		
Britain	94.12 **	41.75 **	15.15	5.91	β1= (1.00,1.68,9.43,-9.79,-283.32) β2= (1.00,-15.58,-6.91,-3.48,234.26)	α1= -0.00 α2= 0.05
Germany	115.53 **	52.93 **	27.49 **	6.55	β1= (1.00,-18.32,52.44,-49.24,-1430.90) β2= (1.00,-4.88,-2.56,5.94,70.63) β3= (1.00,-1.62,-0.08,-0.56,-15.20)	α1= -0.00 α2= -0.09 α3= -0.02
France	85.87 **	38.66 **	21.28 **	9.25 **	β1= (1.00,6.30,-1.68,7.20,-4.78) β2= (1.00,-82.01,-8.19,-8.91,593.58) β3= (1.00,16.86,14.85,-6.16,-539.89) β4= (1.00,-2.70,-1.56,-0.21,34.48)	α1= -0.02 α2= -0.01 α3= -0.03 α4= -0.14
Italy	60.15 **	27.48	13.21	1.85	β1= (1.00,-2.72,-3.70,-1.96,117.90)	α1= -0.02
Spain	100.09 **	50.12 **	22.33 **	6.68	β1= (1.00,-10.66,-1.71,-3.46,73.30) β2= (1.00,-15.30,-22.59,6.91,656.93) β3= (1.00,-18.38,2.11,-4.50,4.08)	α1= 0.01 α2= 0.02 α3= -0.04
Nether-lands	64.09 **	38.41 **	18.84	9.06	β1= (1.00,-7.83,-1.40,-1.45,55.25) β2= (1.00,-8.71,6.32,-17.72,-182.59)	α1= 0.02 α2= 0.01
Austria	91.38 **	38.43 **	16.17	4.38	β1= (1.00,17.71,-8.80,11.88,127.03) β2= (1.00,12.37,14.31,-9.04,-463.09)	α1= -0.04 α2= -0.01
Finland	85.89 **	47.27 **	20.21 **	4.58	β1= (1.00,12.99,-70.21,37.57,1707.86) β2= (1.00,-17.49,62.49,-32.98,-1539.64) β3= (1.00,2.13,-8.92,0.89,195.53)	α1= -0.00 α2= -0.00 α3= -0.31
Sweden	107.72 **	53.69 **	26.16 **	9.70	β1= (1.00,19.27,-51.97,20.81,1266.16) β2= (1.00,-2.70,14.26,-7.65,-390.08) β3= (1.00,1.76,-1.75,-2.26,15.08) β4= (1.00,-2.31,-5.18,-2.45,123.82)	α1= -0.01 α2= -0.05 α3= -0.04 α4= -0.09

Note: Critical Values (c.v.) are from Osterwauld-Lenum (1992).

Table 4.B.2: Johansen Results for β_i = (R, Q, Y, INF, constant)

	Trace Statistic				Long-Run Coefficients	Mean
					β_i= (R, Q, Y, INF, constant)	Reversion
H_o:	r=0	r≤1	r≤1	r≤1		
95% c.v.	53.12	34.91	19.96	9.24		
90% c.v.	49.65	32.00	17.85	7.52		
Britain	78.55 **	42.27 **	22.09 **	6.99	β1= (1.00,-77.99,-28.03,219.95,1080.21) β2= (1.00,-12.39,-13.57,-12.75,400.15) β3= (1.00,2.97,-7.01,-2.25,151.95)	α1= 0.00 α2= 0.03 α3= -0.03
Germany	74.91 **	40.31 **	16.19	5.41	β1= (1.00,-5.25,2.32,-14.99,-66.58) β2= (1.00,-0.50,-1.51,6.82,20.07)	α1= -0.16 α2= -0.28
France	59.81 **	32.65 *	14.12	3.69	β1= (1.00,-8.22,5.17,26.52,-140.09) β2= (1.00,22.34,7.82,10.29,-357.40)	α1= -0.04 α2= -0.02
Italy	59.24 **	25.54	10.47	1.83	β1= (1.00,-2.45,-8.40,-11.34,279.81)	α1= -0.06
Spain	81.14 **	43.90 **	24.40 **	6.42	β1= (1.00,-9.73,-5.42,-22.26,168.55) β2= (1.00,-22.86,-9.36,14.24,333.88) β3= (1.00,-3.73,-4.39,18.18,112.35)	α1= 0.05 α2= -0.01 α3= 0.01
Nethrlds	75.87 **	35.80 **	17.01	7.02	β1= (1.00,-3.30,0.67,23.58,-31.73) β2= (1.00,-3.45,-4.19,-0.13,123.61)	α1= -1.80 α2= -0.05
Austria	78.18 **	39.79 **	14.85	2.87	β1= (1.00,20.39,8.32,74.02,-343.40) β2= (1.00,85.52,54.33,67.24,-1869.76)	α1= -0.03 α2= 0.00
Finland	71.28 **	29.04	10.74	2.53	β1= (1.00,-11.66,100.50,-126.78)	α1= -0.00
Sweden	114.18 **	47.71 **	22.19 **	9.72 **	β1= (1.00,1.91,-3.94,15.86,71.59) β2= (1.00,4.73,-6.72,-2.51,133.27) β3= (1.00,1.91,-9.59,-0.86,222.70) β4= (1.00,-4.35,-17.29,-0.99,454.07)	α1= -0.03 α2= -0.05 α3= 0.00 α4= -0.04

Table 4.B.3: Johansen Results for β_i =(R, Q, Y, constant)

	Trace Statistic			Long-Run Coefficients	Mean
				β_i= (R, Q, Y, constant)	Reversion
H_o:	r ≤0	r ≤1	r≤2		
95% c.v.	34.91	19.96	9.24		
90% c.v.	32.00	17.85	7.52		
Britain	44.61 **	21.45 **	8.92*	β1= (1.00,-12.52,-12.37,367.73) β2= (1.00,6.14,-6.64,127.32) β3= (1.00,-1.38,-5.51,130.82)	α1= 0.04 α2= -0.02 α3= -0.07
Germany	41.29 **	16.52	5.85	β1= (1.00,-1.91,-0.33,-6.83)	α1= -0.49
France	34.80 **	16.73	5.63	β1= (1.00,35.49,8.51,-437.66)	α1= -0.00
Italy	33.43 *	12.50	1.92	β1= (1.00,0.78,-4.37,123.14)	α1= -0.01
Spain	52.36 **	25.62 **	7.12	β1= (1.00,-12.46,-6.11,199.26) β2= (1.00,-61.96,-20.38,809.97)	α1= 0.04 α2= -0.00
Nethlds	35.29 **	16.61	6.53	β1= (1.00,-3.46,-4.16,122.75)	α1=-0.05
Austria	36.45 **	15.20	2.92	β1= (1.00,-19.25,-18.01,546.55)	α1= -0.02
Finland	29.19	11.03	2.36		
Sweden	45.56 **	20.14 **	9.36**	β1= (1.00,4.01,-6.86,140.16) β2= (1.00,-0.10,-11.39,279.06) β3= (1.00,10.84,2.04,-124.36)	α1= -0.07 α2= -0.05 α3= 0.02

Table 4.B.4: Johansen Results for β_i =(R, Q, Y, VOL, constant)

	Trace Statistic				***Long-Run Coefficients*** ***β_i= (R, Q, Y, VOL, constant)***	***Mean Reversion***
H₀:	*r* =0	*r* ≤1	*r* ≤2	*r*≤3		
95% c.v.	53.12	34.91	19.96	9.24		
90% c.v.	49.65	32.00	17.85	7.52		
Britain	59.95 **	34.81 *	14.47	2.69	β1=(1.00,-15.89,-14.08,0.77,431.29) β2=(1.00,-0.85,-3.19,1.74,69.77)	α1= 0.03 α2= -0.01
Germany	73.52 **	35.21 **	15.13	6.08	β1=(1.00,-2.35,0.06,0.08,-15.60) β2=(1.00,-0.84,0.02,-0.31,-22.18)	α1= -0.68 α2= 0.07
France	48.69	27.13	13.72	4.08		
Italy	45.22	21.00	9.36	2.71		
Spain	72.41 **	45.08 **	23.67 **	5.44	β1= (1.00,84.18,14.32,-3.14,-808.48) β2= (1.00,-1.72,-3.84,-0.38,87.06) β3= (1.00,-5.24,-4.88,-0.35,131.23)	α1= -0.01 α2= -0.27 α3= -0.18
Netherlands	67.49 **	29.90	17.19	7.15	β1= (1.00,-2.38,-4.22,0.02,119.91)	α1= -0.14
Austria	59.04 **	27.11	14.24	2.59	β1= (1.00,-3.49,-6.29,-0.91,159.86)	α1= 0.05
Finland	63.29 **	31.10	14.36	3.29	β1= (1.00,-3.31,3.61,-2.98,-105.25)	α1= -0.08
Sweden	59.73 **	34.00 *	19.95	8.87	β1= (1.00,5.27,-7.52,0.22,152.49) β2= (1.00,-3.98,-1.77,-1.35,38.00)	α1= -0.05 α2= -0.06

Table 4.B.5: Johansen Results for β_i =(R, VOL, Y, constant)

	Trace Statistic			***Long-Run Coefficients*** ***β_i= (R, VOL, Y, constant)***	***Mean Reversion***
H₀:	*r* =0	*r* ≤1	*r*≤2		
95% c.v.	34.91	19.96	9.24		
90% c.v.	32.00	17.85	7.52		
Britain	30.96	11.52	2.25		
Germany	38.93 **	19.88 **	5.34	β1= (1.00,-0.52,-0.36,-15.66) β2= (1.00,0.05,-1.56.19.83)	α1= -0.07 α2= -0.35
France	34.38 **	17.57	4.43	β1= (1.00,2.15,6.78,-218.06)	α1= -0.03
Italy	33.28 *	10.54	4.37	β1= (1.00,2.93,-4.03,120.84)	α1= -0.00
Spain	40.04 **	17.24	3.85	β1= (1.00,-0.51,-3.95,81.61)	α1= -0.11
Netherlands	34.86 *	14.04	3.53	β1= (1.00,0.25,-4.79,127.83)	α1= -0.12
Austria	43.00 **	12.23	2.44	β1= (1.00,-0.72,-4.14,86.53)	α1= 0.04
Finland	32.64 *	15.47	3.94	β1= (1.00,-0.57,-3.83,74.04)	α1= -0.20
Sweden	29.86	11.71	2.64		

Table 4.B.6: Johansen Results for β_i =(R, VOL, Y, AP, constant)

	Trace Statistic				Long-Run Coefficients β_i= (R, VOL, Y, AP,constant)	Mean Reversion
H_o:	$r=0$	$r\leq1$	$r\leq2$	$r\leq3$		
95% c.v.	53.12	34.91	19.96	9.24		
90% c.v.	49.65	32.00	17.85	7.52		
Britain	61.78 **	28.37	11.85	2.33	β1=(1.00,-0.89,-6.82,5.71,174.32)	α1= -0.01
Germany	71.62 **	34.53 *	12.11	4.37	β1=(1.00,0.08,-0.76,1.57,1.57) β2=(1.00,-4.48,10.32,2.80,319.07)	α1= -0.72 α2= 0.01
France	54.54 **	30.86	17.28	4.64	β1=(1.00,1.74,6.45,1.92,-203.57)	α1= -0.04
Italy	48.09	22.14	10.25	4.07		
Spain	61.25 **	34.96 **	17.27	5.36	β1= (1.00,-0.54,-0.52,-0.42,-10.58) β2= (1.00, -0.48,-3.20,-0.41,60.36)	α1= 0.02 α2= -0.32
Netherlands	56.78 **	25.38	12.43	3.39	β1= (1.00,0.30,-8.00,-4.63,219.27) β2= (1.00,1.11,-4.85,0.46,132.97)	α1= -0.06 α2= -0.06
Austria	58.10 **	22.30	10.26	2.99	β1= (1.00,-0.64,-4.02,1.30,85.11)	α1= 0.05
Finland	64.41 **	36.16 **	17.61	5.61	β1= (1.00,-1.59,-9.59,-7.10,208.10) β2= (1.00,-0.76,-0.62,2.44,-4.42)	α1= -0.14 α2= -0.16
Sweden	50.18 *	23.22	12.98	3.74	β1= (1.00,-0.49,-7.55,-5.87,167.48)	α1= -0.08

Appendix 4.C: Panel Unit Root Results

Table 4.C.1: Levin and Lin (1993) Panel Unit Root Tests

Variable	Time Means	No Constant No Trend	Constant No Trend	Constant Trend
R	Not Removed	3.48	1.59	-1.04
	Removed	-3.52**	-0.01	0.24
Q	Not Removed	-1.80	0.36	-1.02
	Removed	-3.70**	1.10	0.47
Y	Not Removed	7.18	0.69	-0.30
	Removed	-0.51	-0.42	-0.97
INF	Not Removed	-4.95**	-3.03**	-7.50**
	Removed	-9.13**	-10.48**	-16.79**
P	Not Removed	-2.05**	-12.63**	-3.32**
	Removed	-8.88**	-10.32**	-0.04
VOL	Not Removed	-0.80	-0.93	0.65
	Removed	-1.60	0.02	1.60
AP	Not Removed	0.09	-2.01**	-0.51
	Removed	-0.12	0.48	3.45

Notes: The Critical Value for all tests at the 5% is –1.96. Large negative values reject the null hypothesis of non-stationarity. Stationary time series represented by **.

Chapter Five

Foreign Exchange Intervention and Exchange Rate Volatility: Policy Coordination and Spillovers

5.1 Introduction

In this Chapter, we are interested in whether attempts to remove misalignment in the nominal exchange rate have been at the expense of increased market disorder measured by daily volatility. We do this by considering the higher frequency effect of intervention on exchange rate volatility over the period 1985 to 1987. 1985 represents the return of the US Federal Reserve to the Foreign Exchange market - after a hiatus of five years - and renewed attempts by the G-5 countries to remove Dollar misalignment. The end of our sample is 1987 when exchange rate policy changed again with the Louvre Accord and a determination to keep the US Dollar in more narrow bands.

In one of the earlier papers in the field, which motivated a substantial research effort, Dominguez and Frankel (1993a) found evidence that daily exchange rate intervention significantly influenced the level of the DM-Dollar and Yen-Dollar rates between 1984 and 1990. In this thesis on international reserves we are interested in the influence of daily intervention on exchange rate volatility. There have been a number of recent studies on intervention and volatility using as many different approaches. Our study focuses on daily exchange rates using Generalised AutoRegressive Conditional Heteroskedasticity (GARCH) techniques with daily intervention data. Given Edison's (1993) suggestion that the literature in this area is particularly diffuse, we follow the methods of Dominguez (1998) and complement her approach in a number of ways. Firstly, we use actual intervention data, which has only recently become available from the central banks involved in intervention. This is instead of the reported intervention data used in Dominguez (1998). Using actual data will certainly provide different, and indeed may provide additional, information regarding the influence of intervention.

Also, we consider the third IMF Guiding Principle, which states that countries should take into account the exchange rate interests of others. We do this by examining the spillover effects of intervention, using UK bilateral rates. The existence of such policy spillovers have been found elsewhere in the literature in International Economics (see Cooper, 1968 and Hughes Hallett, 1989) and are a fundamental reason why central banks should coordinate their activities. Additionally, we assess the impact of coordinated intervention and contrast its effect with unilateral intervention. This is interesting because our sample period coincides with the Plaza Agreement when G-5 countries announced they stood "ready to cooperate more closely...when to do so would be most helpful". The impact of intervention on the level of the exchange rate has been found to be dependent upon whether other countries are in the market or not

(Dominguez, 1990 and Catte et al., 1994). Our approach facilitates capturing this effect and to the best of our knowledge this has not been investigated previously. Also, it has been suggested that any evidence of policy spillovers may be explained by the existence of concerted activity by more than one central bank. For example, the impact of Bundesbank intervention on the Dollar-Yen bilateral exchange rate may be due to German intervention coinciding with the intervention activities of the US and Japan. Our approach can shed some light on whether policy externalities are masked by concerted intervention.

The rest of Chapter Five is set out as follows. Section 5.2 reviews the basic theory as to why foreign exchange intervention may have an effect on the exchange rate. We also review recent empirical studies into the effect of intervention on volatility. Section 5.3 describes the GARCH methodology and the data that we use in this study. Section 5.4 presents our results for the impact of US and German foreign exchange activities on the volatility of US Dollar, Deutsche Mark and UK Pound Sterling bilateral rates. Section 5.5 concludes.

In general terms we find consistent evidence that Federal Reserve intervention reduces exchange rate volatility. Additionally, we find evidence that the distinction between unilateral and coordinated intervention is important. Furthermore we find spillover effects from foreign exchange intervention that are not dependent upon coordinated activity.

5.2 A Brief Literature Overview

5.2.1 Theoretical Issues

Before we examine the literature on foreign exchange intervention and exchange rate volatility, it is useful to consider the theoretical channels by which sterilized and non-sterilized intervention effect the level of the exchange rate. It is generally accepted that changes in the relative money supply of two countries will lead to proportionate changes in the nominal level of their bilateral exchange rate; see, for example, the Monetary Approach to exchange rate determination.[41] Since non-sterilized foreign exchange intervention is a central bank operation that involves a change in the money base, it should effect the exchange rate to the extent that there is a change in relative money supplies. In contrast, sterilized intervention combines a foreign exchange market operation with an open market operation to leave the domestic money base unchanged. According to the Monetary Approach this kind of central bank operation has no effect on the exchange rate. This is an important issue given that the US and German monetary authorities sterilize intervention as a matter of routine, leaving domestic monetary aggregates unaffected. However, there are alternative models to the Monetary Approach which predict that sterilized intervention will have an effect upon the exchange rate.

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[41] Although the Monetary Approach has been found wanting, by for example Meese and Rogoff (1983), there is recent evidence in its favour (see MacDonald, 1995, and Mark, 1995). Irrespective of whether the Monetary Approach holds empirically, it is not controversial to suggest that changes in the money supply have an impact on the nominal exchange rate.

The hypothesized channels through which sterilized intervention has an effect on the exchange rate are the Portfolio Balance and the Signaling Channels. In contrast to the Monetary Model, the Portfolio Balance Model of exchange rate determination assumes domestic and foreign assets are imperfect substitutes.[42] Asset holders will allocate their portfolios to balance exchange rate risk against expected rates of return. We can see by defining the risk premium as the expected deviation from uncovered interest-rate parity, or as a function of relative asset supplies.

$$\rho = i - i^* + E(s_{t+1}) - s = f(B / B^*) \qquad (5.1)$$

Central bank foreign exchange activities which are sterilized will change the relative supply of domestic (B) and foreign assets denominated in domestic currency units (B^*), without changing the money supply and relative interest rates (i - i^*). This will result in a change in the risk premium (ρ) and hence a change in expected returns on the assets. A change in expected return will lead to a change in the exchange rate today such that the assets are willingly held.

The Portfolio Balance and the Monetary Approach have different predictions regarding the effect of foreign exchange intervention. Nevertheless, both suggest that intervention can have an indirect effect by providing information about the views and intentions of the monetary authorities. This indirect effect is called the Signaling Channel and was first described by Mussa (1981). Unlike the Portfolio Balance Model,

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[42] Additionally bonds have to be considered as outside assets. This is to say, Ricardian Equivalence does not hold or the public fully anticipates that current government debt will

domestic and foreign assets do not have to be imperfect assets for intervention to influence the exchange rate through the Signaling Channel.[43] The central bank is assumed to convey inside information to the market about the course of future fundamental determinates of the exchange rate, by selling or buying domestic currency. Where the current exchange rate is defined within the context of a simple Monetary Model with Rational Expectations:

$$s_t = (1-\beta)\sum_{j=o}^{\infty} \beta^j E_t z_{t+j} \qquad (5.2)$$

z_{t+j} are fundamental variables including relative supplies and β is a discounting factor. For example, a central bank that intends to contract the money supply in the future - in order to lower expectations of inflation - may signal this intention by buying domestic currency today. If intervention signals a change in future monetary policy, even without a change in the money supply today, expectations of the course of monetary policy will change. As the exchange rate is determined as an asset price, a revision of market expectations of future money supply, will lead to a revision of the expected exchange rate and consequently of the current rate. By staking their money on their expectations, intervention is being used by the central

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be sterilised by future taxation. See Stockman (1979).

[43] The failure of Ricardian Equivalence is also not required for intervention to have any impact through the signalling channel.

bank as a commitment technology and will have more of an impact than 'cheap talk'.[44]

5.2.2 Empirical Evidence

Studies of the impact of foreign exchange intervention on daily exchange rate volatility include Baillie and Humpage (1992), Baillie and Osterberg (1997), Bonser-Neal and Tanner (1996), Connolly and Taylor (1994), Dominguez (1998), Hung (1997) and Mundaca (1990). Among the different measures of volatility used (including Autoregressive Conditional Heteroskedastic models, Implied Volatility and Standard Deviation) and the different measures of intervention (reported, secret and actual, buying and selling Dollars) there is inconsistent evidence that intervention influences the volatility of daily exchange rates. This could be due to the fact that intervention itself is situational: it depends on other factors that vary in the economy to have any effect and also depends upon the varying objectives of policy makers. We focus on research on the sample period January 1985 to February 1987 when there was active intervention to reduce the value of the US Dollar, which was considered seriously misaligned against the main non-Dollar currencies (Krugman, 1985). For this sample period there is a consensus that Fed intervention had a negative effect on Dollar-Mark daily volatility but less evidence that Bundesbank intervention had any effect.

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[44] For a survey of empirical research into the impact of sterilised intervention through the Portfolio Balance and Signaling Channel see Edison (1993) and Dominguez and Frankel (1993a).

Dominguez (1998) concentrates on intervention data derived from press reports and uses Generalised ARCH methods to measure *ex post* volatility. She finds that US Federal Reserve intervention had a significantly negative effect on the Dollar-Mark volatility at the 10% significance level, for the period January 1985 to February 1987. Reported Deutsche Bundesbank intervention has a significant negative effect at the 5% level. Huang (1997) investigates the impact of Federal Reserve intervention in Deutsche Marks on the Dollar-Mark exchange rate using standard deviations for a sub-period of Dominguez's sample (April 1985 - December 1986). Huang's results are consistent with Dominguez (1998): Fed DM intervention had a significant negative impact. Huang does not examine the impact of Bundesbank intervention. Bonser-Neal and Tanner (1996) use implied volatility to measure *ex ante* volatility, with reported and actual intervention data. They find a significant negative effect from reported Fed intervention but only an insignificant effect for actual Fed, and actual and reported Bundesbank intervention, over the sample period January 1st 1985 to February 22nd 1987. Baillie and Osterberg (1997), using GARCH methods, find no significant effect from Fed or Bundesbank activities. These researchers differentiate intervention into buying or selling Dollars, whilst using actual data for the sample period April 1985 to December 1986.

We follow the methods of Dominguez (1998), but complement her approach in a number of ways. First, we use actual intervention data, which has only recently become available from those central banks involved in intervention. This is instead of the reported intervention data used by Dominguez (1998) based on newspaper reports.[45] Osterberg and Westmore Humes (1993) and Klein (1993), suggest that newspaper reports often mention foreign exchange intervention that did not actually occur or fail to note intervention that did occur. In addition, there is a tendency for larger daily Dollar interventions to be reported by newspapers, whereas interventions of a smaller magnitude are unnoticed. This would suggest that using actual intervention data may present different results, and hence additional information, regarding the impact of central bank exchange rate activities. There is also the issue of the limited occasions when the Federal Reserve entered the market during our sub-period. We focus below on issues related to coordinated intervention and hence this is a means of making the most of the limited days of intervention.

We also differentiate our study by considering the Third Guiding Principle of the IMF Executive Board, that countries should take into account the exchange rate interests of others. We do this by examining the spillover effects of Federal Reserve and Bundesbank intervention, using the volatility of UK Pound bilateral exchange rates. As far as we are aware this rate has not been tested before in the literature. The possibility of policy spillovers has been considered in passing by some researchers. For instance, Bonser-Neal and Tanner (1996) find some evidence that

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[45] Dominguez (1998) additionally uses newswire reports available from NEXIS which increases the correlation between reported and actual data.

reported[46] Bank of Japan intervention is significantly associated with an increase in the *ex-ante* volatility of the Dollar-Mark rates (over the sample period 1987 to 1989).[47] Also Dominguez (1998) finds evidence that reported Bundesbank intervention significantly reduced Dollar-Yen *ex-post* volatility (during the period 1985-1987),[48] although there is no evidence that actual Bundesbank intervention had a spillover effect for the Dollar-Yen over the longer sample 1985-1991. We consider these spillover effects for Fed intervention on the Pound-Mark bilateral rate, and the effect of Bundesbank intervention on the Dollar-Pound rate. Both the rates in this instance would not appear to be of direct interest to the policy authorities under examination. There will be an effect only to the extent that there are policy spillovers.

We also consider the impact of coordinated intervention on exchange rate volatility and contrast these results with unilateral intervention. According to the Signaling Channel hypothesis the efficacy of unilateral intervention rests critically on the central bank having inside information about future monetary policy and an incentive to reveal this information truthfully. Central banks coordinate intervention activities to convince the market that both conditions hold. Multiple signals will

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[46] The Bank of Japan does not release daily foreign exchange intervention data to the public.

[47] The point is made by Bonser-Neal and Tanner (1996) that these results could be determined by multicollinearity, given that there was extensive coordinated intervention in this period. We further consider these issues below.

[48] This is rationalised by Sylvester Eijffinger as giving information to the markets about the Bundesbank's 'Dollar target'.

increase the total amount of inside information and it will also increase the probability that the information is true. The cost of lost reputation among coordinated central banks may further restrain the various monetary authorities from sending misleading signals. Additionally, coordination rules out the possibility of contradictory signals and also allows some central banks to free ride on the credibility of others. The latter may be considered as a commitment technology. Not only is the domestic central bank staking cash on intervention (selling Dollars in anticipation that the currency depreciates) the foreign central bank will be at a loss if policy actions did not eventually back up the signal. These are the reasons for using coordinated intervention but what of the evidence? Dominguez (1990) finds coordinated intervention operations over the period 1985-87 consistently influenced long-term market expectations, and therefore, on the basis of the signaling channel, lead to changes in the level of the exchange rate. This would suggest that coordinated intervention could have an impact upon exchange rate volatility quite distinct from any unilateral impact.[49] Interestingly, the distinction between unilateral and coordinated intervention may additionally allow us to flesh out any spillover effects. This non-cooperative/cooperative distinction has often been used to explain previous empirical evidence of policy externalities. For example,

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[49] Also Loopesko (1984), Eijffinger and Gruijters (1992) and Dominguez and Frankel (1993a) find evidence suggesting coordinated intervention has a quantitatively different effect from non coordinated intervention. Additionally Catte et al. (1994) found evidence that coordinated intervention had an important influence on the exchange rate. However, Humpage (1989) and Humpage and Osterberg (1992) find the distinction unimportant for the nominal exchange rate.

where Bundesbank intervention is found to be associated with increased Dollar-Yen volatility, this is attributed to coordinated intervention with the US and Japanese monetary authorities rather than any direct policy externality. By separating these two effects into different variables we hope to get a handle on the spillover effect which is not due to coordinated activities.

We therefore adopt three different models when examining exchange rate volatility: a basic model without intervention variables, an actual model which combines all of a countries intervention activities within a single variable and a marginal model which allows us to differentiate between unilateral intervention and coordinated intervention.

We also pay some attention to the effect of aberrant observations on the estimated residuals. Franses and van Dijk (1997) and Franses and Ghijsels (1999) suggest that evidence of GARCH effects may be susceptible to extreme values in the standardised residuals of our estimated results. As we further explained below, our methods partly take account of extreme values in the returns data given that we use a student-t distribution. An alternative approach considered by Dominguez (1998) is to include dummy variables, which take account of exchange rate policy announcements that could influence the returns data. There has also been an attempt to include the effect of macroeconomic announcements (see Bonser-Neal and Tanner, 1996). Nevertheless, we also consider whether our results are robust to large standardised residuals, which may not be fully accounted for by other methods.

5.3 Estimation Methods and Data Series

5.3.2 GARCH Methodology

Empirical evidence suggests that short-run exchange rate movements display contiguous periods of quiescence and turbulence, with leptokurtic (relatively peaked and fat tailed) unconditional distributions; see, for example, Mandelbrot (1963), Westerfield (1977), Mussa (1979), and Hsieh (1988). These movements are ideally suited to modeling by AutoRegressive Conditional Heteroskedastic (ARCH) methods and this is the framework we use to model exchange rate volatility. In this section we firstly introduce Engle's (1982) original ARCH model before considering the more parsimonious Generalised ARCH (GARCH) model introduced by Bollerslev (1986).

Following the seminal paper by Engle (1982) we shall refer to all discrete time stochastic processes (ε_t) of the form

$$\varepsilon_t = z_t \sqrt{v_t} \qquad (5.3)$$

$$z_t \; iid, \quad E(z_t) = 0, \quad \text{var}(z_t) = 1, \qquad (5.4)$$

with $\sqrt{v_t}$ a time-varying, positive, measurable function of the time t-1 information set, as an ARCH model. By definition ε_t is serially uncorrelated with mean zero, but the conditional variance of ε_t equals v_t, which may be changing through time. In most applications, and in ours, ε_t refers to the innovation in the mean for some other stochastic process, say $\{y_t\}$ where

$$y_t = g(x_{t-1}; \beta) + \varepsilon_t \qquad (5.5)$$

and $g(x_{t-1};\beta)$ denotes a function of x_{t-1} and the parameter vector β, where x_{t-1} is in the time t-1 information set.

Let $f(z_t)$ denote the density function for z_t, and θ be the vector of all the unknown parameters in the model. By the prediction error decomposition, the log-likelihood function for the sample ε_T, $\varepsilon_{T-1},\ldots,\varepsilon_1$ becomes, apart from the initial conditions,

$$L(\theta)=\sum_{t=1}^{T}\left[\log f(\varepsilon_t v_t^{-1/2})-\log\sqrt{v_t}\right] \qquad (5.6)$$

The second term in the summation is the Jacobian term arising from the transformation from z_t to ε_t. Note that equation *(5.6)* also defines the sample log-likelihood for y_T, $y_{T-1},\ldots$, y_1 as given by *(5.5)*. Given a parametric representation for $f(z_t)$, maximum likelihood estimates for the parameters of interest can be computed directly from *(5.6)* by a number of different numerical optimization techniques. In our example we estimated *(5.6)* using the maximum likelihood procedure described in Berndt, *et al.* (1974).

As suggested by Engle (1982) one possible parameterisation for v_t is to express v_t as a linear function of past squared values of the process

$$v_t=\alpha_o+\sum_{i=1}^{q}\alpha_i\varepsilon_{t-1}^2 \qquad (5.7)$$

with α_o>0 and $\alpha_i \geq 0$. This model is known as the linear ARCH model. With financial and exchange rate data it captures the tendency for volatility clustering, i.e. for large (small) price changes to be followed by other large (small) changes, but of unpredictable sign.

In many of the applications with the linear ARCH(q) model, a lag length of q is required. An alternative and more flexible lag structure is often provided by the Generalised ARCH or GARCH(p,q) model in Bollerslev (1986),

$$v_t = \alpha_0 + \sum_{i=1}^{q} \alpha_i \varepsilon_{t-i}^2 + \sum_{i=1}^{p} \beta_i v_{t-i} \tag{5.8}$$

To ensure a well-defined process all the parameters in the infinite order AR representation must be non-negative, where it is assumed that the roots of the polynomial lie outside the unit circle. For a GARCH(1,1) process this amounts to ensuring that both α_1 and β_1 are non-negative. It follows also that ε_t is covariance stationary if and only if $\alpha_1+\beta_1<1$. Of course in that situation the GARCH(p,q) model corresponds exactly to an infinite order linear ARCH model with geometrically declining parameters.

Bollerslev *et al.* (1992) suggest that in most applications a lag length of p=q=1 will suffice. In addition, it has been found by Hsieh (1989a,b) that a simple GARCH(1,1) model did relatively well in describing the returns to five different daily nominal US Dollar rates and this is the model which we utilise in this study. These methods give us a measure of exchange rate volatility, in the form of a conditional variance, and an idea of its statistical relationship to foreign exchange intervention. In the conditional variance equation *(5.9)* we have foreign exchange interventions variables, a holiday dummy and a measure of the spread between domestic and foreign interest rates.

The GARCH(1,1) model that we use is as follows:

$$\Delta s_t = \beta_0 + \beta_1 Fed_{t-1} + \beta_2 Buba_{t-1} + \beta_3 H_t + \beta_4 SPREAD_t + \varepsilon_t \tag{5.9}$$

$$\varepsilon_t \mid \Omega_{t-1} \sim N(0, v_t, \kappa_t)$$

$$v_t = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \alpha_2 v_{t-1} + \psi_1 |Fed_{t-1}| + \psi_2 |Buba_{t-1}| + \psi_3 H_t + \psi_4 SPREAD_t \tag{5.10}$$

Where $\Delta s_t = 100 * \log(S_t/S_{t-1})$ is the return on the spot exchange rate (S_t) between period *t* and *t*-1. H_t is a holiday dummy variable, which is equal to one the day following the market being closed for a public holiday.[50] *Fed* is a variable capturing actual US Federal Reserve intervention operations in billions of US Dollars. The *Fed* intervention variable has two forms: the first includes Federal Reserve intervention in all currencies; the second consists of Fed intervention in Deutsche Marks only. Presumably, there will be a different effect from intervention conducted in Yen from that conducted in Marks for say the Dollar-Mark and the Pound-Mark rates. *Buba* is a variable capturing Deutsche Bundesbank intervention, again in billions of US Dollars.[51] For the conditional mean equation *(5.9)* positive values for *Fed* and *Buba* indicate Dollar purchases by the central bank, while negative values denote Dollar sales. || is the absolute value operator and ε_t is the disturbance term. The conditional distribution of the disturbance term is standardised t, with variance v_t and degrees of freedom κ. This is based on Bollerslev (1986), Hsieh (1989b), and Baillie and Bollerslev (1989), who find evidence that a

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[50] Hsieh (1988) finds evidence that holiday dummy variables should be included as explanatory variables in daily exchange rate GARCH models. We include holidays from both countries that make up the bilateral rates as a single variable.

[51] Federal Reserve and Bundesbank intervention are both lagged by one time period to ensure the intervention variables are predetermined as our exchange rate is 12:00 EST.

conditional student t distribution performs better than a normal distribution for daily exchange rate volatility. For example, exchange rate returns data is typically more leptokurtic than the normal distribution. The t distribution approaches a normal distribution as the parameter κ approaches infinity.

Following Dominguez (1998) we include the spread between domestic and foreign interest rates (*SPREAD*). For example, we include the spread between German and UK interest rates for Mark-Pound spot returns, German and US rates for Dollar-Mark returns, and UK and US interest rates for Dollar-Pound. This is an attempt to control for the impact of changes in relative monetary policy which, according to the Monetary Model, will lead to changes in the exchange rate. Our volatility model will potentially suffer from omitted variable bias if we do not take account of contemporaneous monetary policy or a change in the money supply.

We consider three specifications for the GARCH(1,1) model. The first is the basic model that excludes intervention variables. The second model includes our intervention variables using actual central bank data: this is our actual model. The third specification that we use takes account of contemporaneous intervention by our two central banks and separates this from independent, or marginal, intervention. We therefore have marginal variables for independent intervention *Fed* and *Buba*, and an added variable *Coord* that represents coordinated intervention. This is our marginal model, with the following conditional variance equation:

$$\begin{aligned} v_t = \alpha_0 + \alpha_1\varepsilon_{t-1}^2 + \alpha_2 v_{t-1} + \psi_1|Fed_{t-1}| + \psi_2|Buba_{t-1}| + \psi_3|Coord_{t-1}| \\ + \psi_4 H_t + \psi_5 SPREAD_t \end{aligned} \tag{5.11}$$

5.3.2 Data Series

The exchange rate data is 12:00 noon EST and the interest rate data is overnight 4:00 EST data. Both are from the Federal Reserve Bank of New York. German intervention data is from the Bundesbank and US data is from Chris Neely. Our sample period is from the 2nd January 1985 to 27th February 1987, which spans the period over which the monetary authorities actively engaged in the foreign exchange market in an attempt to reduce the value of the Dollar.

5.4 An Empirical Study of Volatility

5.4.1 Data Analysis and Preliminary Results

Firstly, we visually examine the returns data or, in other words, the first difference of the log of the exchange rate. From Figure 5.1 we see that these data series are characterised by contiguous periods of volatility and stability. This suggests that these data can be usefully represented within a GARCH framework. The graphs are roughly similar for the bilateral US Dollar rate against the Deutsche Mark and UK Pound. This similarity should be reflected in our GARCH results. The graph for the returns to the DM-Pound rate takes a more distinct pattern and we are particularly interested in whether our measures of intervention bear any relationship to these patterns of volatility.

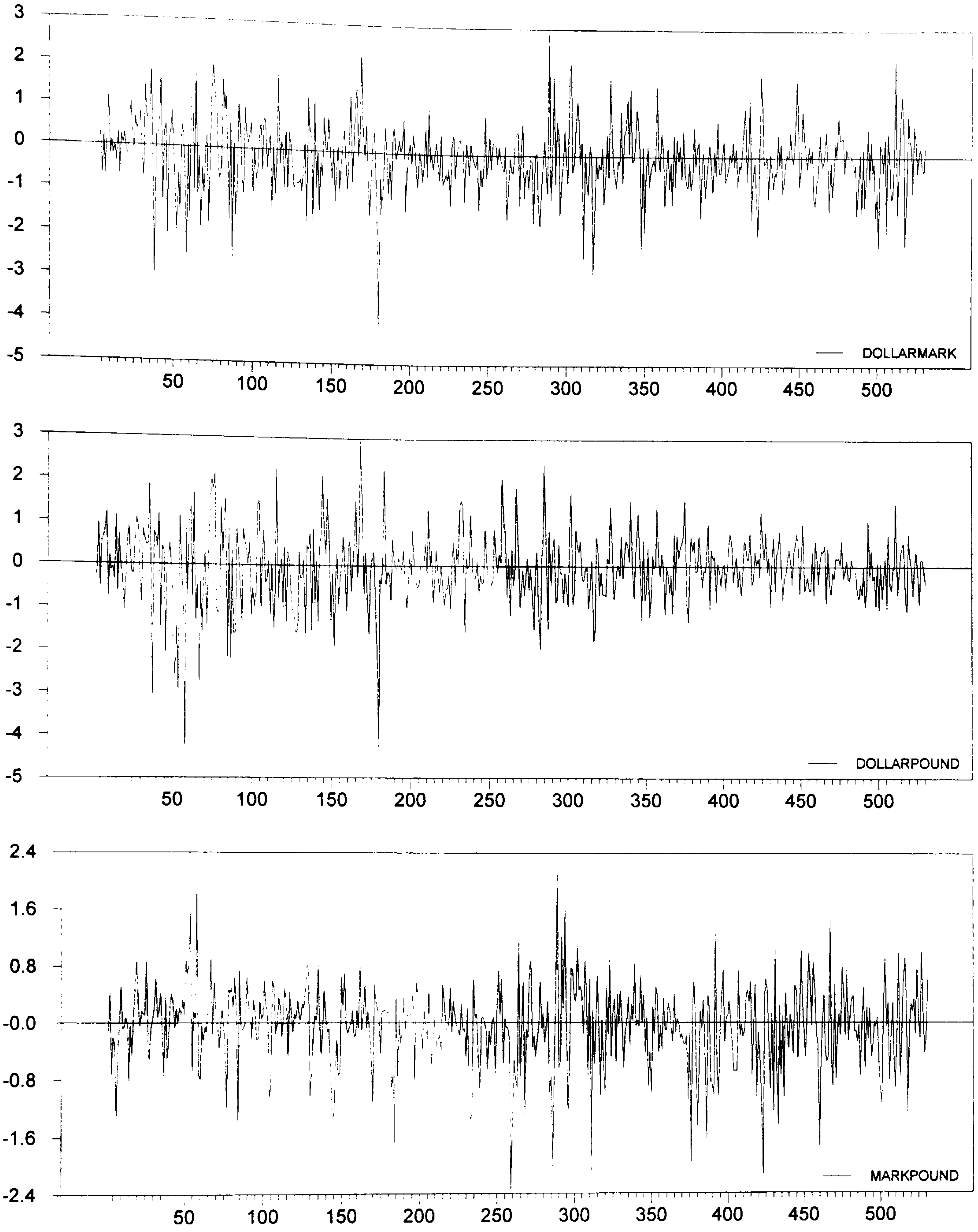

Figure 5.1 The first difference of the logs of the US Dollar- Deutsche Mark, US Dollar-UK Pound and Deutsche Mark-UK Pound bilateral exchange rates

We include in our estimation results the value of the log likelihood function, the number of iterations, the Box-Pierce Q statistic for the standardised residuals ($Q_z(20)$) and the squared standardised residuals ($Q_z^2(20)$). These give an impression of how well our model is specified. Given that in Tables 5.1 to 5.3 none of our Box-Pierce Q statistics are significant, our results do not appear to be misspecified in this sense.

Additionally, our results from Tables 5.1, 5.2 and 5.3 (in Appendix 5.A) suggest that the holiday dummies (H_t) generally have a positive impact on the Dollar-Mark and Dollar-Pound rate, a negative impact for the Pound-Mark, although it is typically insignificant. In some occasions it was found that removing the holiday dummy from the conditional variance equation would have the effect of ensuring that the intercept coefficient was positive without having a qualitative impact on the other coefficients. For comparative purposes therefore we have retained the holiday dummy. The degrees of freedom parameter was consistently significant and between 4 and 8 in magnitude. This result is consistent with the use of the conditional student-t distribution.

The coefficients on the lagged squared residuals (α_1) and lagged conditional variance (α_2), in the conditional variance equations, are always found to be positive and their sum is less than, but close to, one. This indicates that shocks to exchange rate volatility do not die out very quickly. Lamoureux and Lastrapes (1990) suggest that substantial volatility persistence represents misspecification and may be due to structural change in the unconditional variance of the process, as represented by a change in α_0. Lamoureux and Lastrapes (1990) deal with this problem by introducing dummy variables that take account of changing policy regimes. Our results should be robust to this problem since we have a sample period when exchange rate policy did not change.

5.4.2 Exchange Rate Volatility and Intervention

Table 5.1 in Appendix 5.A examines the relationship between the volatility of the daily US Dollar-Deutsche Mark rate and our two countries' intervention variables (*Fed*, *Buba* and *Coord*). The coefficient on *Fed* intervention is significant and negatively associated with *ex post* volatility. This is irrespective of whether we use total Federal Reserve intervention (as in the actual model) or unilateral intervention (as represented by the marginal model). Also, this significant and negative association is not conditional upon whether the data is Fed intervention in all currencies or intervention conducted in DM only.

German intervention has a positive effect on volatility, although nowhere in Table 5.1 is it significant at the 5% level. There is some indication that it is significant at the 10% level but this is dependent upon Fed intervention being conducted in all currencies. We would generally expect the *Buba* intervention variable's coefficient to change between the two actual specifications where there was some relationship between *Buba* and *Fed* variables. Since intervention was often coordinated during this period we should not be surprised by this result. Hence the change in significance of the *Buba* coefficient is an indication of collinearity between our two countries' intervention variables, and is *prima facie* evidence that we should differentiate between our actual and marginal models. Further evidence for this differentiation is provided by the log likelihood test statistic. The value of the function is greatest for the marginal model, and following Taylor (1994), we consider this further evidence in favour of our distinction between actual, unilateral and coordinated intervention. The coefficient on our variable for coordinated intervention (*Coord*) does not have a significant effect on volatility at the 10% level. Interestingly, this is a marginal failure to indicate that *Coord* is significant (with t-values equal to 1.63 and 1.48), much in the way that *Buba* fails marginally but classically to be accepted in the actual model. Although this would lead us to suspect that there is still an effect present, we nevertheless fail to find any clear statistical evidence in favour of this possibility for the US Dollar-DM exchange rate.

Next we examine the relationship between foreign exchange intervention and the *ex post* volatility of the UK Pound-US Dollar bilateral rate. This rate is of direct concern for US policy makers.[52] Indeed, in Table 5.2, we find a significant and negative association between all measures of Fed intervention and volatility. On the other hand, given that this bilateral rate is of no direct interest to German policy makers, Bundesbank intervention is unlikely to have a significant effect. So our results from the actual model which suggests that Bundesbank intervention produces a

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[52] Also this is of concern for UK policy makers. Unfortunately, as mentioned above, the Bank of England does not permit access to daily intervention data.

strong positive effect on this rate, not directly related to German policy, appear to be counter-intuitive. Alternatively, this could be due to strong spillovers from Bundesbank intervention, as found by Dominguez (1998) on the Dollar-Yen rate.

Again in Table 5.2 our log-likelihood function suggests that our preferred specification is the marginal model which takes account of the interaction between the two countries intervention activities. Once we separate Bundesbank intervention into unilateral and coordinated intervention we find that there is no spillover from non-coordinated Bundesbank intervention. There remains some indication from the *Coord* variable that when Germany is in the market with the US, Federal Reserve intervention has a strongly significant and positive impact upon the volatility of the Pound-Dollar rate. This is contrary to the impact of typical Fed intervention and not inconsistent with the results from Table 5.1, regarding the Mark-Dollar rate.

It is important to note that evidence of coordinated intervention being positively associated with volatility could be a result of the nature of G-5 intervention at the time. If a number of countries were attempting to induce large negative changes in the Dollar bilateral rates, we should not be surprised if coordinated intervention was able to increase volatility. With this view, what is disappointing is Fed intervention does not facilitate changes in the bilateral rates of the same order. Coordinated intervention has a more powerful impact than unilateral intervention, in the sense that it is associated with large movements in the exchange rate. This coordinated intervention was also characterised by UK involvement, but any further examination of UK intervention is constrained by data limitations. To summarise the results so far: we find a significant and negative effect for Fed intervention across all specifications for the Pound-Dollar rate as we do for the Mark-Dollar rate. Compared to the Mark-Dollar rate we find more evidence that Bundesbank intervention is important, although these occasions seem constrained to whether the Fed is also in the market. This provides minimal evidence of policy spillovers but greater evidence that the non-coordinated/coordinated intervention distinction is important.

Thirdly, we focus on the spillovers from US Federal Reserve intervention and the volatility of the UK Pound-Deutsche Mark rate. This rate should be of no direct interest to the US Federal Reserve and little direct concern for Germany, given that for our mid-1980s sample Germany and Britain had no formal exchange rate arrangement.[53] In Table 5.3 there is strong evidence that Federal Reserve policy is having an effect to the extent that its intervention is unilateral and conducted in DM. Bundesbank foreign exchange operations have a significant negative effect when the intervention is unilateral. This particular specification is preferred on the basis of the size of the log likelihood function. For this rate we do not find any detrimental spillovers, only benign externalities. The results for the Pound-Dollar rate contrast somewhat with those for the Pound-Mark rate,

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[53] See Lawson (1992). The UK policy of shadowing the DM began around the Louvre Accord of February 1997, after our sample period.

although marginal German intervention has more of a negative effect on exchange rate volatility than actual intervention.

From Table 5.3 we see that coordinated intervention does not appear to have a statistically significant effect on the Pound-Mark rate. The coefficient on this intervention variable is also rather small. This can quite sensibly be interpreted as further evidence of the success of coordinated intervention, in terms of its objectives of the time. G-5 intervention attempted to reduce the level of the overvalued Dollar: if both currencies (Deutsche Mark and UK Pound) benefit approximately equivalently then it is sensible that the volatility of the cross rate is unaffected. The equivalent benefit of coordinated intervention on the volatility of Dollar bilateral rates with the DM and Pound from Table 5.1 and Table 5.2 are consistent with this argument.

5.4.3 Robustness of Results

We also consider the impact of outliers on our results. There are reasons to believe that our analysis should already take account of this problem to a certain extent. For instance, we use the leptokurtic student-t distribution to take account of fat tails in the exchange rate returns data. Additionally, and consistent with the point made by Bayoumi and Eichengreen (p197, 1998), extreme changes in the exchange rate and in reserves - induced in our study by intervention - is a inherent characteristic of the data used here. To leave out these observations would be to lose the kind of information we are especially interested. Following the approaches of Franses (1998) and Nelson (1990), we include dummies for standardised residuals that are outside the bandwidth [-4, +4]. Our results are not qualitatively different from those not corrected for aberrant observations, standardised residuals are within our bandwidth and the results do not suffer from serial correlation.

Table 5.4 in Appendix 5.B provides evidence that Federal Reserve intervention consistently reduces DM-Dollar volatility across our different measures of intervention and specifications even when we take account of outliers. Generally, Bundesbank and coordinated intervention are not related to volatility. Table 5.5 suggests Fed intervention is associated with a reduction in volatility and actual Buba intervention and coordinated intervention increase the volatility of the Pound-Dollar exchange rate. There is some evidence that Fed intervention reduces volatility for the Pound-DM rate in Table 5.6 and that unilateral German intervention also reduces volatility taking account of outliers.

5.5 Conclusion

In this Chapter we utilise a GARCH methodology to consider the relationship between exchange rate volatility and foreign exchange intervention. We extend the previous literature in this area in two important ways. Firstly, using actual intervention data we consider joint activity by the Bundesbank and Federal Reserve; making a distinction between actual, unilateral and coordinated intervention. Secondly, we examine the Pound-DM and Pound-Dollar rates, and, in particular, whether there is any evidence of policy externalities to these rates from German Bundesbank and US Federal Reserve intervention.

Our results consistently provide evidence that unilateral Federal Reserve intervention is associated with a reduction in DM-Dollar and Pound-Dollar daily exchange rate volatility, and with benign spillovers to the Pound-DM rate. In contrast, there is an indication that Bundesbank intervention increased volatility for the DM-Dollar and Pound-DM bilateral rates. Additionally there is evidence of spillovers from Buba activity to the DM-Dollar rate. There is a potential problem with collinearity given the fact that both central banks have a tendency to engage in coordinated activities. Once we separated these effects out we find that joint German and US intervention is significantly related to Pound-Dollar volatility rate. (There is also evidence that coordinated intervention increases volatility for the Dollar-DM rate but this marginally fails to be significant.) However the evidence that intervention appears to increase volatility must be considered within the context of the objectives of Central Banks' foreign exchange activities. With a policy which attempts to induce large changes in the level of the exchange rate consistent with some medium term objective it should not be surprising that we find evidence intervention has increased volatility. Such was the nature of exchange rate policy during our sample period, when there was extensive attempts to reduce the value of the Dollar.

Our results highlight the different impact of unilateral intervention and coordinated intervention. This may be symptomatic of the existence of

John Williamson's (1985) delineation of exchange rate movements. Williamson emphasises the distinction between medium term movements and short run movements. Policy may alternate between attempts to change trend movements and reduce short run fluctuations in the exchange rate. We find evidence that G-5 activity to move the nominal exchange rate in the medium term has involved a daily 'cost' in increased short run movements. However, this additional cost has to be considered in the light of the express intention of changing the level of the exchange rate. Unilateral activity, on the other hand, has apparently been more concerned by limiting fluctuation of any kind and seems to have been more successful in this regard.

Interestingly, to the extent that the creation of the Euro has created a forum for the implementation of monetary policy, we can suggest that this is likely to provide more opportunities to coordinated intervention at a European level. This is dependent upon the underlying preferences of monetary authorities and their willingness to intervene. There has been a disinclination to consider the depreciation of the Euro against the Dollar and Yen as a problem necessitating intervention. This was mainly due to these exchange rate movements being beneficial to large areas of Euroland in a cyclical downturn. But to the extent that such swings in the exchange rate conflict with the cyclical implementation of monetary policy (in general terms) then we suggest that sterilised intervention will have a more powerful impact because of greater coordination.

A suggestion for further research is to consider Hamilton and Susmel's (1994) Switching ARCH (SWARCH) methodology which allows the parameters of an ARCH process to come from one of several different regimes. For example, Lamoureux and Lastrapes (1990) argue that substantial persistent in volatility may be due to structural change in the unconditional variance of the process or changing objectives within a particular regime. In terms of exchange rate policy that would involve switching between a medium-term exchange rate target and an attempt to ameliorate market volatility on a day-to-day basis. One approach would be to include Markov-switching parameters in the ARCH model, as suggested by Hamilton and Sunsel (1994). Our methods have attempted to pick this up by using the distinction between unilateral and coordinated intervention. To the extent that intervention is coordinated then it would be an indication that G-5 countries were trying to induce large changes whereas unilateral intervention was aimed at reducing day-to-day volatility. It would be interesting to consider whether we obtain some indication of these differences within other frameworks. Additionally, the SWARCH approach may take account of whether the effectiveness of signals conveyed through sterilised intervention is dependent on the central bank's objectives being "consistent with underlying economic fundamentals" Group of Ten Deputies (1993). This kind of argument suggests that if an exchange rate is consistent with fundamentals and policy makers objectives are clearly stated then the effect of intervention will be different from when there is misalignment and policy makers aims are unclear.

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Appendix 5.A

Table 5.1: Daily Exchange Rate GARCH(1,1) Model for Deutsche Mark-U.S. Dollar

$$v_t = \alpha_o + \alpha_1\varepsilon^2_{t-1} + \alpha_2 v_{t-1} + \psi_1|Fed_{t-1}| + \psi_2|Buba_{t-1}| + \psi_3|Coord_{t-1}| + \psi_4 H_t + \psi_5 SPREAD_t$$

	Basic	Actual[a]	Actual[b]	Marginal[a]	Marginal[b]
α_0	0.023	-0.004	0.001	0.001	0.001
	(0.898)	(-0.465)	(0.091)	(0.103)	(0.094)
α_1	0.085	0.010	0.042	0.035	0.044
	(2.468)**	(0.909)	(2.071)**	(2.183)**	(2.127)**
α_2	0.881	0.975	0.941	0.944	0.939
	(17.661)***	(71.905)***	(33.943)***	(56.277)***	(34.691)***
ψ_1		-0.956	-0.633	-2.055	-2.938
		(-3.960)***	(-3.484)**	(-5.738)***	(-6.224)***
ψ_2		0.132	0.158	0.094	0.119
		(1.726)*	(1.516)	(0.603)	(0.767)
ψ_3				0.158	0.140
				(1.630)	(1.482)
ψ_4	0.097	-0.026	0.049	0.031	0.070
	(0.716)	(-0.319)	(0.437)	(0.290)	(0.614)
ψ_5	-0.001	-0.006	-0.003	-0.005	-0.003
	(-0.099)	(-2.159)**	(-0.865)	(-1.270)	(-0.812)
ρ[c]	18	134	371	79	28
Ln(L)[d]	-629.918	-622.688	-624.347	-618.150	-617.272
$Q_z(20)$[e]	21.510	22.976	21.024	19.139	19.724
$Q_z^2(20)$[f]	19.442	25.332	20.892	21.425	20.216

Notes: Model includes holiday variable (H_t) and interest rate spread ($SPREAD_t$) in conditional mean and variance. U.S. Federal Reserve intervention is in Billions of Dollars. Returns = 100*log(s/s{1}).***=1% **=5% *=10%. Exchange rate data is noon U.S. EST. Buba intervention is in Dollars and DM. The sample period is from 2 January 1985 to 22 February 1987.

(a) Total official Fed intervention in DM, Yen and other currencies.
(b) Official Fed intervention in DM only.
(c) ρ is the number of convergence iterations.
(d) ln(L) is the value of the log likelihood function.
(e) $Q_z(20)$ denotes the Box-Pierce Q-statistic (with 20 lags) for the standardised residuals
(f) $Q_z^2(20)$ is the Box-Pierce Q-statistic for the squared standardised residuals.

Table 5.2: Daily Exchange Rate GARCH(1,1) Model for UK Pound-U.S. Dollar

$$v_t = \alpha_0 + \alpha_1 \varepsilon^2_{t-1} + \alpha_2 v_{t-1} + \psi_1|Fed_{t-1}| + \psi_2|Buba_{t-1}| + \psi_3|Coord_{t-1}| + \psi_4 H_t + \psi_5 SPREAD_t$$

	Basic	**Actual**[a]	**Actual**[b]	**Marginal**[a]	**Marginal**[b]
α_0	-0.018	-0.004	-0.004	-0.000	0.004
	(-1.118)	(-0.306)	(-0.325)	(-0.005)	(0.322)
α_1	0.034	0.014	0.015	0.009	0.005
	(2.175)**	(1.541)	(1.590)	(0.786)	(0.516)
α_2	0.951	0.983	0.983	0.976	0.983
	(48.730)***	(94.452)***	(93.128)***	(80.457)***	(114.899)***
ψ_1		-0.218	-0.393	-1.812	-2.116
		(-2.189)**	(-2.192)**	(-5.737)***	(-4.385)***
ψ_2		0.119	0.125	-0.035	-0.039
		(2.637)***	(2.730)***	(-0.381)	(-0.486)
ψ_3				0.285	0.294
				(4.582)***	(4.620)***
ψ_4	0.116	0.135	0.135	0.102	0.092
	(1.291)	(1.504)	(1.505)	(1.041)	(1.019)
ψ_5	0.006	-0.001	-0.001	0.001	-0.001
	(1.377)	(-0.198)	(-0.194)	(0.234)	(-0.211)
ρ^c	37	123	171	865	38
$Ln(L)^d$	-626.751	-620.737	-620.855	-617.036	-615.482
$Q_z(20)^e$	17.218	16.743	16.753	19.298	18.323
$Q_z^2(20)^f$	23.223	26.689	26.778	28.035	29.550

Notes: See Table 5.1.

Table 5.3: Daily Exchange Rate GARCH(1,1) Model for UK Pound-Deutsche Mark

$$v_t = \alpha_0 + \alpha_1 \varepsilon^2_{t-1} + \alpha_2 v_{t-1} + \psi_1 |Fed_{t-1}| + \psi_2 |Buba_{t-1}| + \psi_3 |Coord_{t-1}| + \psi_4 H_t + \psi_5 SPREAD_t$$

	Basic	**Actual^a**	**Actual^b**	**Marginal^a**	**Marginal^b**
α_0	0.131	0.163	0.171	0.099	0.095
	(1.718)*	(1.873)*	(1.855)*	(1.787)*	(1.614)
α_1	0.210	0.213	0.227	0.171	0.176
	(2.638)***	(2.583)***	(2.573)***	(2.493)**	(2.564)**
α_2	0.639	0.599	0.585	0.664	0.672
	(5.689)***	(5.078)***	(4.798)***	(6.515)***	(7.119)***
ψ_1		-0.601	-0.194	-0.577	-0.724
		(-1.335)	(-1.168)	(-1.334)	(-6.269)***
ψ_2		-0.024	-0.020	-0.244	-0.240
		(-0.250)	(-0.192)	(-3.996)***	(-3.366)***
ψ_3				-0.044	-0.035
				(-1.322)	(-0.829)
ψ_4	-0.037	-0.054	-0.049	-0.054	-0.051
	(-0.514)	(-0.742)	(-0.649)	(-0.828)	(-0.778)
ψ_5	-0.010	-0.012	-0.012	-0.004	-0.004
	(-1.127)	(-1.227)	(-1.227)	(-0.657)	(-0.634)
ρ^c	13	91	1036	1150	770
ln(L)^d	-433.782	-432.598	-430.481	-424.120	-423.519
$Q_z(20)^e$	23.980	23.349	23.359	20.076	20.029
$Q_z^2(20)^f$	4.679	4.724	4.703	7.088	7.501

*Notes: S*ee Table 5.1.

Appendix 5.B

Table 5.4: Daily Exchange Rate GARCH(1,1) Model for Deutsche Mark - US Dollar Corrected for Aberrant Standardised Residuals

$$v_t = \alpha_0 + \alpha_1\varepsilon^2_{t-1} + \alpha_2 v_{t-1} + \psi_1|Fed_{t-1}| + \psi_2|Buba_{t-1}| + \psi_3|Coord_{t-1}| + \psi_4 H_t + \psi_5 SPREAD_t$$

	Basic	**Actual**[a]	**Actual**[b]	**Marginal**[a]	**Marginal**[b]
α_0	0.039	0.012	0.010	0.0221	0.017
	(1.372)	(0.743)	(0.701)	(0.928)	(0.881)
α_1	0.087	0.053	0.043	0.068	0.057
	(2.493)**	(2.156)**	(1.926)*	(2.244)**	(2.137)
α_2	0.856	0.910	0.925	0.876	0.902
	(17.387)***	(30.244)***	(32.394)***	(22.430)***	(27.428)***
ψ_1		-0.496	-0.758	-2.026	-3.526
		(-2.560)**	(-2.939)***	(-2.452)**	(-3.573)***
ψ_2		0.172	0.189	0.161	0.169
		(1.438)	(1.725)*	(0.701)	(0.878)
ψ_3				0.125	0.160
				(0.828)	(1.354)
ψ_4	0.1216	0.099	0.092	0.103	0.126
	(0.897)	(0.858)	(0.831)	(0.800)	(1.057)
ψ_5	0.002	-0.001	-0.001	-0.003	-0.000
	(0.327)	(-0.303)	(-0.195)	(-0.407)	(-0.024)
ρ^c	39	22	20	47	71
Ln(L)[d]	-627.326	-621.195	-621.614	-619.706	-616.813
$Q_z(20)^e$	17.505	15.818	16.738	14.008	14.725
$Q_z^2(20)^f$	23.743	25.140	25.671	24.093	21.239

Notes: These results are corrected by including dummy variables for aberrant values in the standardised residuals. These are defined, following Franses (1998), as standardised residuals whose modulus value is greater than four standard errors. This approach suggests the inclusion of dummy variables for observation t=180 and 289. We include these dummy variables in both conditional mean and variance, which consequently results in the standardised residuals all being within our band [-4, +4]. For other notation see Table 5.1.

Table 5.5: Daily Exchange Rate GARCH(1,1) Model for UK Pound-US Dollar Corrected for Aberrant Standardised Residuals

$$v_t = \alpha_0 + \alpha_1 \varepsilon^2_{t-1} + \alpha_2 v_{t-1} + \psi_1 |Fed_{t-1}| + \psi_2 |Buba_{t-1}| + \psi_3 |Coord_{t-1}| + \psi_4 SPREAD_t$$

	Basic	**Actual^a**	**Actual^b**	**Marginal^a**	**Marginal^b**
α_0	0.004	0.012	0.013	0.011	0.012
	(0.374)	(1.083)	(1.168)	(0.963)	(1.151)
α_1	0.053	0.030	0.033	0.023	0.014
	(2.780)***	(2.414)**	(2.402)**	(1.542)	(1.083)
α_2	0.936	0.970	0.967	0.964	0.976
	(42.114)***	(67.331)***	(62.318)***	(60.135)***	(78.892)***
ψ_1		-0.357	-0.216	-1.581	-1.896
		(-2.296)**	(-2.506)**	(-4.232)***	(-3.279)***
ψ_2		0.129	0.132	-0.046	-0.061
		(2.402)**	(2.543)***	(-0.475)	(-0.696)
ψ_3				0.226	0.249
				(2.901)***	(2.991)***
ψ_4	0.001	-0.004	-0.003	-0.001	-0.002
	(0.180)	(-1.123)	(-1.151)	(-0.319)	(-0.638)
ρ^c	29	22	29	348	36
Ln(L)^d	-615.287	-609.241	-609.104	-605.262	-603.956
$Q_z(20)^e$	13.778	12.563	12.754	14.025	13.075
$Q_z^2(20)^f$	19.611	21.855	22.056	23.896	27.068

Notes: These results are corrected by including dummy variables for aberrant values in the standardised residuals. These are defined, following Franses (1998), as standardised residuals whose modulus value is greater than four standard errors. This approach suggests the inclusion of dummy variables for observation t=38 and 180. We include these dummy variables in the conditional mean, which consequently results in the standardised residuals all being within our band [-4, +4]. Additionally we exclude the variable for market holidays given that our results which included the holiday variable often had a negative constant in the conditional variance equation. For other notation see Table 5.1.

Table 5.6: Daily Exchange Rate GARCH(1,1) Model for UK Pound-Deutsche Mark Corrected for Aberrant Standardised Residuals

$$v_t = \alpha_0 + \alpha_1\varepsilon^2_{t-1} + \alpha_2 v_{t-1} + \psi_1|Fed_{t-1}| + \psi_2|Buba_{t-1}| + \psi_3|Coord_{t-1}| + \psi_4 H_t + \psi_5 SPREAD_t$$

	Basic	**Actual**^{a}	**Actual**^{b}	**Marginal**^{a}	**Marginal**^{b}
α_0	0.114	0.103	0.099	0.090	0.091
	(1.706)*	(1.603)	(1.593)	(1.597)	(1.621)
α_1	0.197	0.184	0.182	0.163	0.165
	(2.792)***	(2.727)**	(2.757)***	(2.600)***	(2.611)***
α_2	0.651	0.681	0.680	0.676	0.664
	(6.240)***	(6.836)***	(6.900)***	(6.814)***	(6.812)***
ψ_1		-0.018	0.003	-0.508	-0.691
		(-0.107)	(0.011)	(-1.256)	(-2.924)***
ψ_2		-0.030	-0.033	-0.202	-0.233
		(-0.468)	(-0.527)	(-2.340)**	(-3.514)***
ψ_3				-0.033	-0.037
				(-0.811)	(-1.001)
ψ_4	-0.030	-0.0366	-0.036	-0.044	-0.045
	(-0.454)	(-0.544)	(-0.565)	(-0.718)	(-0.701)
ψ_5	-0.008	-0.007	-0.006	-0.004	-0.004
	(-1.075)	(-0.969)	(-0.923)	(-0.659)	(-0.571)
ρ^c	18	28	22	58	505
Ln(L)^d	-432.569	-431.821	-431.814	-426.458	-422.972
$Q_z(20)^e$	24.892	24.067	24.204	20.929	20.195
$Q_z^2(20)^f$	8.403	8.778	8.888	10.442	11.940

Note: These results are corrected for aberrant values in the standardised residuals by including dummy variables. Aberrant observations are defined, following Franses (1998), as standardised residuals whose modulus value is greater than four standard errors. This approach suggests the inclusion of dummy variables for observation t=259. We include a dummy variable in both the conditional mean and variance, which consequently results in the standardised residuals all being within our band [-4, +4]. For other notation see Table 5.1.

Chapter Six

Conclusion and Suggestions for Further Research

In this work we have considered a number of issues related to foreign exchange reserves. This is an important area for research give recent evidence highlighting the influence of foreign exchange operations on the exchange rate. Intervention is conditional on central bank holdings of foreign currency. Hence reserves and intervention are intrinsically interdependent. Empirical research in this area of international economics has concerned two broad issues: what factors influence reserves and what impact do reserve operations have on other macroeconomic variables? In particular, our research concerns the long-run determination of reserves and the high frequency impact of foreign exchange operations on exchange rate volatility.

The traditional literature concerning reserve determination was framed with respect to balance of payments disequilibria. Reserves were held as a buffer-stock to finance external disequilibrium and to avoid the costly real economy adjustments required to restore external payments equilibrium. Early academic research was conducted in an era of fixed rates, under the Bretton Woods agreement, where no adjustment occurred through the nominal exchange rates. The collapse of Bretton Woods meant economic shocks could be ameliorated by adjustment through exchange rates and removed much of the reason for reserve holdings. Nevertheless, international reserves continued to be held throughout the 1970s, and thereafter. This can be attributed to the desire of policy makers to engage in foreign exchange intervention. Any study of the long-run determination of reserves should take account of this.

Recent developments in econometrics are another important justification for revisiting reserves and another important contribution that this thesis makes to the existing literature. In particular, we apply methods that take account of potential non-stationarity in our data and the possibility of spurious relationships. These methods have not been widely used in the literature examining reserves. Our approach uses tests for cointegration as evidence in favour of the existence of long-run relationships. We consider the short-run interaction between our variables and the speed of mean reversion to our long-run relationships. In particular, we benefit from the multivariate Johansen (1988) methodology which has become a popular workhorse in applied research when examining long and short run issues. Additionally we utilise a LQAC model and the estimation approach developed by Dolado *et al.* (1990) and Gregory *et al.* (1993). This indicates whether our results are robust when a less *ad hoc* framework, than that used by others, is adopted.

Studying the UK, we found evidence that income and the real effective exchange rate were important determinants of reserves using the single equation LQAC model and the multivariate approach. Both income and the exchange rate were significant in a long-run relationship, which was also cointegrating. The estimated coefficient on the exchange rate was significantly positive, consistent with "leaning-against-the-wind". For example, when the exchange rate is appreciating and this is believed to be unmerited, the monetary authorities will accumulate reserves. This is to prevent the deleterious impact of such an appreciation on exporting sectors. We found little evidence of the statistical significance of prices in a long-run specification using methods robust to non-stationarity. The inclusion of prices was not necessary to produce a reasonably signed cointegrating relation within the Johansen approach. Mean reversion to our long-run relationship was found in a number of specifications, with a half-life of around four quarters. This suggests UK policy-makers will correct any disequilibrium in the long-run relationship for reserves with a lag, highlighting the importance of a disequilibrium model in any descriptive study of reserves. We believe the lag in adjustment represents policy makers desire to avoid any undesirable side effect that operations to restore a long-run reserve equilibrium have on the domestic monetary situation or on the exchange rate. Additionally, we found evidence of a small discount factor within the LQAC model, which was quite distinct from the size estimated by studies applying these methods in other areas. This questions how forward-looking the monetary authority actually is in setting reserves.

In Chapter Three we considered the interaction of reserves and domestic monetary variables in more detail. We had four main motivations in this Chapter. We extended our short-run model for reserves to take account of the monetary disequilibria of private agents, as suggested by the Monetary Approach to the Balance of Payments. This view proposes that any short-run model of reserves based only on disequilibrium from long-run reserves will be misspecified, resulting in biased estimated coefficients. This brings into question the degree of mean reversion that we found in Chapter Two. To produce a disequilibruim term for private agents' real money balances it was necessary to construct a long-run relationship for money demand. We believe that the latter is also of independent interest.

There has been considerable research effort since the mid-1970s examining whether a stable money demand relationship existed. This was initiated by Goldfeld's (1973 and 1976) examinations of US data. Using cointegration methods as the benchmark evidence of a money demand relation, researchers have often found it necessary to include a cumulative measure of interest rates. Recent research suggests that including a cumulative measure may not be needed with a sample that includes data from the 1980s and early 1990s, when nominal interest rates were less high and variable than in the 1970s. Additionally, we explicitly test for the inclusion of deterministic components in our analysis which researchers, to the best of our knowledge, have not considered before within the context of money demand. Combining a specification of the domestic monetary situation with our analysis of open economy issues allows us to consider Juselius' (1996) proposal concerning a monetary model estimated within a non-stationary VAR. Juselius studied domestic monetary interaction and suggested this would benefit from making reference to the open economy. We do so in our study. Finally, we considered the degree of monetary sterilisation and whether changes in reserves have implications for the money supply.

We obtained evidence of a cointegrating money demand relation, which does not use a cumulative *ad hoc* measure of interest rates. We then use this long-run relationship and a reserve equation to construct a short-run model for reserve adjustment. Disequilibrium in private agents' real money balances was not found to have a significant impact on changes in reserves. We consequently discount the idea that balance of payments disequilibrium is self-correcting by inducing changes in reserves. Our short-run equation for reserves from Chapter Two is, on this basis, not mis-specified. Disequilibrium in reserves is again significant and mean reverting in a short-run reserves equation, but with a slightly slower rate of adjustment than in Chapter Two. Testing the short run model for the degree of reserve sterilisation did suggest, however, that M0 is dependent upon reserves. We verified this evidence using impulse response functions. Although there is no immediate impact of reserves on money, this relationship unwinds over time. However in the long-run we do find a re-adjustment to remove this monetary effect. This is slightly more consistent with long run evidence that reserves do not have an impact on domestic monetary aggregates as found by other authors. But we find clear evidence of less than complete sterilisation through time, consistent with Neumann and von Hagen's (1992) study of German data.

Interestingly, when we consider the monetary sector our impulse response functions indicated that M0 has a clear impact on inflation. This implies that M0 would be of some use as a leading indicator of inflation. Although there has been a move away from monitoring ranges for broad and narrow money and a move towards explicitly targeting inflation, given the lag in monetary policy in the UK, this effectively means that the Bank of England targets an inflation forecast. Since M0 has useful properties as a leading indicator in this setting, it may remain of some operational significance in the conduct of UK monetary policy.

A number of interesting possibilities are opened up for further research in the area of joint modelling of money demand and reserves. It would be of a great deal of interest given the significance of M4 in central bank policy, to use our approach with a broader monetary aggregate. This would allow us to consider whether M4 has any use as a leading indicator of inflation in comparison to M0. We could also consider whether the degree of sterilisation of reserve changes was of the same order and with the same time pattern for the broader monetary aggregate. Within the context of narrow or broad money we could extend our analysis to other European countries. This would be useful for considering the transmission mechanism of monetary policy and whether the UK and other Euroland countries are similar in their responses to monetary impulses.

The Fourth Chapter returns to specification issues in the determination of reserves. We do so by widening our sample and re-considering our results from before, which suggest an important role for the real effective exchange rate in reserve determination. This approach has the benefit of a broader sample, which considers how relevant these results are for other countries. Having broadened our study we can also use developments in non-stationary panel estimation, which has become an informative and popular method of analysis for applied researchers. To the best of our knowledge, recent developments in panel data have not been used in empirical studies of reserve demand. We consider this to be a significant contribution to the literature. It is a view consistent with Gonzalo (1994) and Gregory (1994), that given the susceptibility of even the latest developments in time series estimation to small sample bias we can corroborate our time series results by using non-stationary panel methods.

Our time series results are not definitive but they highlight the relevance of a newer specification for reserve holdings for European countries. They replicate results for the UK from previous Chapters, emphasising the importance of income and the real effective exchange rate. From the panel studies we find that once we ensure cross sectional independence (i.e. remove a potential common factor problem) we find that evidence in favour of a newer specification dominates evidence for a reserves specification based on the traditional portfolio-balance approach. For example, we find evidence of a positive and significant estimated coefficient for income and the real exchange rate, and greater evidence of cointegration using a number of panel tests.

For a combined sample of nine European countries, we also find that average mean reversion is much quicker with a newer specification in comparison to results only for the UK. This emphasises the importance of reserve holdings for countries participating in the ERM and is consistent with more formal financing arrangements post-1987.

A suggestion for further research in this area would be the incorporation of multiple cointegrating vectors within a non-stationary panel environment. This would allow us to consider whether monetary disequilibrium has an important short-run effect on reserves as suggested by the monetary approach to the balance of payments. The non-stationary panel methods we use here are single equation approaches. Although Larsson *et al.* (1998) introduce a panel Johansen Trace test and the ability to test long-run hypotheses, no short-run developments dealing with multiple cointegrating vectors have been made. Such a development would allow us to consider whether both reserve disequilibria and private agents' monetary disequilibria were important determinants of short run changes in reserves. As the empirical literature stands at the moment, it would presently be only possible to extend a time-series approach to consider these questions for a greater sample of countries.

Remaining within the broad area of international reserves in Chapter Five, we switch our analysis to the impact of foreign exchange operations on the exchange rate. We considered the relationship between high frequency foreign exchange intervention and the second moment of the exchange rate. We illustrated that reserves continue to be held by central banks to facilitate foreign exchange operations, with a view to influencing the external value of the domestic currency. Recent studies have extended research on the level of the exchange rate to consider the impact of foreign exchange operations on the volatility of returns data. We extent the volatility literature in a number of ways.

We explicitly introduce coordinated intervention, which has not, to the best of our knowledge, been considered before in the literature on exchange rate volatility. We replicate studies examining the level of the exchange rate by differentiating the impact of unilateral and coordinated intervention. This approach additionally allows us to model spillovers in more detail. The implicit suggestion is that spillovers were masked in previous studies by the possibility of coordinated intervention. Using actual intervention data rather than reported, we also revisit Dominguez's (1998) study; consistent with the general criticism by Edison (1993) that the literature in this area is particularly fractured. This should provide distinct and possibly additional information to Dominguez's (1998) study, regarding the impact of foreign exchange intervention. Finally, we consider UK Pound bilateral rates - not previously examined in the literature - and assess whether there are spillovers from Fed and Bundesbank policy to the Pound.

In our results in Chapter Five, unilateral US Federal Reserve intervention is associated with a fall in the volatility of the Dollar bilateral rates against the Pound and DM. There are also 'benign' spillovers (i.e. intervention associated with a fall in volatility) to the Pound-DM rate. In contrast, Bundesbank intervention leads to an increase in volatility for the Dollar and Pound against the Mark, and is associated with increased volatility of the Pound-Dollar rate. Our specification that includes both intervention variables may have a problem with collinearity, since the US and German central banks often use concerted intervention. This will be removed by introducing our distinction between unilateral and coordinated intervention.

Fed and Buba coordinated intervention is significantly and positively related to Pound-Dollar volatility. On the other hand, unilateral US intervention is statistically associated with a fall in volatility. Coordinated intervention is also distinct from unilateral German intervention, which is negatively associated with volatility. This highlights the importance of our distinction, which was also found for the DM-Dollar rate and the DM-Pound.

In general terms it should not necessarily be considered as a malign feature of intervention, that it is associated with an increase in volatility. This may be due to our chosen sample period coinciding with an explicit policy that attempted to induce changes in the level of the exchange rate. The success of this policy may manifest itself in increased volatility. What is especially interesting is that unilateral US policy has lead to a fall in volatility but when US policy is coordinated volatility increases. This suggests that policy objectives or policy effectiveness changes. The latter could well explain the situational nature of results in this field - different sample periods produce different results.

As a proposal for further research in this area, using a Switching ARCH (SWARCH) method within the context of the changing influence of foreign exchange intervention may be useful. We could implement this approach by including a Markov switching parameter into the ARCH model (as suggested by Hamilton and Susmel, 1994). Alternatively we use some form of Kalman filter estimation. This would also be a direct extension of the distinction we make between unilateral and coordinated intervention, in the sense that it may pick up the changing objectives of policy.

It may also be an idea for further research to check that our central hypotheses, for instance, whether the distinction between coordinated and unilateral intervention is important, are robust to different measures of volatility. For example, we could adopt Anderson and Bollerslev's (1998) approach of constructing a measure of daily volatility from ultra-high frequency exchange rate data. Dominguez (1999) has utilised this method in a recent paper on intervention. It would also be interesting to consider if we can obtain a measure of volatility that is related to a model of exchange rate behaviour (e.g. the monetary model of exchange rates). This would indicate whether volatility has been excessive on the basis of economic fundamentals. This could, in addition to assessing the impact of intervention, have the benefit of rationalising some of the failure of traditional fundamentalist models to fit exchange rate behaviour. This excessive measure of volatility has been advocated as an argument for throwing sand into the wheels of currency markets. This would give us an indication of whether volatility is due to fundamentals and whether this is statistically reduced by intervention.

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