ESSAYS IN MACROECONOMICS AND INTERNATIONAL FINANCE

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BY

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DECLARATION

I declare that four papers, titled "Real Interest Rates, Liquidity Constraints and Financial Deregulation: Private Consumption Behaviour in the UK", "Capital Flows to Developing Countries: Short-Term and Long-Term Determinants", "European Capital Flows and Regional Risk" and "The Behaviour of Real Exchange Rates During the Post-Bretton Woods Period", drawn on Chapters 1, 2, 3 and 4 of this thesis, are forthcoming in the *Journal of Macroeconomics*, *The World Bank Economic Review*, *The Manchester School* and the *Journal of International Economics* respectively.

I declare that the University Librarian is granted powers of discretion to allow this thesis to be copied in part or in whole without further reference to the author.

Signed

Lucio Sarno

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ABSTRACT:'Essays in Macroeconomics and International Finance' by Lucio Sarno

In this thesis we examine four different issues in macroeconomics and international finance which have received widespread attention from researchers and which still represent very debated issues.

In Chapter 1, we investigate the effect of financial deregulation on consumption expenditure in the UK and in France. A non-linear model for consumption which allows for liquidity constraints through a time-varying parameter dependent on a proxy for financial deregulation is estimated using non-linear instrumental variables. It is concluded that financial deregulation has significantly reduced liquidity constraints faced by consumers, allowing a higher percentage of the population to smooth consumption over time. The estimated path of the time-varying parameter, interpreted as the proportion of income going to liquidity-constrained consumers, is consistent with the process of financial deregulation in the UK and in France respectively.

In Chapter 2, we look at a range of issues concerning capital flows in Europe. We gauge the extent of international capital mobility through an examination of savings-investment correlations and investigate the difference between the short-run and the long-run savings-investment correlation coefficient, in order to shed light both on the validity of the Feldstein-Horioka regression as a means of measuring the degree of capital mobility and on its implications. Using quarterly UK data, we also examine the effectiveness of the abolition of exchange control which, in October 1979, ended a long period of restrictions on capital flows between the UK and the international economy. We find that, consistent with the logical implication of the Feldstein-Horioka regression, the short-run correlation is significantly higher than the long-run correlation. In contrast with much of the literature employing the Feldstein-Horioka interpretation, however, the results suggest that the UK is highly financially integrated with the global economy post 1979. In addition, we provide further evidence on the extent and effect of European capital mobility through an examination of GNP/GDP ratios across the largest European countries and a comparison of these with similar calculations executed for regions within the UK. We also examine the role of private financial markets in reducing regional risk by developing an empirical model which is then estimated on data for the European Union.

In Chapter 3, we focus on the determinants of the large US portfolio flows received by Latin American and Asian countries over the sample period 1988-92. Using cointegration techniques, we find that both domestic and global factors contribute to explaining capital flows to developing countries in the form of bonds and equities and represent significant long-run determinants of portfolio flows. We also investigate the dynamics of portfolio flows by estimating seemingly unrelated error-correction models. Global and country-specific factors seem to be equally important in determining the long-run movements in equity flows for both Asian and Latin American countries, while global factors appear to be much more important than domestic factors in explaining the dynamics of bond flows. In particular, US interest rates are found to be a particularly important determinant of the short-run dynamics of portfolio, especially bond, flows to developing countries. We also measure the relative size and statistical significance of permanent and temporary components of portfolio flows to developing countries using Kalman filtering maximum likelihood techniques and variance ratio tests corrected for small sample upward-bias. Overall, the empirical results strongly indicate that if there is a permanent component in equity and bond flows to developing countries, it is very small in relative size, suggesting that capital flows to developing countries over the sample period were highly temporary. This, in turn, has important policy implications.

In Chapter 4, we examine the behaviour of real exchange rates during the post-Bretton Woods period. Since standard tests for mean reversion in real exchange rates may lack power with sample periods corresponding to the span of the recent float, researchers have employed more powerful multivariate unit root tests. We point out a potential problem with such tests, namely that the null hypothesis of joint non-stationarity may be rejected even when only one of the series is a realization of a mean-reverting process. We suggest another multivariate test in which the null hypothesis is violated only when *all* of the processes in question are stationary. This test is easily constructed and has a known limiting distribution. We investigate the finite-sample empirical performance of both tests using Monte Carlo techniques. Finally, we apply the test procedures to quarterly data on real exchange rates among the G5 countries over the recent floating rate period.

In sum, this thesis adds to the relevant literature on the four different issues examined by providing insights and evidence to researchers and indicating potential avenues for future research.

To my parents, Raffaele and Teresa

'In primisque hominis est propria veri inquisitio atque investigatio.'

(Cicero, *De Officiis*, Book I.iv)

INTRODUCTION AND OVERVIEW

In this thesis we examine four different topics in macroeconomics and international finance which have received widespread attention from researchers and which still represent very debated issues, in an attempt to provide interesting contributions to the literature as well as results of interest to practioners and policy-makers.

Macroeconomics and international finance are alive with highly controversial theoretical issues, unexplained empirical evidence and great practical questions, whereas consensus and theory-consistent stylised facts are still relatively limited in those fields. We especially focus on four puzzling and unresolved issues which particularly attract our personal interest. These are concerned with the excess sensitivity of consumption to income and the effect of financial policies on consumption expenditure, the finding of relatively low international capital mobility implied by the high correlation between savings and investment shares of gross domestic product across countries, the identification of the main factors generating the recent large surge of capital flows towards emerging markets and the validity of the purchasing power parity condition during the post-Bretton Woods period respectively.

The consumption function literature provides some of the most fascinating reading in macroeconomics and we certainly agree that 'attempts by economists to understand the saving and consumption patterns of households have generated some of the best science of economics' (Deaton, 1992). After the earlier literature on modelling consumption expenditure during the 1950s, when the basis of the modern theory of consumption was provided, the last twenty years - precisely since the publication of Hall's (1978) paper on the forward-looking consumption function -

have been characterised by 'a great flurry of activity' (Deaton, 1992). This activity has been concerned mainly with testing the Hall model of consumption and the empirical failure of the theory has often been attributed to the presence of liquidity constraints. The empirical evidence from the estimation of forward-looking consumption function models which allow for liquidity-constraints has, however, implications which are quite counterintuitive: in particular, extant empirical studies appear to show that the process of financial deregulation in the UK and in France has generated a *perverse* effect - increasing the proportion of income going to liquidity-constrained consumers - in the UK and *no significant* effect in France (see Muellbauer, 1994 for a survey). It is in this sense that we regard the empirical evidence on the excess sensitivity of consumption a puzzle, an awkward empirical fact that refuses to fit onto the relevant theoretical framework.

The second topic analysed in this thesis is also a puzzle. Since the late 1970s, it has become common among economists to describe the world financial system as characterised by perfect international capital mobility, following the drastic reduction or complete removal of capital controls in the mid 1970s by many of the major industrialised countries, in addition to the ongoing process of widespread financial deregulation, the improvement of information and communication technology in financial markets and the recycling of OPEC surpluses to developing countries (Frankel, 1993; Coakley, Kulasi and Smith, 1995a). One of the most widely used measures of the degree of international capital mobility, suggested by Feldstein and Horioka (1980), is based on the correlation between domestic saving and investment. The rationale is that if capital is mobile internationally, there need not be any correlation between saving and investment. In a world of perfect capital mobility, savings may be expected to flow from countries with a relatively high propensity to save to countries with relatively more favourable investment opportunities. Nevertheless, the empirical evidence based on saving-investment correlations strongly indicates, apart from very few exceptions, that the degree of international capital mobility is very low in the world economy. This finding is in sharp contrast with the assumption of perfect capital mobility, routinely used in the literature on exchange rate determination, with the exception of variants of the Mundell-Fleming model and the portfolio balance model (Taylor, 1995; Taylor and Sarno, 1997, Chapter 4). Traditionally, attempts to explain the Feldstein-Horioka puzzle have been based on attacks against the meaning of the association between savings and investment, rather than against the the adequateness of the econometric techniques employed by researchers. In this thesis, we provide an alternative explanation of the Feldstein-Horioka puzzle in that we show how the Feldstein-Horioka definition of international capital mobility may be informative if an adequate econometric strategy is used, while the estimation procedures employed by the previous literature may be misleading.

The third topic examined is concerned with the identification of the main determinants of the recent, enormous surge of US capital flows to developing countries. The sharp expansion of net and gross capital flows together with a marked increase in the participation of foreign investors and foreign financial institutions in the financial markets of developing countries is one of the main features of the recent development of international capital flows, also enhanced by the ongoing process of abolition of various sorts of impediments and capital controls and the broader liberalization of financial markets in developing countries from the late 1980s. Another main feature of the recent trend of capital flows to developing countries is that US private (bond and equity) as opposed to US official capital flows are increasingly a crucial source of financing of large current account imbalances (World Bank, 1997). These trends in the pattern of net and gross capital flows raise important issues concerning which factors motivate the flows and how these flows affect the performance of developing countries. In particular, it is of interest to assess the relative importance - on the one hand - of the improved economic performances of developing countries (country-specific or "pull" factors) and of the stimulus - on the other hand - provided by the decline of US interest rates and the slowdown of the US economy in the early 1990s (global or "push" factors) (Chuhan, Claessens and Mamingi, 1993; Fernandez-Arias, 1996; Agenor, 1997). The reason why the identification of the relative importance of push and pull factors is important for the effective design of policy - and therefore worthy of investigation - has recently been made forcefully by Fernandez-Arias and Montiel (1996), who list a number of arguments describing why large capital flows may, under various circumstances, have adverse effects on developing countries unless proper policies designed to neutralise such effects are adopted. Unlike the other themes treated in this thesis, the issue of the investigation of the forcing factors inducing capital flows to developing countries in the 1990s could neither generate a puzzle nor support a stylised fact as it is a quite new and hence only recently explored issue on which there is not sufficient evidence to generate a puzzle or a stylised fact.

The fourth and last topic investigated in this thesis, the behaviour of real exchange rates during the post-Bretton Woods period, is certainly a puzzle, as noted recently in an excellent survey of the relevant literature by Rogoff (1996). Purchasing power parity (PPP), viewed as a theory of exchange rate determination or as an equilibrium condition or also as an efficient arbitrage condition in either goods or asset markets, has the strong implication that movements in the nominal exchange rate are proportional to the ratio of national price levels or that the real exchange rate is constant over time.

The modern articulation of PPP starts in the immediate post World War I, after the collapse of the world financial system and as an attempt to restore a stable international financial environment. Since the original formulation due to Cassell (1921, 1922), however, the profession has radically switched position several times with regard to the validity of PPP. During the pre-Bretton Woods period, the professional consensus was quite supportive of PPP as a condition holding over long periods of time (eg. Friedman and Schwartz, 1963; Gaillot, 1970). A much stronger position was, however, taken in the early 1970s by some proponents of the monetary approach to exchange rate determination, who posited continuous, rather than long-run, purchasing power parity (eg. McCloskey and Zecher, 1976; Frenkel, 1976; Frenkel and Johnson, 1978). A radical change in the position of the profession on the behaviour of real exchange rates occurred in the mid to late 1970s. In fact, following the very high variability of real exchange rates after the collapse of the Bretton Woods system, the extreme proposition of continuous PPP was largely abandoned (eg. Krugman, 1978; Frenkel, 1981). In addition, the relevant literature of the 1980s could not reject - by large - the hypothesis of random walk behaviour in real exchange rates (eg. Roll, 1979; Adler and Lehmann, 1983; Piggott and Sweeney, 1985), consistently with the failure to find cointegration between nominal exchange rates and relative prices (eg. Taylor, 1988; Corbae and Ouliaris, 1988; Enders, 1988; Mark, 1990), further reducing the confidence of the profession in purchasing power parity and leading to the widespread belief that PPP was of little or no use empirically (eg. Dornbusch, 1988). Overall, the validity of PPP is still one of the most debated issues in international finance.

The remainder of the thesis is set out as follows. In Chapter 1, we investigate the effect of financial deregulation on consumption expenditure in the UK and in France. A non-linear model for consumption which allows for liquidity constraints through a time-varying parameter dependent on a proxy for financial deregulation is estimated using non-linear instrumental variables. It is concluded that financial deregulation has significantly reduced liquidity constraints faced by consumers, allowing a higher percentage of the population to smooth consumption over time. The estimated path of the time-varying parameter, interpreted as the proportion of income going to liquidity-constrained consumers, is consistent with the process of financial deregulation in the UK and in France respectively.

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A final chapter briefly summarises, highlights the key findings of the thesis and concludes.

CHAPTER 1

Real Interest Rates, Liquidity Constraints and Financial Deregulation: Private Consumption Behaviour in the UK and in France

'Large errors in forecasting consumption in the late 1980s and early 1990s were a major contribution to the poor performance in forecasting overall economic developments. These errors have been attributed to the failure of existing consumption functions adequately to capture important influences on consumer behaviour, such as the effects of financial deregulation ...'

Church, Smith and Wallis (1994, p. 71)

1.1 Introduction

According to the permanent income hypothesis, changes in consumption should be unpredictable, since revisions to the planned consumption path can only arise in this model from news about future income which, by definition, is unpredictable (see eg. Deaton, 1992, p. 83). A similar result is obtained in the representative agent model of Hall (1978, 1988), with the assumption of either a constant real interest rate or a zero intertemporal elasticity of substitution of consumption. Following Campbell and Mankiw (1989, 1990, 1991), this study focuses on the Hall consumption function allowing for the presence of liquidity constraints and, in addition, for financial deregulation and a time-varying real rate of interest. The relevant literature argues that if a significant proportion of consumers are unable to smooth their consumption effectively because of liquidity constraints, then movements in current income may be an additional explanatory variable of consumption. The main objective of this study is to investigate the behaviour of aggregate consumption when the proportion of liquidity-constrained consumers shifts over time due to the effects of financial deregulation - as one might expect to have been the case in the UK and in France.

Using quarterly data for the UK during the period 1963-1994 and for France during the period 1970-1993, the analysis results in the application of a time-varying parameter empirical model of consumption behaviour which is tested against an alternative error correction model which has generally been found to perform well, especially on UK consumption data.

The remainder of the chapter is set out as follows. In Section 1.2, we provide a review of the relevant literature and develop the estimating equations. In Section 1.3, we provide a brief historical excursus concerning the process of financial deregulation in the UK and in France as a motivation for the analysis. In Section 1.4, we discuss the data set employed. In Section 1.5, we provide a discussion of some of the estimation issues and discuss the estimation results. A final section concludes.

1.2 Modelling consumption: theory and evidence

1.2.1 Intertemporal choice, excess sensitivity and random walks

Consider a representative consumer facing a time horizon of T future periods. Assuming additivity of the utility function, the consumer's optimization problem in period 0 involves maximising the discounted present value of expected future consumption - ie. maximising the objective function:

$$V_0 = E_0 \sum_{t=0}^{T} (1 + \rho)^{-t} U(C_t) \tag{1.1}$$

where C_t is real consumption at time t, ρ is the individual rate of time preference, U(.) is the single-period utility function; E_0 is the mathematical expectation operator conditional on information at time 0. We assume $U'>0$, $U''<0$, implying strict concavity and twice differentiability of the utility function. In each period the utility maximisation problem of the consumer is governed by the asset evolution equation or budget constraint:

$$A_{t+1} = (1 + r_t)(A_t + Y_t^{\alpha} - C_t), \qquad t=0,\ldots,T \tag{1.2}$$

where A_t represents net real assets at the beginning of the period t, Y^{α}_t is real labour income during time t and r_t is the real rate of interest at which the representative consumer can borrow or lend from the beginning of period t to the beginning of period t+1. The first-order conditions for the maximization of (1.1) subject to (1.2) are a set of stochastic Euler equations of the form[1]:

$$U'(C_t) = E_t[(1 + \rho)^{-1}(1 + r_t)U'(C_{t+1})] \tag{1.3}$$

Assuming an isoelastic utility function of the form $U(C_t) = \sigma(\sigma-1)^{-1}C_t^{(\sigma-1)/\sigma}$, where σ is the intertemporal elasticity of substitution, equation (1.3) may be closely

[1] In order to derive the forward-looking or permanent income consumption from (1.1) and (1.2) it is necessary to specify in addition a transversality condition on the value of assets. This is needed in order to rule out Ponzi schemes - ie. borrowing to finance consumption and then borrowing forever in order to service the increased debt. Although a transversality condition is implicitly understood to apply, it need not be stated explicitly in the present analysis because of the focus of this study on the Euler equations.

approximated by[2]:

$$E_t \Delta c_{t+1} = \sigma (E_t r_t - \rho) + 0.5\sigma^{-1}\omega^2_{t+1} \tag{1.4}$$

where $\Delta c_{t+1} \equiv \log(C_{t+1}) - \log(C_t)$ and where

$$\omega^2_{t+1} = Var_t(\Delta c_{t+1} - \sigma r_t) \tag{1.5}$$

where $Var_t(.)$ denotes conditional variance.

Typically, ω^2_{t+1} is implicitly assumed to be constant and absorbed into the intercept term (Hall, 1978, 1988), although, as we discuss in the next sub-section, it may be interpreted as the contribution of the precautionary motive for saving. For the moment, we hold this term constant so that (1.4) collapses to[3]:

$$\Delta c_{t+1} = \mu + \sigma r_t + \epsilon_{1\,t+1} \tag{1.6}$$

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[2] The approximation arises only because of the normal approximations made with logarithmic transforms (Deaton 1992, p. 64); note that if we assume that $\log(C_{t+1})$ and r_t are joint normally distributed, then (1.4) holds exactly, given (1.3) (see Appendix 1.1).

[3] In the empirical work reported below, we tested for autoregressive conditional heteroscedasticity (ARCH) effects in the residual series for the estimated models as well as in the growth of consumption directly, but were in no case able to discover significant ARCH effects. This therefore implies that we may be justified in treating this conditional variance term as constant.

where $\epsilon_{1\,t+1}$ may be correlated with r_t, but is uncorrelated with lagged variables[4].

If agents are able to smooth their consumption patterns through the capital market, then equation (1.6) implies that an increase in ex ante real interest rates will generate an increase in current consumption by increasing the amount of future consumption which can be undertaken with the same total resources, some of which is transferred to the present. How much is transferred to the present depends directly on the magnitude of the elasticity of intertemporal substitution.

Hall (1978) notes that if the real rate of interest is constant, then (1.6) implies that consumption follows a random walk with drift. Hall (1978) provides empirical evidence on US data which shows that the change in consumption is orthogonal to lagged income, but not to lagged stock prices. Despite the formal rejection of the random-walk model, however, he argues that past stock prices provide such a small contribution to the variation in consumption in his empirical work that the random walk model constitutes a satisfactory approximation. Hall (1988) interprets the empirical evidence supporting the random walk behaviour of consumption as evidence that the intertemporal elasticity of substitution σ is very close to zero[5].

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[4] There are a large number of implicit and explicit assumptions underlying the Hall random walk consumption function, including no credit restrictions or other non-linearities in the budget constraint; no habits or adjustment costs; non-durable goods; the subjective discount rate, ρ, is equal for all consumers and also equal to the market real rate of interest, r; no measurement errors or transitory shocks to consumption; the coincidence of the frequency of consumers' decisions with the frequency of the data used; a constant real rate of interest; rational expectations (Muellbauer, 1994). Most of the literature on consumption following Hall (1978) is concerned with the examination of the implications of relaxing one or more of the underlying assumptions.

[5] Note that, under the assumption of an isoelastic utility function, the intertemporal elasticity of substitution is, strictly speaking, the reciprocal of the

A large number of empirical studies for different sample periods and countries have rejected, however, the random walk model even as an approximation. In particular, for the UK, Davidson and Hendry (1981) and Daly and Hadjimatheou (1981) find that the orthogonality condition on the disturbance term is violated when tested against lagged income, consumption and liquidity measures. One of the most interesting findings of the work of Davidson and Hendry (1981) is that they show, using Monte Carlo simulations, that a random walk in consumption may well be accepted when, in fact, it does not hold, simply because it would be hard to distinguish from an error-correction model for consumption.

Another well-known important strand of the literature has been stimulated by Flavin (1981), who, using US data, finds "excess sensitivity" of the consumption pattern to current and lagged income, ie. the contribution of lagged changes in income to forecasting changes in consumption is significantly greater than implied by the life cycle-permanent income hypothesis (LCPIH)[6]. A number of empirical

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coefficient of relative risk aversion, so that as the former approaches zero, the latter approaches infinity. Hall (1988), in order to avoid this implication of the random walk consumption function, provides a number of arguments which attempt 'to eliminate the automatic connection' between these two coefficients (see Hall, 1988, pp. 343-345).

[6] Mankiw and Shapiro (1985, 1986), however, argue that if the time series for real income possesses a unit root - as data for most major industrialised countries seems to demonstrate - then the Flavin procedure may be heavily biased towards rejection of the permanent income hypothesis. There are, however, other test procedures which are immune to the unit root problem and which also generate evidence of significant excess sensitivity (see eg. Stock and West, 1988; Deaton, 1992, pp. 92-96). An example is the regression of consumption growth on lags in income growth and on lagged saving, which is a stationary variable, even if income is not (Campbell and Mankiw, 1989).

studies investigating the excess sensitivity of consumption are able to reject the Hall model (Deaton, 1992).

1.2.2 Precautionary saving and liquidity constraints

Two interesting issues have been at the centre of an important strand of recent research on consumption: precautionary saving and liquidity constraints, both of which may be adduced as an explanation of excess sensitivity.

If the utility function has a positive third derivative - ie. the marginal utility function is convex - then saving is required to self-insure against the uncertainty of future labor income (Kimball, 1990). In equation (1.4), the term $0.5\sigma^{-1}\omega_{t+1}^2$ represents precautionary saving, since increases in the uncertainty of future consumption entail a reduction in current consumption and hence an increase in the growth of consumption between current and future periods. In the present case (isoelastic utility functions), the degree of prudence governing this term - measured by σ^{-1} - is identical to the coefficient of relative risk aversion; in general, however, statistics governing prudence and risk aversion will be different (Kimball, 1990). As equation (1.5) demonstrates and Carroll (1992) notes, any variable which helps predict future variability in consumption will help predict the rate of growth of consumption. Thus, precautionary saving may be adduced as an explanation of excess sensitivity of consumption to income movements.

Investigation of the effect of precautionary saving on consumption patterns has depended on numerical solutions (Skinner, 1988; Zeldes, 1989a; Carroll, 1992; Deaton, 1991; see Deaton, 1992 and Browning and Lusardi, 1996 for surveys), which tend to indicate - *inter alia* - that households will save more earlier in the life cycle than indicated by the permanent income model. The empirical importance of precautionary saving is a topic of current debate in the literature (Deaton, 1992; Browning and Lusardi, 1996). In the present study, it is implicitly assumed that precautionary savings have a small or constant effect on consumption behaviour.

Another way in which excess sensitivity of consumption to income may arise is where some agents encounter binding liquidity constraints and therefore are forced to consume entirely out of current income (Hall and Mishkin, 1982). An individual or an household is facing liquidity constraints if he or she is unable - for whatever reason - to borrow against future earnings beyond a certain non-negative limit[7]. It is apparent that there are close similarities as well as differences between the effects on consumption behaviour generated by precautionary saving and by liquidity constraints. Both cases result in a relationship between consumption and income which has a 'Keynesian flavor' (Deaton, 1992), but for different reasons: liquidity-constrained models because consumption is assumed to respond to income directly, and precautionary saving models because uncertainty concerning the future is discounted into the present (Deaton, 1992)[8].

Apart from affecting changes in consumption expenditure and inducing a

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[7] An alternative, weaker definition of liquidity constraints is also given by Attanasio (1995). According to the second definition, a liquidity-constrained consumer faces a difference between borrowing and lending interest rates and, more generally, interest rates offered to him or her are not independent of his or her net asset position.

[8] Attention to liquidity constraints was given in the 1970s by Flemming (1973), who argues that consumption expenditure may depend on current disposable income in the short run as an effect of the inability to borrow against future income. Another early study is due to Tobin and Dolde (1971), who argue that, in presence of binding liquidity constraints, the marginal propensity to consume out of available resources is much higher than it would be in a simple life-cycle model without liquidity constraints.

positive relationship between consumption and income growth, binding liquidity constraints also have other implications. In particular, the presence of liquidity constraints may lead to an increase of the aggregate saving rate in steady state. This is shown by Jappelli and Pagano (1994) using an overlapping generation model. An important piece of evidence is due to Guiso, Jappelli and Terlizzese (1994), who find that the relatively high Italian saving rate may be largely generated by the slow development of the Italian financial markets where borrowing is generally difficult and costly.

With regard to the welfare consequences, clearly the presence of liquidity constraints generates a welfare loss, which depends on - *inter alia* - the tightness of the constraints, the technology available for self-insurance, the variability of idiosyncratic shocks, the specific nature of the constraints. Imrohoroglou (1989), for example, assesses the welfare loss caused by liquidity constraints in an equilibrium model with infinitely lived agents. Imrohoroglou's study suggests that welfare decreases considerably as an effect of the inability to insure idiosyncratic risk and the size of the cost depends upon the variability of the idiosyncratic shocks and on the curvature of the utility function of the representative agent.

Another important implication of the presence of liquidity constraints is concerned with policy issues and, in particular, with the Ricardian equivalence proposition. The Ricardian equivalence proposition states that government debt should not be considered net wealth as rational agents foresee that future taxes will be needed to repay it; this implies neutrality of fiscal policy (Barro, 1974). If liquidity constraints prevent the intertemporal allocation of consumption, however, government debt becomes a means for transferring resources from the future to the present, implying that the Ricardian equivalence proposition does not hold and, therefore, fiscal policy may be effective and beneficial (Tobin, 1980).

Zeldes (1989b) uses time-series and cross-section data on families from the Panel Study of Income Dynamics to show that an inability to borrow against future labour income affects the consumption of a significant fraction of the US population.

Jappelli (1990) relies on survey questions in order to assess the importance of liquidity constraints in the US, using the 1983 Survey of Consumer Finances (SCF), which contains questions not only on the availability of credit, but also on whether the lack of consumer credit is due to unssuccessful loan applications or also to the pessimism of individuals who did not apply for a loan because they felt their applications would have been unsuccessful. Using the SCF, Jappelli (1990) can classify liquidity constrained consumers in various categories and provide evidence that lower-educated, younger and black people are more likely to encounter liquidity constraints. Unfortunately, however, no information is provided on consumption expenditure in the SCF and, therefore, differences in consumption behaviour of the liquidity constrained consumers cannot be characterised.

Campbell and Mankiw (1989) argue that the empirical link between consumption and income may be due to the fact that while the permanent income model of consumption may be applicable to some consumers, other consumers may be guided by "rules of thumb" according to which they consume their current rather than their permanent income because they encounter binding liquidity constraints in some time periods. There is, moreover, no inconsistency in some groups of consumers encountering liquidity constraints in some time periods while other groups do not. Stiglitz and Weiss (1981), for example, apply an argument based on the principle of adverse selection to demonstrate why there may be credit rationing even when interest rates are fully flexible[9].

Specifically, for the liquidity-constrained group of consumers, consumption growth may be assumed to be directly proportional to the growth in current disposable income, while the remainder of consumers may be assumed to satisfy the general Euler equation (1.6). If the proportion of income which goes to liquidity-constrained consumers at time t is λ_t, then the following equation may be derived:

$$\Delta c_t = \lambda_t \Delta y_t + \theta_t r_{t-1} + \epsilon_t \tag{1.7}$$

where:

$$\epsilon_t = (1 - \lambda_t)\epsilon_{1t} \tag{1.8}$$

and

$$\theta_t = (1 - \lambda_t)\sigma \tag{1.9}$$

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[9] Work on liquidity constraints has been carried out by, *inter alios*, Hall and Mishkin (1982), Clarida (1987), Zeldes (1989b), Jappelli and Pagano (1989), Deaton (1991), Nicoletti-Altimari and Thomson (1995). For a recent, comprehensive review of the implications of the presence of liquidity constraints on consumption behaviour and a survey of the theoretical and empirical relevant literature, see Attanasio (1995).

In the framework of Campbell and Mankiw (1989) the proportion of income going to liquidity-constrained consumers is treated as constant. Using US data, their empirical point estimates of λ imply that the fraction of income received by liquidity-constrained consumers is about fifty percent, indicating a significant departure from the life cycle-permanent income hypothesis.

In a further paper, Campbell and Mankiw (1991) investigate whether the same model fits quarterly data for the UK over the period 1957-1988 and for Canada, France, Japan and Sweden over the period 1972-1988, exploring various generalizations of the basic model. They find - in common with Jappelli and Pagano (1989) - that the international pattern of results across a broad range of countries, with λ held constant, suggests that λ is lower for countries with better developed credit markets.

In principle, however, one might expect λ_t to vary over time for a number of reasons. In particular, the effects of financial deregulation might be expected to reduce λ_t, the proportion of income going to liquidity-constrained consumers[10]. Thus, if k_t is an index of financial liberalization, we might expect to find λ_t to be a decreasing function of k_t:

$$\lambda_t = \lambda(k_t), \qquad \lambda' < 0 \tag{1.10}$$

Campbell and Mankiw (1991), as well as executing tests of a purely liquidity

[10] Muellbauer and Murphy (1990) and Miles (1992) - among others - also stress the importance of financial liberalization in increasing the marginal propensity to consume in the UK in the mid-1980s.

constrained model ($\lambda=1$), estimate a very simple form of such a model, with k_t set equal to a linear time trend. For the UK, they find some evidence of an increase in λ over the estimation period. This finding is perverse in the sense that it seems to imply that credit rationing tightened over the period 1957-88 in the UK. A similar result is reported by Jappelli and Pagano (1989). Neither of these studies includes a variable real interest rate in the estimating equation, however, and Muellbauer (1994, p. 21) argues that this may to some extent explain this result. More generally, Campbell and Mankiw (1991, p. 753) suggest that their methods, 'applied to aggregate data, are simply not powerful enough to detect movements in λ over time'.

Bayoumi and Koujianou (1990) examine the relationship between liquidity constraints, consumption and financial innovation for six different countries (the US, the UK, Canada, Japan, France and Sweden) all of which have recently experienced financial deregulation, and find a decline in the fraction of liquidity-constrained consumers after the deregulation. Moreover, an approach similar in some ways to the one undertaken in the present study is followed by Bayoumi (1993), who investigates the effect of the process of UK financial deregulation on liquidity constraints and consumption. Using regional data for the UK, Bayoumi (1993) shows both that financial deregulation has been effective in reducing the fraction of aggregate consumption subject to liquidity constraints from around 60 percent before deregulation to around 30 percent by the late 1980s. Bayoumi's regional study is also one of the few studies to find strong statistical evidence of a non-zero intertemporal elasticity of substitution, which suggests that explicit allowance for financial deregulation may be important in this context[11]. Using some non-parametric tests of the optimality of regional consumption patterns in the UK (ie. the absence of an effect of liquidity constraints on consumption), Bayoumi identifies a clear change in UK consumption pattern in the early 1980s, moving decisively towards optimality in the second half of the 1980s.

The present paper may be seen as an extension of the literature reviewed in this section, in that we estimate equation (1.7) while allowing for variation in the real rate of interest and by explicitly modelling λ as a non-linear function of an index of financial deregulation.

1.2.3 The excess smoothness debate and the link to excess sensitivity

An issue receiving large attention by the literature on consumption function is concerned with the magnitude of the response of consumption expenditure to current income, the marginal propensity to consume out of current income. Under the LCPIH, assuming rational expectations, the responsiveness of expected permanent income to current income depends upon the degree of persistence of the income generation process (Muellbauer, 1994)[12].

In fact, the responsiveness of permanent income to current income depends on whether the income process has a unit root. By and large, the literature is

[11] Note, however, that because of the lack of regional data on consumption of nondurables, Bayoumi is forced to use data on total consumption, which includes durable goods. These data, in principle, are expected to be more sensitive to income and the real rate of interest than nondurable goods' consumption, the more usual dependent variable (see Bayoumi, 1993, p. 537).

[12] This issue was first highlighted by Deaton (1987, 1992) and Campbell and Deaton (1989), and is usually recalled as the "Deaton paradox" (see also Deaton, 1992, Chapter 4; and Muellbauer, 1994).

unable to reject the hypothesis that income is a unit root process, while the hypothesis of nonstationarity of income changes can be easily rejected at conventional nominal levels of significance, implying that income is a first-difference stationary process. If income contains a unit root, however, permanent income is expected to respond very significantly to shocks hitting current income. Also, under the Euler equation, if income is difference-stationary and if consumption shocks are not transitory and there are no measurement errors, then the innovation in consumption is equal to the innovation in permanent income, which implies that the variance of the consumption innovation should be comparable in size to the innovation in current income (Muellbauer, 1994). This implication is, however, strongly rejected by the data: consumption patterns appear to be excessively smooth, ie. not volatile as much as they should be if income was a unit root process. A paradox arises, however, as the LCPIH is intended to explain consumption smoothing of agents over time.

The Deaton paradox is not robust, however, to many changes in the underlying model. First, it is well known that, using conventional unit root tests, it is very difficult to distinguish processes that contain a unit root from near-unit root processes (Perron, 1989; Stock, 1990, 1993). This is especially true using the relatively short postwar sample period, if income shocks are very persistent and damp at a very slow rate. Also, not taking into account possible structural breaks due to regime shifts in the series (say shocks of the kind of the 1973 oil shock) is known to bias conventional nonstationarity tests towards the non-rejection region, creating illusory persistence. Overall, the presence of a unit root in the income process is controversial, and describing income dynamics by a single univariate process over a number of years is very questionable (Muellbauer, 1994).

Nevertheless, alternative econometric specifications for the income process have been proposed by researchers. An example is an approach which uses segmented trends, where occasional large shocks shift the slope or the constant in what would otherwise be a deterministic trend (Perron, 1989; Banerjee, Lumsdaine and Stock, 1992). Also stationary nonlinear models - eg. threshold autoregressive models and smooth transition autoregressive models - are good candidates for fitting the income process (see Tong, 1990; Granger and Teräsvirta, 1993; Teräsvirta and Anderson, 1992; Teräsvirta, 1994, 1995). Muellbauer and Murphy (1993) use two alternative econometric specifications for fitting US data on real personal disposable income and provide, with both approaches, strong evidence against the hypothesis that income is a difference-stationary variable. The interpretation suggested by the authors is that there is an upper boundary to the output pattern in the short-run, caused, for example, by a slowly evolving supply side, and also there are forces, both endogenous and induced by macroeconomic policy changes, which stimulate recovery if income falls below its long-run trend. If this interpretation is correct, then a rational consumer will tend to be more pessimist about future growth if growth has been relatively high lately, and *vice versa*. Also, this argument can also explain the excess smoothness of consumption, because a very large proportion of income shocks would not be permanent and, therefore, the Deaton paradox does not arise.

Nevertheless, in principle, the existence of liquidity constraints can explain *also* the excess smoothness phenomenon. Consider again the Campbell-Mankiw model, derived assuming that the random walk model of consumption holds only for unconstrained consumers, whose consumption expenditure may be described by the process $c_t^u = c_{t-1}^u + e_t$ where e_t is white noise. Credit-constrained consumers, however, consume out of their current income, so that their consumption pattern is described by the equation $c_t^c = y_t^c$. Using first differences, the following equation for the change in consumption may be derived:

$$\begin{aligned} \Delta c_t &= (1-\lambda)\,\Delta c_t^u + \lambda \Delta c_t^c \\ &= (1-\lambda)\, e_t + \lambda \Delta y_t^c \end{aligned} \tag{1.11}$$

which corresponds to equation (1.7) assuming a constant value of λ and a zero intertemporal elasticity of substitution of consumption. Equation (1.11) establishes a direct link between excess sensitivity and excess smoothness, in the sense that excess smoothness simply follows from the fact that, given (1.11), changes in consumption are forecastable as the consumption innovation is equal to a weighted average of two elements: the innovation in permanent income for the consumers who follow the LCPIH (e_t) and the current income innovation for the liquidity-constrained consumers (say v_t). This implies that, if e_t and v_t are not highly correlated and the variance of e_t is not greater than the variance of v_t, then it is very likely that the variance of the consumption innovation is smaller than the variance of the income innovation: again no Deaton paradox arises (Campbell and Mankiw, 1991; Deaton, 1992; Muellbauer, 1994).[13] [14]

1.2.4 Further issues beyond the rational expectations permanent income hypothesis

Most studies following Hall's seminal paper (1978) focus on the implications of relaxing one or more of the underlying assumptions of the rational expectations permanent income hypothesis (REPIH). Some of the issues have already been widely discussed, such as liquidity constraints and precautionary saving in the face of income uncertainty. Without aiming to provide a review of the enormous literature on the consumption function, some other issues examined by the relevant literature are briefly described for completeness, even if they are relatively less directly related to the main focus of this study, which is concerned with the effects of financial liberalization on liquidity constraints and hence on consumption expenditure.

First, the analysis conducted until now assumes separability of individuals' utility functions between consumption and leisure. If this assumption does not hold,

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13 Some researchers argue that also precautionary behaviour of consumers in the face of income uncertainty may provide an alternative feasible explanation of the excess smoothness of consumption. This is because 'a large positive income shock is likely to be taken as an indicator that there is more uncertainty about the new, higher level of income, therefore leading to a smaller rise in consumption than in the absence of a precautionary motive' (Muellbauer, 1994).

14 Also the relaxation of the assumption of no habits and no adjustment costs, used in deriving the Hall model of consumption, strongly affects the implications of the theory. Habits and convex adjustment costs imply partial adjustment of consumption to life-cycle wealth, providing an explanation of excess smoothness, since any period's shocks are only partially adjusted within the period. Nevertheless, the empirical evidence appears to suggest that these factors cannot fully explain the failure of the random walk consumption model (Muellbauer, 1988, 1994).

then consumption decisions depend on permanent income and also on leisure and leisure expectations; that is to say, if separability is violated, then, in each period, utility functions depends on both consumption and leisure. The derived Euler equation, under this framework, implies a negative relationship between changes in consumption expenditure and changes in unemployment, providing another possible explanation of the excess sensitivity phenomenon (Heckman, 1974; Bean, 1986; Deaton, 1992).

Another possible explanation of the failure of the REPIH may be the inadequacy of the assumption that expectations are formed rationally, even if it is difficult to suggest valid alternatives. First, the validity of the rational expectation (RE) hypothesis is very controversial. Second, if the RE hypothesis is in fact correct, the estimation of many consumption function executed by the literature is subject to the well known Lucas (1976) critique to econometric models, unless one constructs a model of consumption which includes only "deep structure", regime-invariant parameters, such as - in this context - the intertemporal elasticity of substitution. In addition, of course, parameters which are expected to be regime-variant may be included in a consumption function without assuming their constancy. This is the case, in fact, in the consumption function proposed in this study, where the parameters estimated are the time-varying parameter λ_t, representing the proportion of income going to liquidity-constrained consumers, and the intertemporal elasticity of substitution σ.[15]

Nevertheless, Muellbauer (1994) argues that the marginal propensity to

[15] Clearly, estimations of the original Campbell-Mankiw model of consumption are also subject to a Lucas critique problem if the proportion of income going to liquidity constrained consumers is assumed to be constant.

consume is less for illiquid assets than for liquid ones. Muellbauer's argument is the following. Households typically hold some liquid assets, which may be easily converted in expenditure, and illiquid assets, usually yielding a relatively higher rate of return. Among a range of assets with different liquidity characteristics, housing, pension funds and life insurance funds may reasonably be considered as the most illiquid assets. Nevertheless, if illiquid assets may be used as collateral for borrowing money, then the market value can be accessed to a certain degree. The key insight is that differences in liquidity create different spendability weights associated with different types of assets and debts. Muellbauer (1994) argues that financial deregulation increases the spendability weights on illiquid assets; rather, the increased spendability of illiquid assets may be the most important implication of financial liberalization. This would be consistent with the empirical evidence suggesting an upward trend of the proportion of liquidity constrained consumers over time in the UK, even during the period in which the financial deregulation was fully operating in the UK (Jappelli and Pagano, 1989; Campbell and Mankiw, 1991). As discussed below, however, we find the converse result: the model proposed in this study suggests that the proportion of income going to liquidity constrained consumers (λ) in the UK and in France has declined over time, as an effect of the undergoing process of financial deregulation. Previous empirical results suggesting that λ has followed an upward linear trend in the UK and that has not significantly changed in France - intuitively implying a tightening of UK financial policies during the late 1980s and no significant changes in French financial policies during the same period - may simply be due to the fact that the history of financial deregulation in France and, to a greater extent, in the UK, are too complex to admit a simple monotonic behaviour of λ.

1.2.5 The macroeconomics and the microeconomics of the consumption function

Econometric tests of the LCPIH based on Euler equations using aggregate per capita data generally tend to reject the LCPIH. Nevertheless, aggregation issues are very serious and, according to some researchers, not sufficiently taken into consideration by the empirical literature on the Euler equation - eg. see Attanasio (1995), Deaton, (1992), Muellbauer (1994).[16] Consider, for example, a chain of entering households overlapping existing households: aggregate consumption is carried out by different groups of people on different times and there is no reason for which aggregate consumption on a particular date should be related to aggregate consumption on other dates. For example, there is not necessarily a relationship between consumption of the young at a time t+1 and consumption of the old at time t, even if the Euler equation implies a relationship between the consumption of the young at time t and the consumption at time t+1 of the same people, who are one-period older. Important contributions on the implications of aggregation in the consumption function literature are due to Gali (1990, 1991) and Clarida (1991) who demonstrate that a world of finite-lived households ought be expected to generate a consumption pattern exhibiting both excess sensitivity and excess smoothness in the aggregate. This implies that the empirical finding of excess sensitivity and excess smoothness of consumption using aggregate per capita data may be consistent with the LCPIH rather than indicating its rejection.

In particular, Attanasio (1995) argues that the empirical results provided by

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[16] See also Obstfeld and Rogoff (1996, Chapter 3).

Campbell and Mankiw (1989) and similar subsequent studies are questionable, as some of the assumptions required may not be plausible. These assumptions are that (a) consumption and leisure are separable in the utility function, (b) the proportion of income going to liquidity constrained consumers, λ, is constant over time and (c) liquidity constrained consumers always consume their labour income in both an up-swing and down-swing of the business cycle. Even if there is some evidence against the hypothesis of separability between consumption and leisure (eg. Browning and Meghir, 1991), separability is generally assumed in most of the relevant literature as its relaxation generates high analytical difficulties. The second assumption is certainly not plausible and the importance of a Lucas critique problem has already been stressed in this context. The model we estimate in this study, however, is not affected by this problem since we relax the assumption that λ is constant over time. The third assumption, also commonly used in the consumption function literature, even if it may not hold exactly, is not expected to have a very significant effect (Deaton, 1992).

Overall, while the results from estimating the Campbell-Mankiw consumption function should be taken with caution, they may still provide important information on the determinants of consumption if the models are constructed in a way which is robust to the Lucas critique. Moreover, as noted by Attanasio (1995), given the very small number of observations provided by the micro data available, estimation of consumption models using aggregate data are 'extremely useful' for practical purposes and particularly for forecasting.

Many empirical applications have been carried out on micro data, mainly using the Panel Study of Income Dynamics (PSID)[17] or the Consumer Expenditure Survey (CEX) for US households and the British Family Expenditure Survey (BFES) for British households. The results from studies of Hall and Mishkin (1982), Zeldes (1989a), Jappelli and Pagano (1989) have already been discussed above. In addition to those results, recent studies using the CEX (Lusardi, 1996; Attanasio and Weber, 1995) and BFES (Attanasio and Browning, 1995; Attanasio and Weber, 1993; Blundell, Browning and Meghir, 1994) report quite favourable results to the Euler equation.

Attanasio and Weber (1993) also argue that the use of improperly aggregated data may explain the rejection of the Euler equation usually reported using aggregate data. They show that, when aggregation is based on arithmetic means, the Euler equation is rejected on aggregate data while, when aggregation is executed in a theory-consistent way, the Euler equation is not rejected[18]. The empirical evidence on the relationship between consumption changes and income changes using micro data at relatively short frequencies mainly suggests that excess sensitivity of consumption is relatively small, if one allows for the effects of demographic and labor-supply variables on the marginal utility of consumption.

The empirical results provided by studies using micro data, however, should also be taken with much caution for a number of reasons. First, the introduction of demographic and labor-supply variables may introduce noise which reduces the power of the excess sensitivity tests (Attanasio, 1995). Second, excess sensitivity

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[17] Other empirical applications on the PSID are due to Altonji and Siow (1987), Runkle (1991) and Keane and Runkle (1992).

[18] For a general survey of the literature on aggregation, see Stocker (1993).

may be important at relatively low frequency (life-cycle) and, therefore, the relatively high frequency (business-cycle) used in empirical studies based on micro data may simply not be sufficiently powerful to detect excess sensitivity. In particular, this argument is consistent with the empirical evidence provided by Carroll and Summers (1991), who estimate age profiles for both consumption and income for various population groups (disaggregated by occupation and education) for a number of countries and find that 'consumption parallels income', in contrast to the prediction of the LCPIH.

Moreover, additional doubts on the reliability of the results provided by empirical studies using micro data arise because of certain characteristics of the data sets. The PSID only provides data on food consumption (about 17 percent of total consumption in the National Accounts) and, in fact, empirical work on the PSID requires strong assumptions, such as separability in preferences of food consumption and other forms of consumption. Also, the CEX is a rotating panel of households in operation since 1980, where households are visited five times during a fifteen-months period and are asked about income in two of the visits and about consumption in all but the first visit. Using the CEX, therefore, time-series variation is usually replaced by cross-sectional variation and this requires additional and 'dangerous' (Deaton, 1992) assumptions. The BFES provides, however, time-series of independent cross-sections of annual data on some 7,000 households a year since 1954. Note that no household or group of households is observed more than once. Methods for constructing a usable data set are provided by Browning, Deaton and Irish (1985) and Deaton (1985), but the econometrics becomes very complex and the results ambiguous because of the lack of a long-panel data, which represents the ideal data set for testing the LCPIH. Moreover, large measurement errors may be present in these data sets, affecting significantly the estimation results (see Deaton, 1992, p. 144).

Overall, therefore, while a well-constructed micro data set represents the ideal data set for testing the LCPIH, it is very questionable that currently available micro data sets may be very useful for this purpose.

1.3 The process of financial deregulation

In this section we provide a brief discussion of the process of financial regulation and deregulation in the UK and in France over the past twenty-five years or so, as a motivation of our contention that one would have expected the proportion of income going to liquidity-constrained consumers to have diminished over time as binding liquidity constraints on individuals and households became less common.

1.3.1 The UK experience[19]

One of the first set of financial reforms in the UK during the period under consideration was the introduction of "Competition and Credit Control" (CCC) in 1971. CCC removed the distortions induced by a number of impediments previously characterising the banking sector: ceiling controls on bank lending in sterling, the Clearing Bank Cartel, which was in operation since the mid 1930s, and

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[19] Hall (1991) and Llewellyn (1991) provide discussions of financial deregulation in the UK, while OECD (1987) and Mathieson and Rojas-Suarez (1990) provide a general summary of financial liberalization in industrial countries; see also Bayoumi (1990).

the previously agreed cash and liquidity ratios of the clearing banks. Changes in dealing tactics in the gilt-edged market were also introduced in order to limit the extent to which the Bank of England could create money through its operations in that market. Also a reserve ratio of 12.5%, to be observed by all banks, and the power to call special deposits from the Bank of England were implemented in order to increase the effectiveness and the predictability with which the authorities' open market operations would impinge on bank lending and the money supply. This package of reforms had the objective of increasing competition in the banking sector and of enhancing the authorities' ability to control the money stock through market rather than "direct" control mechanisms.

The effect of CCC was partly offset, however, by three actions. First, the "Supplementary Special Deposits" scheme (SSD) - the "Corset" - was introduced in 1973. In that year, significant increases in nominal interest rates occurred and several calls for special deposits were made. The SSD was designed to restrain the growth of the money supply, in some sense moving back to direct controls, since non-interest bearing deposits were called on a bank-by-bank basis whenever the growth of their balances exceeded a certain rate. Hall (1991) argues that the corset was particularly effective during its final phase of operation, June 1978 to June 1980. Second, hire purchase terms control, abolished under CCC, was reintroduced in December 1973 and remained in force until July 1982. Third, in 1978, the market-related formula for deriving the minimum lending rate (MLR), which replaced the Bank's related rediscount rate (bank rate) in 1972, was removed. From 1978, the MLR began once again, therefore to be determined by administrative action by the monetary authorities.

A return to the initial spirit of the CCC was signalled by the publication of the Green Paper on monetary controls in 1980 (HMSO, 1980), where the deficiencies of the current financial system and an alternative "reregulating" approach were emphasised.

In August 1981, the MLR was suspended. Also, the Bank of England explicitly adopted an essentially *laissez-faire* approach to the prudential regulations of the banks. The new deregulating - or, perhaps better, self-regulating - spirit in the banking system is clearly reflected by the Banking Act in 1983, which took effect from 1987.

Furthermore, the legislation of the Building Societies was strongly amended during the 1980s, after a Green Paper (HMSO, 1984), which relaxed the legal restraints under which building societies operated, subject to the general rule 'that the building societies continue primarily in their traditional roles.' These proposals were largely confirmed in the Building Societies Bill, which received Royal Assent in July 1986. They allowed the building societies to enter, subject to certain restrictions, into a wide range of financial activities. This move was given further impetus by the 1987 Financial Services Act which, among other things, allowed building societies to compete more effectively in the retail financial services market place.

1.3.2 The French experience

Quantitative controls on credit lasted in France for more than one decade (Pecha and Sicsic, 1988). In November 1972, advertisements for personal credits were forbidden and the length of credit agreements was limited to two years at most. All bank credits were subject to a maximum sliding yearly growth rate in 1973 and 1974. Credit to consumers was not even allowed to exceed its level reached in early July 1973. This measure was cancelled only within the expansionary fiscal policy initiated in September 1975. Over the 1976 through 1982 period, the growth of consumer credit was limited to the same norm as all other types of bank credit, within the credit control procedure. More restrictive measures, similar to those implemented over the 1973-75 period were again implemented for a two-year period beginning in late November 1982. In early February 1983, however, limits were lifted for consumer credit assigned to the financing of a specific good. It proved difficult to apply in practice a distinction between assigned and unassigned credit. In early December 1984, the decision taken in 1982 was cancelled.

During that period, the absence of respect of the ceilings on the growth of bank credit was sanctioned by "supplementary" required reserves with the Banque de France (Patat, 1993). These reserves were proportional to excess credit creation and were a rapidly increasing function of it. Some flexibility, however, was given to this system. For instance, when a bank did not reach its ceiling during a given month, it could "capitalise" this gap and use it over the following semester to exceed the ceiling. In addition, banks could trade their gaps between themselves during a given month. Moreover, banks were enabled to subtract from the credits they granted the increase in their permanent capital (own funds and banks' bonds issued net of bonds acquired).

The very rapid increase in consumer credit between 1986 and 1989 resulted both from a catching up after the constraints its growth faced during the credit control period and from an active strategy pursued by banks in search of market share.

1.3.3 Concluding remarks

Overall, the last twenty-five years or so for the UK and France have been characterised by two complementary trends of financial deregulation and prudential regulation, although the pace has accelerated markedly in the mid-1990s. The competition allowed by this deregulation has also stimulated the development of financial innovations, contributing positively to increase both structural and allocative efficiency. As one would expect, increased deregulation and competition in the financial sector has contributed to a reduction in excess demand for financial products, with allocation of finance becoming more determined by the price mechanism than by rationing. *A priori*, one would expect this process to have affected the proportion of consumers in the economy who face liquidity constraints of one kind or another.

1.4 Data

Quarterly data were obtained for the sample period 1963Q1-1994Q4 for the UK from *Economic Trends* (Central Statistical Office) and for the sample period 1970Q1-1993Q4 for France from the *INSEE Quarterly National Accounts* for the following variables: real and nominal personal disposable income, the consumption deflator and consumption expenditure on non-durables and services. Quarterly data on three UK short term interest rates - the three-month Treasury Bill rate, the deposit rate and the bank rate - and quarterly data on three French short term interest rates - the three-month Interbank rate, the Pibor rate and the money market rate - were taken from the International Monetary Fund *International Financial Statistics* data base. Data on population - used to construct a per capita series for consumption - were taken from the OECD *Main Economic Indicators*. Data on consumer credit were taken from *Economic Trends* (Central Statistical Office) for the UK and from the Banque de France for France in order to construct an index, k_t, of the total stock of consumer credit outstanding as a proportion of personal disposable income. The actual quarterly percentage change in the consumption deflator was subtracted from each of the nominal interest rates (expressed on a quarterly basis) to construct measures of *ex post* real interest rates (expressed in decimals). For the UK, the real deposit rate (r_t) was used in the regression equations with lags of the real bank rate (i_t) and the real treasury bill rate (tb_t) included in the instrument set (discussed below). For France, the real money market rate (r_t) was used in the regression equations with lags of the real interbank rate (i_t) and the real Pibor rate (tb_t) included in the instrument set[20].

1.5 Estimation and empirical results[21]

In estimating the Campbell-Mankiw consumption function, even under the assumption of constant λ, the ordinary least squares (OLS) estimator will in general be biased and inconsistent, since the error term in (1.7) is orthogonal to lagged variables, but not necessarily to Δy_t. In constructing a suitable instrumental

[20] The choice of the sample period for both countries considered was dictated by data availability for consumer credit outstanding.

[21] In what follows, we apply a five percent nominal significance level to all test procedures, unless explicitly stated otherwise.

variables (IV) estimator, any lagged stationary variable, which is correlated with Δy_t is a valid instrument, since it is by construction orthogonal to the error term in (1.7). If no good instrument is found, ie. no instrument is found which is orthogonal to ϵ_t yet correlated with Δy_t, then the procedure breaks down since permanent income is equal to current income and so λ_t is unidentified (Campbell and Mankiw, 1990). The choice of the instrument set is, therefore, crucial. Lagged values of Δy_t are usually considered in the literature. Lagged values of c_t cannot be used as instruments because the series is typically non-stationary, although the cointegrated (Engle and Granger, 1987) series $(c_t - y_t)$ may be used[22]. When lags of Δy_t, Δc_t and $(c_t - y_t)$ are included as instruments, then the instrument set is tantamount to the explanatory variables of a consumption simple error correction model (ECM) of the form suggested by Davidson, Hendry, Srba and Yeo (1978), Davidson and Hendry (1981) and Hendry and von Ungern-Sternberg (1981)[23]. A test of the overidentifying restrictions, given this information set, is therefore tantamount to testing the estimated model against the alternative of a general error correction model. In our empirical work we also include additional lagged real interest rates in the instrument set, since these are often found to be powerful in forecasting changes in income (Fischer and Merton, 1984; Litterman and Weiss, 1985; Sims, 1980).

Another potential problem which is encountered in estimating Euler equations for consumption is that consumption and income are measured as

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[22] Preliminary unit root tests on the present data set indicated that the stochastic processes generating r_t, Δy_t, Δc_t and $(c_t - y_t)$ are all apparently I(0).

[23] See also Patterson (1984a,1984b, 1987).

quarterly averages rather than at points in time. If the representative agent's decision period is shorter than the frequency of observation of the data (in the present study one quarter), then, even under the assumption of a constant real interest rate or a zero intertemporal elasticity of substitution, measured consumption would be the time average of a random walk. As originally noted by Working (1960), this will induce positive first-order moving average (MA) serial correlation (see also Patterson and Pesaran, 1992). In principle, therefore, consistent estimation requires an instrument set with at least a two-period gap between the regressors and the instruments.

Patterson and Pesaran (1992) suggest that it may be more appropriate to estimate the first-order MA process in the disturbance parametrically and test for its significance. This is because there may be a substantial loss in quality of the instrument set by lagging twice and this may be avoided if the first-order serial correlation is not empirically significant (Nelson and Startz, 1990a,b; Maddala and Jeong, 1992; Staiger and Stock, 1994; Patterson and Pesaran, 1992). In addition, the argument of Campbell and Mankiw (1991) that an MA process may be induced because of small elements of durability in the category of consumption labelled non-durables and services does not hold with the same force in the present investigation since in most of our estimations we allow time-varying real interest rates (Mankiw, 1982). Following Patterson and Pesaran (1992), therefore, we explicitly test for the presence of first order MA terms in the disturbance using an instrumental variables-moving average estimation (IVMA) procedure (Pesaran, 1990; Patterson and Pesaran, 1992).

In addition, we allow for heteroscedasticity in the errors using a Hansen

(1982) method-of-moments adjustment to the covariance matrix[24]. Note that heteroscedasticity would be expected if λ_t is indeed time varying - see (1.8).

1.5.1 The liquidity-constrained model assuming a constant real rate of interest

In Table 1.1 we report the results from the estimation of the pure liquidity constrained consumption model, assuming a constant value of λ and a constant real rate of interest, which is then absorbed into the constant term. The instrument set used corresponds to an error-correction model for consumption. In order to give an indication of the quality of the instrument set we report the adjusted R^2 from the first-stage regression and the p-value of a Wald test of the null hypothesis that the coefficients in the first-stage regression, excluding the constant, are jointly zero[25].

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[24] In the case of heteroscedasticity, Hansen (1982) shows that it is possible to compute consistent estimators for the covariance matrix in a wide range of situations using a procedure that imposes little structure upon the covariance matrix. Given a regression model $Y=X\beta+u$ with covariance matrix $V=E(uu')=\sigma^2 I$ and defined a matrix of instruments Z, the instrumental variables estimator is:

$$(X'P_Z X)^{-1} X'Z (Z'Z)^{-1} mcov(Z, u) (Z'Z)^{-1} Z'X (X'P_Z X)^{-1}$$

where $P_Z=Z(Z'Z)^{-1}Z'$ and

$$mcov(Z, u) = \sum_{k=-i}^{i} \sum_{t=1}^{T} u_t Z_t' Z_{t-k} u_{t-k}$$

with u_t denoting the residual at time t (see also White, 1980).

[25] Note that we do not in any way attempt to optimise these measures of instrument quality since it can be demonstrated that 'such pre-estimation screening may exaggerate rather than alleviate the size distortion described by Nelson and Startz (1990a,b). In essence, this error arises because those instruments that are identified as having high relevance for the regressors in the sample are also likely to have higher endogeneity in the sample' (Hall, Rudebusch and Wilcox, 1994).

We also report a Sargan test for the validity of the instruments which is in effect a test of the over-identifying restrictions of the model (Sargan, 1958). This is asymptotically distributed as $\chi^2(d)$ under the null hypothesis of no misspecification, where d=m-n, m being the number of instruments and n the number of endogenous variables on the right-hand side of the regression. In the present context, rejection of the over-identifying restrictions is tantamount to a rejection of the estimated model against the alternative of an error correction consumption model[26]. The final column lists, where appropriate, the asymptotic t-ratio for the first-order moving average coefficient given by instrumental-moving average error estimation of the model (Patterson and Pesaran, 1992). Ordinary least square regressions are reported for purpose of comparison only, since they are likely to be biased and inconsistent.

The results for the UK and for France are reported in Panel A and Panel B respectively. Panel A of Table 1.1 shows that the estimated value of λ for the UK,

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[26] Given an instrument set Z, if Z is independent of the structural error term in the equation, say ϵ, the regression of ϵ on Z is expected to yield a relatively low R^2. Using the residuals estimated by instrumental variables, say ϵ_{IV}, in place of the structural ones (unobservable), the R^2 from the regression of ϵ_{IV} on the instruments is used to construct the Sargan test:

$$Sargan = (T - k)\, R^2 \sim \chi^2(d)$$

where T and k denote the number of observations and the number of parameters in the structural equation respectively. An alternative, equivalent formulation of the Sargan test statistic, commonly used by the literature, is:

$$Sargan = (\epsilon_{IV}' P_Z \epsilon_{IV}) / s^2$$

where P_Z is the projection matrix of the instruments $[=Z(Z'Z)^{-1}Z']$ and $s^2=(\epsilon_{IV}'\epsilon_{IV})/(T-k)$.

using IV estimation, is significantly different from zero and indicates a proportion of income going to liquidity-constrained consumers of around 21%. The results are qualitatively unchanged when IVMA is used, and the estimated first-order MA coefficient is insignificant, indicating that single lags of the instruments are adequate. Most tellingly, however, the Sargan test statistic indicates a strong rejection of the over-identifying restrictions, indicating rejection of purely liquidity-constrained consumption model against the ECM alternative.

As Panel B of Table 1.1 shows, using French data IVMA estimation yields a significant error structure, suggesting that once-lagged instruments are inappropriate. Therefore we estimate the model by using a twice-lagged instrument set. The estimated value of λ for France, using the twice-lagged instrument set, is significantly different from zero and indicates a proportion of income going to liquidity-constrained consumers of around 36%. Also, the model, estimated with a twice-lagged instrument set, cannot be rejected against the error correction alternative using the Sargan test statistic.

1.5.2 The model with only unconstrained consumers

Table 1.2 contains the results of estimates of the log-linearised Euler equation (1.6), under the assumption of no liquidity constraints ($\lambda=0$). The results for the UK and for France are reported in Panel A and Panel B respectively.

As Panel A of Table 1.2 shows, there is again no indication of statistically significant first-order moving average effects in the disturbance, using the UK data. These results, with estimates of σ significantly different from zero, provide *prima facie* evidence against Hall's (1988) contention that the random walk model of consumption is correct because the elasticity of substitution, σ, is close to zero. Nevertheless, the Sargan test again rejects the model against the ECM alternative.

Panel B of Table 1.2 shows, however, indication of a statistically significant first-order moving average effect in the disturbance, using French data. Therefore, we reestimate the model by IV using a twice-lagged instrument set. Interestingly, for France, the results, with estimates of σ not significantly different from zero, provide evidence supporting Hall's (1988) random walk model. Also for France, however, the Sargan test indicates suggests the rejection of the model against the ECM alternative.

1.5.3 The model with both liquidity-constrained consumers and unconstrained consumers

In Table 1.3 we report the results of estimation of the full Campbell-Mankiw model, equation (1.7), assuming a constant value of λ. The results for the UK and France are reported in Panel A and Panel B respectively.

As shown in Panel A of Table 1.3, for the UK IVMA estimation again suggests that no significant first-order autocorrelation of the error term is present. Although the estimated value of λ is not strongly statistically significantly different from zero, θ [$=(1-\lambda)\sigma$] is strongly significant, implying a value of σ significantly different from zero. Again the evidence is against the hypothesis that changes in consumption are unpredictable. Once again, however, the Sargan test statistic indicates a strong rejection of the model against the ECM alternative.

In Panel B of Table 1.3, using French data, IVMA estimation suggests that significant first-order autocorrelation of the error term is present so that the results obtained using an instrument set with at least two lags are relatively more appropriate. These indicate an estimated value of λ strongly significantly different from zero, which is also very similar in size (0.344) to the results reported in Table 1.1 for the pure liquidity-constrained consumers model; the estimated value of σ, however, is not significantly different from zero[27]. Also, using the Sargan test statistic, the model cannot be rejected against the ECM alternative.

1.5.4 The liquidity-constrained model with time-varying coefficients

Our final exercise involves allowing the proportion of income going to liquidity constrained consumers, λ_t, to vary over time. Specifically, because of the ongoing process of financial deregulation both in the UK and in France, one might expect λ_t to have diminished over the sample period although - given our discussion in Section 1.3 - not necessarily in a monotonic fashion. Following Bayoumi (1993), we take the stock of consumer credit outstanding as a proportion of personal disposable income (normalised to unity at its initial value over the estimation period) as an index of the level of financial deregulation, k_t. This leads to estimation of the following simple non-linear model:

$$\Delta c_t = \lambda_t \Delta y_t + (1-\lambda_t)\sigma r_{t-1} \tag{1.12}$$

where λ_t is forced to lie in the closed interval [0,1] with its value in between

[27] This result is in line with what Campbell and Mankiw (1991) obtained for France over the 1970-1988 period.

postulated to be approximated by a non-linear power function of k_t[28]:

$$\lambda_t = [1 - (\phi_0 + \phi_1 k_t + \phi_2 k_t^b)] \tag{1.13}$$

The general non-linear formulation of λ_t is expected to capture more accurately overlapping financial regulatory and deregulatory regimes. Also, the model (1.12)-(1.13) nests the simple linear model of λ_t more usually employed in the literature when ϕ_2 or b are not found to be statistically significantly different from zero. Thus, allowing λ_t to be a non-linear function of an index of financial deregulation is a means of exploring whether previous studies which have modelled this variable linearly have been unduly restrictive.

The results of estimating this model by non-linear instrumental variables are reported in Table 1.4. Once again, the results for the UK (where a once-lagged instrument set is used) and for France (where a twice-lagged instrument set is used) are reported in Panel A and Panel B respectively.

As shown in Panel A of Table 1.4, for the UK the coefficient ϕ_0, always found to be insignificantly different from zero, is restricted to this value in our final

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[28] Note that all of the variables entering the model (1.12)-(1.13) are realizations of stationary processes, since y_t and c_t are difference-stationary series (ie. Δy_t and Δc_t are stationary) r_t is stationary, (see footnote 22) and k_t must have a finite variance since it is bounded between zero and unity. This is important since, as shown by Stock and West (1988), in the presence of unit roots in consumption and income, standard regression models relating these series will only result in consistent parameter estimates if the equation can be reparameterised so that the coefficients of interest are on stationary variables. This would be a particular problem in a non-linear model unless, as here, the variables already enter in stationary form.

results[29]. The estimated values of ϕ_1 and ϕ_2 are, instead, strongly jointly significant. Although ϕ_2 is not individually strongly significant, restricting ϕ_2 to zero results in the estimated value ϕ_1 becoming insignificant. Therefore, it seems appropriate to interpret the joint significance of ϕ_1 and ϕ_2 as a test of the significance of the non-linear power function, $\lambda_t = [1-(\phi_1 k_t+\phi_2 k_t^b)]$, with an estimated value of b of 3.1, itself strongly significantly different from zero. The estimated value of the intertemporal elasticity of substitution, σ, is also strongly significantly different from zero, with a point estimate of 0.15.

We also tested for the over-identifying restrictions of the model using a generalization of the Sargan test statistic, due to Hansen (1982). The Hansen test statistic indicates that the model cannot be rejected against the ECM alternative at the five percent nominal significance level.

Table 1.4 also reports the results of tests to ensure that neither serial correlation nor heteroscedasticity are present in the residuals.

Overall, therefore, the estimation results for the UK are extremely encouraging. We have uncovered evidence of significant time variation in λ which, when explicitly modelled as a non-linear function of the degree of financial deregulation, results in a statistically significant, plausible estimate of the elasticity of intertemporal substitution, σ, in an empirical model which cannot be rejected against the most popular and widely used alternative specification for UK consumption behaviour - the error correction specification.

As shown in Figure 1.1, which graphs the estimated trajectory of λ_t, the

[29] Estimating the model with the exclusion of ϕ_0 implies that λ_t cannot be constant any longer. The estimates obtained including ϕ_0 in the non-linear power function are, however, qualitatively invariant.

estimated proportion of income going to liquidity-constrained consumers is highly sensitive to changes in borrowing constraints. The estimated trajectory of λ_t suggests that, while the importance of liquidity constraints has declined rapidly during the last decade, it has been of varying importance over the full estimation period. In particular, the effect of Competition and Credit Control in the early 1970s appears to have been largely offset, as discussed in Section 1.3.1, by the reintroduction of hire purchase terms control, the reintroduction of administered interest rates and, perhaps most importantly, the introduction of the corset. The rise of λ_t during the late 1970s accords with Hall's view (1991) that the corset was particularly effective over this period. On the other hand, the relaxation of the corset in 1980, the abandonment of hire purchase terms control in July 1982, and the series of regulatory reforms in the retail financial services sector culminating in the 1987 Financial Services Act, all appear to have led to a rapid decline in λ_t from the early 1980s, becoming zero towards the end of 1988.

The estimated path of λ_t is also instructive for understanding the results obtained by previous researchers. In particular, the observed variation in λ_t - and especially its rise during the late 1970s - may account for why some researchers have detected an upward linear trend in its path (Campbell and Mankiw, 1991; Jappelli and Pagano, 1989). As discussed in Section 1.3.1, however, and as borne out by Figure 1.1, the history of financial deregulation in the UK is probably too complex to admit such simple monotonic behaviour.

In order to make a comparison with Bayoumi's (1993) study of patterns of UK regional consumption, we also estimated the non-linear model (1.12)-(1.13) for total consumption, even though total consumption is not the most appropriate category of consumption expenditure for our purposes since it includes expenditure on durables. The results are, in fact, very similar to the results obtained using nondurable consumption as the dependent variable, with the estimated ϕ_1 and estimated ϕ_2 equal to 2.09 and -6.42 respectively, estimated b equal to 3.09 and the estimated σ to 0.19 (all parameters are statistically significantly different from zero at the five percent nominal level of significance). The Hansen test indicates that the model is not rejected against the ECM alternative. Interestingly, therefore, the time varying path of λ_t is very similar to the one displayed above for the results using nondurable consumption and is fairly consistent with the results of Bayoumi's (1993) regional study. For example, the estimated value of λ_t is just below 50 percent in the early 1980s (Bayoumi suggests around 60 percent), 32 percent in 1985 and 26 percent at the beginning of 1987 (Bayoumi suggests about 30 percent). These results are also consistent with Bayoumi's conclusion that a clear change in UK consumption pattern occurs during the 1980s, moving decisively towards the absence of liquidity constraints during the latter half of the 1980s.

Turning to the results for France, reported in Panel B of Table 1.4, again the coefficient ϕ_0 is always found to be insignificantly different from zero and therefore restricted to this value in the final reported results. Using the French data, however, we also restrict to zero, in the final estimation, the intertemporal elasticity of substitution, σ, which was always found to be very low (about 0.008) and strongly insignificantly different from zero at conventional nominal levels of significance. The estimated values of ϕ_1 and ϕ_2 are, instead, individually statistically significant and the estimated value of b is 1.550, itself strongly significantly different from zero.

Also for France, the Hansen test statistic indicates that the model cannot be rejected against the ECM alternative at the five percent nominal significance level[30]. Finally, the last three columns of Panel B of Table 1.4 ensure that there is no presence of serial correlation or heteroscedasticity in the residuals.

The results for France are also satisfactory, even if less encouraging than the results for the UK. For France, in fact, while there is strong evidence of significant time variation in λ and the model cannot be rejected against the alternative specification of an error-correction model for French consumption, the estimated intertemporal elasticity of substitution is very low and insignificantly different from zero.

As shown in Figure 1.2, which graphs the estimated trajectory of λ_t for France, the importance of liquidity constraints has declined rapidly during the last decade, while being of varying importance over the full estimation period.

The effect of very severe quantitative limits - discussed in Section 1.3.2, forbidding the growth of consumer credit in 1973 and 1974 is reflected in the sharp rise of λ_t over this period. In the late 1970s, less severe restrictions, allowing some growth in consumer credit outstanding, may explain the fact that the estimated proportion of income going to liquidity constrained consumers was slightly reduced. The gradual removal of these restrictions over 1983 to 1985 is exactly consistent with the plummeting estimated value of λ_t in the second half of the 1980s.

In addition, the estimated path of λ_t is very instructive for understanding the results obtained by previous researchers using French data. In fact, the observed

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[30] As in all previous estimated models, strong evidence of MA(1) serial correlation in the residuals is present when the IV estimation is based on a once-lagged instrument set.

variation in λ_t may account for why some researchers have detected no *significant* change in its path (Campbell and Mankiw, 1991). As discussed in Section 1.3.2, however, and as borne out by Figure 1.2, the history of financial deregulation in France is probably too complex to admit the simple monotonic behaviour considered by previous research in this area.

1.6 Conclusions

In this study we have examined the effect of movements in the real interest rate on changes in real consumption for the UK over the period 1963-1994 and for France over the period 1970-1993. In particular we tested the random-walk representative agent consumption model of Hall (1978, 1988) as well as a more general model which allows for some consumers to encounter liquidity constraints. Our main conclusions are as follows.

First, we find no evidence for the UK but strong favourable evidence for France to support the hypothesis that the intertemporal elasticity of substitution is zero and that the real rate of interest and changes in consumption are, therefore, orthogonal.

Estimating a form of the consumption function due originally to Campbell and Mankiw (1989), but with λ held constant, we were easily able to reject the model against an alternative error correction specification for the UK, not for France. Modelling λ_t as a non-linear function of an index of financial deregulation, however, results in an estimated model which cannot be rejected against the ECM specification for both countries examined, yielding sensible and well-determined estimates of the parameters, especially for the UK, where the intertemporal elasticity of substitution is statistically significant and of plausible magnitude. Moreover, the implied time-varying behaviour of λ_t accords well with the observed pace and pattern of financial deregulation in the two countries considered, approaching zero in the second half of the 1980s. Further, the non-monotonic behaviour of λ_t may explain why previous researchers have detected evidence of a linear upward trend in its behaviour for the UK and have not detected any significant change in its path for France.

Appendix 1.1 Modelling the consumption function

Using the same notation as in Section 1.2.1 and under the assumption that the representative consumer's utility function is strictly concave and twice differentiable, the consumer's optimization problem involves the maximisation of the objective function:

$$V_0 = E_0 \sum_{t=0}^{T} (1 + \rho)^{-t} U(C_t) \tag{A1.1.1}$$

subject to the budget constraint:

$$A_{t+1} = (1 + r_t)(A_t + Y_t^{\alpha} - C_t), \qquad t=0,\dots,T \tag{A1.1.2}$$

Solving for C_1 from the constraint (A1.1.2) as

$$C_1 = (1+r_0)(A_0+Y_0^{\alpha}-C_0) + Y_1^{\alpha} - A_2(1+r_1)^{-1} \tag{A1.1.3}$$

and using (A1.1.3) onto (A1.1.1) yields:

$$\begin{aligned} V_0 &= U(C_0) + \\ &E_0(1+\rho)^{-1}U\left((1+r_0)(A_0+Y_0^{\alpha}-C_0)+Y_1^{\alpha}-\frac{A_2}{(1+r_1)}\right)+ \\ &E_0(1+\rho)^{-2}U(C_2) + \dots \end{aligned} \tag{A1.1.4}$$

Maximising (A1.1.4) with respect to C_0 - given A_0 - gives:

$$0 = \frac{\partial V_0}{\partial C_0} = U'(C_0) + E_0 (1 + \rho)^{-1} U'(C_1) [-(1 + r_0)] \quad \text{(A1.1.5)}$$

which implies:

$$U'(C_0) = E_0\left[(1 + \rho)^{-1} (1 + r_0) U'(C_1)\right] \quad \text{(A1.1.6)}$$

Analogously, maximising (A1.1.4) with respect to C_1 - given A_1 - yields:

$$U'(C_1) = E_1\left[(1 + \rho)^{-1} (1 + r_1) U'(C_2)\right] \quad \text{(A1.1.7)}$$

and, therefore, in general:

$$U'(C_t) = E_t\left[(1 + \rho)^{-1} (1 + r_t) U'(C_{t+1})\right] \quad \text{(A1.1.8)}$$

Alternatively, the set of stochastic Euler equations (A1.1.8) may be derived using the Lagrangean method. At time t=0, for example, the maximand may be written as:

$$\begin{aligned} L = V_0 + E_0\{\mu_0 [A_1 - (1 + r_0) (A_0 + Y_0^{\alpha} - C_0)] + \\ \mu_1 [A_2 - (1 + r_1) (A_1 + Y_1^{\alpha} - C_1)] + \ldots\} \end{aligned} \quad \text{(A1.1.9)}$$

and the first-order conditions with respect to C_0, C_1, ..., A_1, ... are:

$$
\begin{aligned}
0 &= \frac{\partial L}{\partial C_0} = U'(C_0) + E_0 \mu_0 (1 + r_0) \\
0 &= \frac{\partial L}{\partial C_1} = E_0 (1+\rho)^{-1} U'(C_1) + E_0 \mu_1 (1 + r_1) \\
\vdots & \quad \vdots \qquad\qquad \vdots \\
0 &= \frac{\partial L}{\partial A_1} = E_0 \mu_0 - E_0 \mu_1 (1 + r_1) \\
\vdots & \quad \vdots \qquad\qquad \vdots
\end{aligned} \tag{A1.1.10}
$$

Hence, using the first-order conditions (A1.1.10), a relationship may be derived between the marginal utility of consumption and the gross rate of interest in one period and the next, which - at time 0 - takes the form:

$$
\frac{U'(C_0)}{E_0 (1+\rho)^{-1} U'(C_1)} = \frac{E_0 \mu_0 (1+r_0)}{E_0 \mu_1 (1+r_1)} \tag{A1.1.11}
$$

which, by substitution, yields exactly (A1.1.6) and, in general, the set of Euler equations (A1.1.8).

Assuming an isoelastic utility function, ie. $U(C_t) = \sigma(\sigma-1)^{-1} C_t^{(\sigma-1)/\sigma}$, equation (A1.1.8) becomes:

$$
C_t^{-\sigma^{-1}} = E_t [(1+\rho)^{-1} (1+r_t) C_{t+1}^{-\sigma^{-1}}] \tag{A1.1.12}
$$

or

$$
C_t^{-\gamma} = E_t \left(\frac{1+r_t}{1+\rho} C_{t+1}^{-\gamma} \right) \tag{A1.1.13}
$$

where γ [$\equiv \sigma^{-1}$] denotes the coefficient of relative risk aversion. Taking logs in (A1.1.13) and using the approximation $\log(1+x) \approx x$ for small x, equation (A1.1.13) may be rewritten as:

$$E_t e^{-\gamma(\Delta c_{t+1} - \gamma^{-1} r_t) - \rho} = 1 \qquad \text{(A1.1.14)}$$

Then, under the assumption that $\log(C_{t+1})$ and r_t are joint normally distributed, computing the expected value in equation (A1.1.14) yields[31]:

$$e^{-\gamma E_t(\Delta c_{t+1} - \gamma^{-1} r_t) - \rho + \frac{1}{2}\gamma^2 Var_t(\Delta c_{t+1} - \gamma^{-1} r_t)} = 1 \qquad \text{(A1.1.15)}$$

Also, rearranging equation (A1.1.15) yields:

$$E_t \Delta c_{t+1} = -\gamma^{-1}(E_t r_t - \rho) + \frac{1}{2}\gamma\, Var_t(\Delta c_{t+1} - \gamma^{-1} r_t) \qquad \text{(A1.1.16)}$$

which is identical to equation (1.1.4) in Section 1.2.1. Also, if the variance term in (A1.1.16) is assumed to be constant and absorbed in the intercept term, the Hall random walk model of consumption - equation (1.1.6) - follows directly.

Appendix 1.2 Instrumental variables-moving average estimation

Patterson and Pesaran (1992) argue that, if a statistically significant first-order disturbance is not present in a consumption function, then once lagged variables are potentially valid instruments, regardless of the econometric problems discussed in Section 1.5. In general, for example, the presence of serial correlation in the residuals may simply be due to some form of misspecification of the model. Also, a poor choice of the instrument set may induce bias in the estimated coefficients relative to their standard error, when the instrument set is not substantially correlated to the explanatory variables of the estimated regression (Nelson and Startz, 1990a,b; Patterson and Pesaran, 1992; Hall, Rudebusch and

[31] The result needed is that if $Y = \log(X) \sim N(\mu, \sigma^2)$, then:

$$E(X) = e^{\mu + \frac{1}{2}\sigma^2}$$

Wilcox, 1994). Following the approach proposed by Pesaran (1990) and Patterson and Pesaran (1992), in the estimations executed in Section 1.5 we test whether a significant MA(1) error is present in order to decide whether a once-lagged or a twice-lagged instrument set is appropriate. Thus, we employ instrumental variable estimation which allows for MA error processes (IVMA), also including the first lag of the instruments, and test whether a MA(1) error is significant. If no significant MA(1) error is discovered in a model estimated using IVMA estimation, we reestimate the model with the once-lagged instrument set using conventional IV estimation.

Consider a linear regression model of the form:

$$y_t = x_t'\beta + u_t, \qquad t=1,2,\ldots,n \qquad \text{(A1.2.1)}$$

where y_t is a nx1 vector of dependent variables, x_t' is a nxk vector of explanatory variables, β is a kx1 vector of parameters and the nx1 vector u_t is described by:

$$u_t = \gamma\epsilon_{t-1} + \epsilon_t, \qquad \epsilon_t \sim N(0,\sigma_\epsilon^2) \qquad \text{(A1.2.2)}$$

Given a nxs matrix Z containing observations on the s instrumental variables z_t, the IVMA estimators of β and γ are computed by minimising the criterion function:

$$S(\beta,\gamma) = \left(\frac{T}{2}\right)u^{*\prime}P_Z u^* + \left(\frac{1}{2}\right)\log|\Omega| \qquad \text{(A1.2.3)}$$

where T is the number of observations, $u^{*\prime}$ is a vector of forward filtered values of $u=y-x'\beta$, $E(uu')=\sigma_\epsilon^2\Omega$, P_Z is the projection matrix $Z(Z'Z)^{-1}Z'$, and Ω is the variance-covariance matrix. The forward filter procedure used is similar to a Kalman filter procedure and it is only applied to the dependent variable and the regressors, not to the instruments[32] (Pesaran, 1988, 1990). Also, the second term in (A2.3) is an additional Jacobian term which is asymptotically negligible when the MA process is invertible, but its inclusion in the criterion function is important to ensure that in small samples the criterion function achieves a minimum, when the roots of the MA process are close to the unit circle (Pesaran, 1990).

For given values of γ, β may be estimated by the IV regression of the forward filtered variables y^* on X^*, using conditional maximum likelihood (ML) estimation. In large samples, however, the estimates given by the conditional ML method should be identical to the estimates obtained by the exact ML method. Also, γ is estimated iteratively using a modified Powell's method of conjugate directions, which has the advantage that, unlike other algorithms (eg. Gauss-Newton), it does not require the computation of complex derivatives of the IV minimand (A2.3) (Powell, 1964; Pesaran and Pesaran, 1991).

Also, using the adjusted residuals defined as:

$$\tilde{\epsilon}_t = -\gamma\tilde{\epsilon}_{t-1} + \tilde{u}_t, \qquad t=1,2,\ldots,n \qquad \text{(A1.2.4)}$$

where

$$\tilde{u}_t = y_t - \sum_{i=1}^{k} \beta_i x_{it}, \qquad t=1,2,\ldots,n \qquad \text{(A1.2.5)}$$

Pesaran (1990) also shows how to derive various summary statistics (R^2, Durbin-Watson, F-statistic, etc.).

[32] Pesaran (1990) provides a comprehensive treatment of the IVMA estimation procedure for the general case of a MA(q) error.

Appendix 1.3 The nonlinear instrumental variables estimator

The econometric literature provides nonlinear instrumental variable (NLIV) estimators for three different models: (a) nonlinear only in the parameters, (b) nonlinear only in the variables and (c) nonlinear in both variables and parameters. Case (a) is analysed by Zellner, Huang and Chau (1965); for case (b) see Kelejian (1971) and other references cited in Goldfeld and Quandt (1972), while the analysis of the third case is largely due to Amemiya (1974)[33].

In this appendix, following Amemiya (1974), we describe the NLIV estimator for a model which is nonlinear in both variables and parameters. The NLIV estimator defined by Amemiya (1974) reduces to the NLIV estimator of Kelejian (1971) if nonlinearity is present only in the variables, to the NLIV estimator of Zellner, Huang and Chau (1965) if nonlinearity is present only in the parameters, and to the simple IV estimator if no nonlinearity is present in the regression function. Amemiya (1974) also shows that the optimality properties of the IV estimator also extend to the special case of a regression function which is nonlinear only in the parameters, which represents exactly the case of the consumption function proposed in this study.

Consider the nonlinear regression:

$$y_t = f(x_t, \beta) + u_t, \qquad u_t \sim (0, \sigma^2) \tag{A1.3.1}$$

where y_t is a scalar random variable, x_t is a H-component vector which includes both endogenous variables - correlated with u_t - and exogenous variables - known constants; β is a G-component vector of parameters and f is a function nonlinear in both x and β, and twice differentiable with respect to β. Equation (A1.3.1) is just one of the equations forming the simultaneous equations model from which the endogenous variables in (A1.3.1) are generated. Given T observations on y_t and x_t, estimates of β and σ^2 may be derived using the NLIV estimator.

[33] For a very comprehensive treatment of nonlinear instrumental variables, and, more generally, of nonlinear statistical models, see Amemiya (1985) and Gallant (1987).

The NLIV estimator of β, β_{NLIV}, is obtained by minimizing the criterion function:

$$\Phi_{IV}(\beta) = (y - f)' P_Z (y - f) \qquad \text{(A1.3.2)}$$

where y and f are T-component vectors whose t-th elements are y_t and $f(x_t,\beta)$ respectively; the projection matrix $P_Z=Z(Z'Z)^{-1}Z'$, where Z is a TxK matrix of known constants. Under the assumptions that (a) the paramater space is compact (ie. bounded and closed), (b) u_t is independently identically distributed, (c) the lim $(1/T)Z'Z$ exists and is nonsingular, (d) the plim of $(1/T)(\partial f'/\partial\beta)Z$ converges to a constant matrix of rank G uniformly in β, (e) the plim of $(1/T)(\partial^2 f'/\partial\beta_i\partial\beta)Z$ converges to a constant matrix uniformly in β for i = 1, 2,..., G, where β_i is the i-th element of the vector of parameters β, Amemiya (1974) proves *Theorem 1*, which states that:

i) the estimated β, β^{NLIV}, converges in probability to the true value β_0, and

ii) $T^{1/2}(\beta^{NLIV} - \beta_0)$ converges in distribution to

$$N\left\{0,\ \sigma^2\left[plim\left(\frac{1}{T}\right)\left(\frac{\partial f'}{\partial\beta}\right)\bigg|_{\beta_0}\cdot P_Z\left(\frac{\partial f}{\partial\beta'}\right)\bigg|_{\beta_0}\right]^{-1}\right\} \qquad \text{(A1.3.3)}$$

The iterative procedure which allows to execute the minimization of the criterion function (A1.3.2) is the Gauss-Newton method, based on the algorithm:

$$\beta_n^{NLIV} = \beta_{n-1}^{NLIV} + \left[\left(\frac{\partial f'}{\partial\beta}\right)P_Z\left(\frac{\partial f}{\partial\beta'}\right)\right]^{-1}\left(\frac{\partial f'}{\partial\beta}\right)P_Z\,(y - f) \qquad \text{(A1.3.4)}$$

where f and $(\partial f'/\partial\beta)$ on the right hand side of the equation are estimated during the (n-1)-th iteration[34].

The analysis of the efficiency properties of the NLIV estimator presents great difficulties, particularly in deriving its asymptotic distribution. For the linear two-stage least squares (2SLS) estimator, it is a well-known result that its asymptotic distribution is the same as the limited information maximum likelihood (LIML) estimator[35], having the smallest asymptotic variance-covariance matrix in the class of instrumental variables estimators. For both a model nonlinear in variables and parameters and a model nonlinear only in the variables, the econometric literature has neither defined the variance-covariance matrix of the maximum likelihood estimator nor suggested a precise set of criteria for defining the optimal choice of Z. Amemiya (1974) shows, however, how the efficiency properties of 2SLS may be extended to nonlinear 2SLS (NL2SLS) when the regression equation is assumed to be nonlinear only in the parameters. Obviously, those properties can be extended to the NLIV estimator in a straithforward way, since the 2SLS estimator is just a special case of the IV estimator.

Consider the model:

$$y = X\,\alpha(\beta) + u \qquad \text{(A1.3.5)}$$

where X represents the matrix whose t-th row is x_t and α is an H-component vector (H > G), whose elements are functions of β. It follows from (ii) of *Theorem 1* that the asymptotic variance-covariance matrix of the NLIV estimator, given (A1.3.5), is:

[34] For a more comprehensive treatment of the convergence properties of the Gauss-Newton algorithm and its more common applications, see Davidson and MacKinnon (1993, Chapter 6).

[35] The LIML method, also called least variance ratio (LVR) method, is the first method developed for estimating simultaneous-equations models. It was suggested by Anderson and Rubin (1949) and was widely used until the development of the two-stage least squares estimator by Theil (1953) and Basmann (1957).

$$\sigma^2\left[\left(\frac{\partial\alpha'}{\partial\beta}\right)\bigg|_{\beta_0} plim\left(\frac{X'P_Z X}{T}\,\frac{\partial\alpha}{\partial\beta'}\right)\bigg|_{\beta_0}\right]^{-1} \tag{A1.3.6}$$

The variance-covariance matrix (A1.3.6) is minimised when Z consists of all the exogenous variables of the system, say Z_0. The log-likelihood function concentrated[36] in β is:

$$\log L = \gamma + \left(\frac{T}{2}\right)\log\left[\frac{(y - X\alpha)'M(y - X\alpha)}{(y - X\alpha)'(y - X\alpha)}\right] \tag{A1.3.7}$$

where γ is a constant term and $M = [I-Z_0(Z_0'Z_0)^{-1}Z_0']$. Amemiya (1974) shows that the value of β, say β^*, which maximises the log-likelihood function (A1.3.7) is exactly the LIML estimator of β. Amemyia (1974) also shows that, following the same steps and under the same assumptions used in order to derive *Theorem 1*, $T^{(1/2)}(\beta^*-\beta_0)$ is asymptotically normally distributed with mean zero and variance equal to (A1.3.7), or, in other words, NLIV has the same asymptotic properties as the LIML.

[36] The meaning of "concentrating the likelihood function" needs to be clarified. Consider a parameter vector A, containing two subvectors A_1 and A_2. It is often possible to concentrate the likelihood function, that is to deal with subsets of the parameter vector A, which turns out to be very convenient, both analytically and numerically. For example, if we know values for A_1, then we can express A_2 in terms of A_1. If, for example, the formula relating the two subvectors is of the form $A_2 = g(A_1)$, then we can write the likelihood function as $L(A_1, g(A_1))$, which can be restated as the concentrated likelihood function $L^*(A_1)$ (Cuthbertson, Hall and Taylor, 1992).

Table 1.1 The liquidity-constrained model with a constant real rate of interest

Panel A: Results for the UK

Method	λ (s.e.)	1st-stage regress: Δy_t [p.v.]	Sargan test statistic [p.v.]	MA(1) test statistic [p.v.]
OLS	0.249 (0.052)	---	---	---
IV	0.210 (0.080)	0.237 [0.000]	31.20 [0.013]	---
IVMA	0.250 (0.055)	0.237 [0.000]	30.75 [0.014]	-1.64[0.101]

Panel B: Results for France

Method	λ (s.e.)	1st-stage regres.: Δy_t [p.v.]	Sargan test statistic [p.v.]	MA(1) test statistic
OLS	0.194 (0.073)	---	---	---
IVMA	0.360 (0.123)	0.287 [0.000]	27.94 [0.063]	-3.00 [0.00]
IVMA*	0.117 (0.108)	0.320 [0.000]	59.76 [0.000]	-3.09 [0.00]

The estimated equation is:

$$\Delta c_t = \mu + \lambda \Delta y_t + \epsilon_t$$

Notes: The instrument set for the UK regression is $\{(c\text{-}y)_{t-1}, \Delta c_{t-1}, \Delta c_{t-2}, \Delta c_{t-3}, \Delta y_{t-1}, \Delta y_{t-2}, \Delta y_{t-3}, \Delta\pi_{t-1}, \Delta\pi_{t-2}, \Delta\pi_{t-3}, r_{t-2}, r_{t-3}, r_{t-4}, tb_{t-2}, tb_{t-3}, tb_{t-4}, i_{t-2}, i_{t-3}, i_{t-4}, \text{constant}\}$, where π denotes the consumption deflator, r is the real Deposit Rate, tb the real Treasury Bill Rate and i the real Bank Rate. The instrument set for the regression on French data is $\{(c\text{-}y)_{t-2}, \Delta c_{t-2}, \Delta c_{t-3}, \Delta c_{t-4}, \Delta y_{t-2}, \Delta y_{t-3}, \Delta y_{t-4}, \Delta\pi_{t-2}, \Delta\pi_{t-3}, \Delta\pi_{t-4}, r_{t-3}, r_{t-4}, r_{t-5}, ib_{t-3}, ib_{t-4}, ib_{t-5}, i_{t-3}, i_{t-4}, i_{t-5}, \text{constant}\}$, where π denotes the consumption deflator, r is the real money market rate, ib the real interbank rate and i the real Pibor rate. The column labelled "first-stage regression" report the adjusted R^2 for the OLS regression of the variable considered onto the instruments; in square brackets is the p-value for the test of the null hypothesis that all the coefficients except the constant are zero. Estimated standard errors (s.e.) are reported in parentheses; p-values (p.v.) are reported in square brackets. The Sargan test statistic for the validity of the instruments is asymptotically distributed as $\chi^2(18)$ under the null hypothesis of valid instruments. All tests and estimated standard errors are heteroscedasticity-robust. In the last column, the asymptotic *t*-ratio for the significance of the MA(1) coefficient in the disturbance is reported.

* The instrument set is $\{(c\text{-}y)_{t-1}, \Delta c_{t-1}, \Delta c_{t-2}, \Delta c_{t-3}, \Delta y_{t-1}, \Delta y_{t-2}, \Delta y_{t-3}, \Delta\pi_{t-1}, \Delta\pi_{t-2}, \Delta\pi_{t-3}, r_{t-2}, r_{t-3}, r_{t-4}, ib_{t-2}, ib_{t-3}, ib_{t-4}, i_{t-2}, i_{t-3}, i_{t-4}, \text{constant}\}$.

Table 1.2 The model with only unconstrained consumers

Panel A: Results for the UK

Method	σ (s.e.)	1st-stage regression r_{t-1}	Sargan test statistic [p.v.]	MA(1) test statistic [p.v.]
OLS	0.245 (0.068)	---	---	---
IV	0.238 (0.090)	0.540 (0.000)	28.66 [0.026]	---
IVMA	0.188 (0.053)	0.540 (0.000)	28.29 [0.029]	0.702[0.483]

Panel B: Results for France

Method	σ (s.e.)	1st-stage regression r_{t-1}	Sargan test statistic [p.v.]	MA(1) test statistic [p.v.]
OLS	0.009 (0.069)	---	---	---
IVMA	0.082 (0.049)	0.775 (0.000)	35.51 [0.008]	-2.11[0.035]
IVMA*	0.075 (0.041)	0.787 (0.000)	55.47 [0.000]	-2.61[0.009]

The estimated equation is

$$\Delta c_t = \mu + \sigma r_{t-1} + \epsilon_t$$

See Notes to Table 1.1.

Table 1.3 The model with both constrained and unconstrained consumers

Panel A: Results for the UK

Method	λ (s.e.)	θ (s.e.)	Implied σ (s.e.)	Sargan test statistic [p.v.]	MA(1) test statistic [p.v.]
OLS	0.207 (0.052)	0.188 (0.721)	0.237 (0.087)	---	---
IV	0.148 (0.083)	0.198 (0.088)	0.233 (0.104)	29.52 [0.021]	---
IVMA	0.184 (0.056)	0.119 (0.049)	0.146 (0.062)	29.54 [0.020]	-1.236 [0.216]

Panel B: Results for France

Method	λ (s.e.)	θ (s.e.)	Implied σ (s.e.)	Sargan test statistic [p.v.]	MA(1) test statistic [p.v.]
OLS	0.194 (0.074)	0.012 (0.065)	0.015 (0.081)	---	---
IVMA	0.344 (0.126)	0.050 (0.031)	0.076 (0.121)	28.329 [0.005]	-3.042 [0.002]
IVMA*	0.088 (0.118)	0.061 (0.039)	0.067 (0.087)	57.456 [0.000]	-3.082 [0.00]

Notes: The estimated equation is

$$\Delta c_t = \mu + \lambda \Delta y_t + \theta r_{t-1} + \epsilon_t$$

Notes: The instruments are as reported in Table 1.1. The implied σ is calculated from $\sigma = \theta/(1-\lambda)$. The Sargan test statistic is asymptotically distributed as $\chi^2(17)$ under the null hypothesis of valid instruments. See also Notes to Table 1.1.

Table 1.4 The liquidity-constrained model with time-varying coefficients

Panel A: Results for the UK

ϕ_0 (s.e.)	ϕ_1 (s.e.)	ϕ_2 (s.e.)	σ (s.e.)	b (s.e.)	Test for joint sign. of ϕ_1 and ϕ_2 [p.v.]
0.000 (-)	2.120 (0.967)	-6.519 (5.077)	0.153 (0.076)	3.128 (0.327)	11.07 [0.004]

Hansen test statistic [p.v.]	23.86 [0.067]

Order	Autocorrelation coefficient (s.e.)	ARCH test statistic [p.v.]
Order 1	-0.158 (0.090)	3.200 [0.074]
Order 4	0.087 (0.097)	5.640 [0.228]
Order 8	0.094 (0.098)	5.930 [0.655]

Panel B: Results for France

ϕ_0 (s.e.)	ϕ_1 (s.e.)	ϕ_2 (s.e.)	σ (s.e.)	b (s.e.)
0.000 (-)	6.447 (1.697)	-10.997 (3.725)	0.000 (-)	1.550 (0.219)

Hansen test statistic [p.v.]	18.920 [0.28]

Order	Autocorrelation coefficient (s.e.)	ARCH test statistic [p.v.]
Order 1	-0.336 (0.207)	0.244 [0.124]
Order 4	-0.233 (0.227)	0.009 [0.110]
Order 8	-0.015 (0.133)	0.058 [0.079]

(Table 1.4 continued ...)

(... continued)

The estimated model is

$$\Delta c_t = \mu + \lambda_t \Delta y_t + (1 - \lambda_t)\sigma r_{t-1} + \epsilon_t$$

$$\lambda_t = 1 \quad for \ (\phi_0 + \phi_1 k_t + \phi_2 k_t^b) \leq 0$$

$$\lambda_t = [1 - (\phi_0 + \phi_1 k_t + \phi_2 k_t^b)] \quad for \ 0 \leq (\phi_0 + \phi_1 k_t + \phi_2 k_t^b) \leq 1$$

$$\lambda_t = 0 \quad for \ (\phi_0 + \phi_1 k_t + \phi_2 {k_t}^b) \geq 1$$

Notes: The instruments are as reported in Table 1.1, in addition to a linear time trend and k_{t-1}. Estimation is by non-linear instrumental variables (Amemyia, 1974; Gallant, 1987). The Hansen test statistic for the validity of the instruments is asymptotically distributed as $\chi^2(16)$ under the null hypothesis of valid instruments. The ARCH test statistics test for autoregressive conditional heteroscedasticity (ARCH) in the residuals of the specified order. See also Notes to Table 1.1.

Figure 1.1 Estimated time trajectory of λ_t in the UK

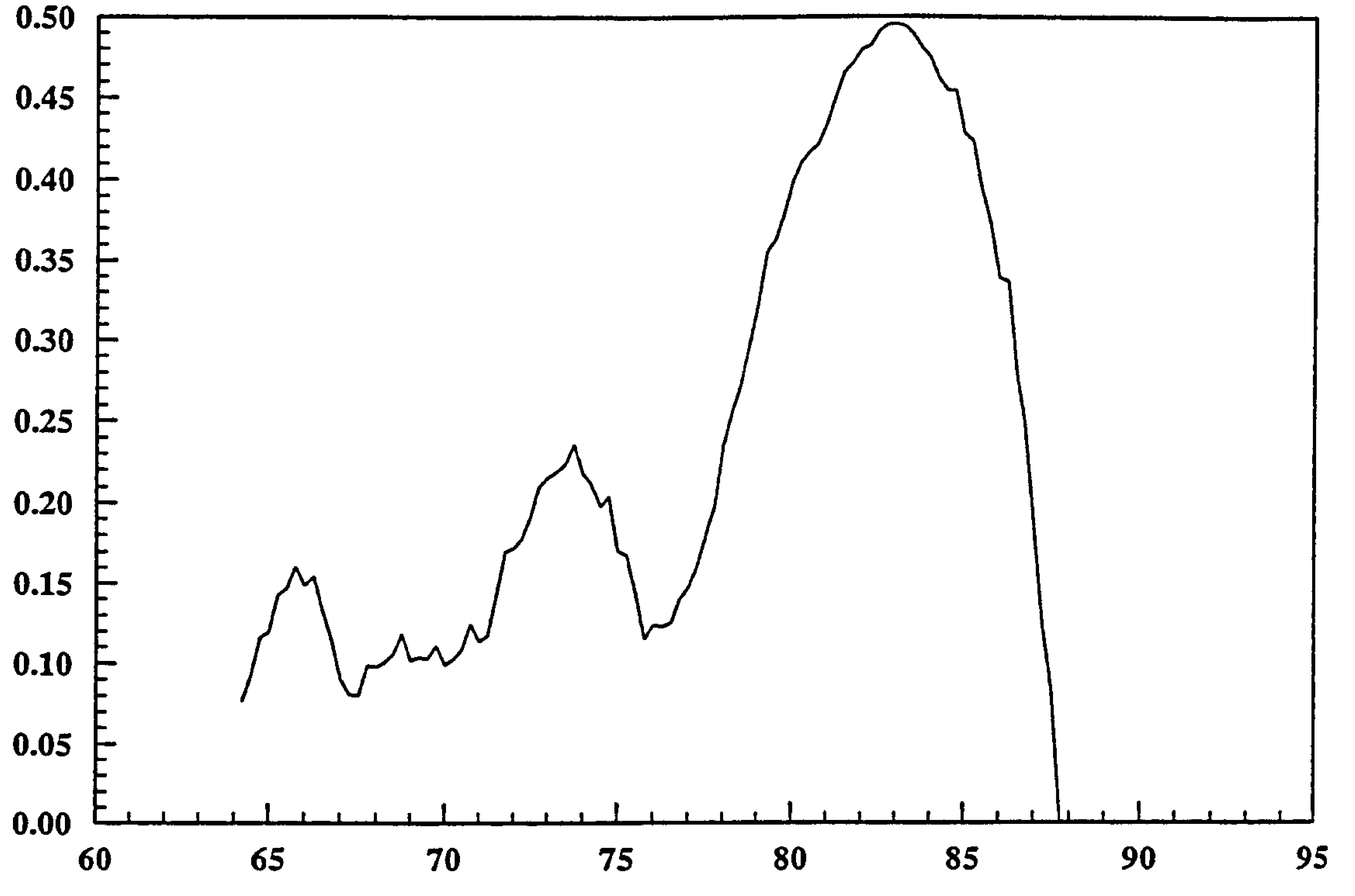
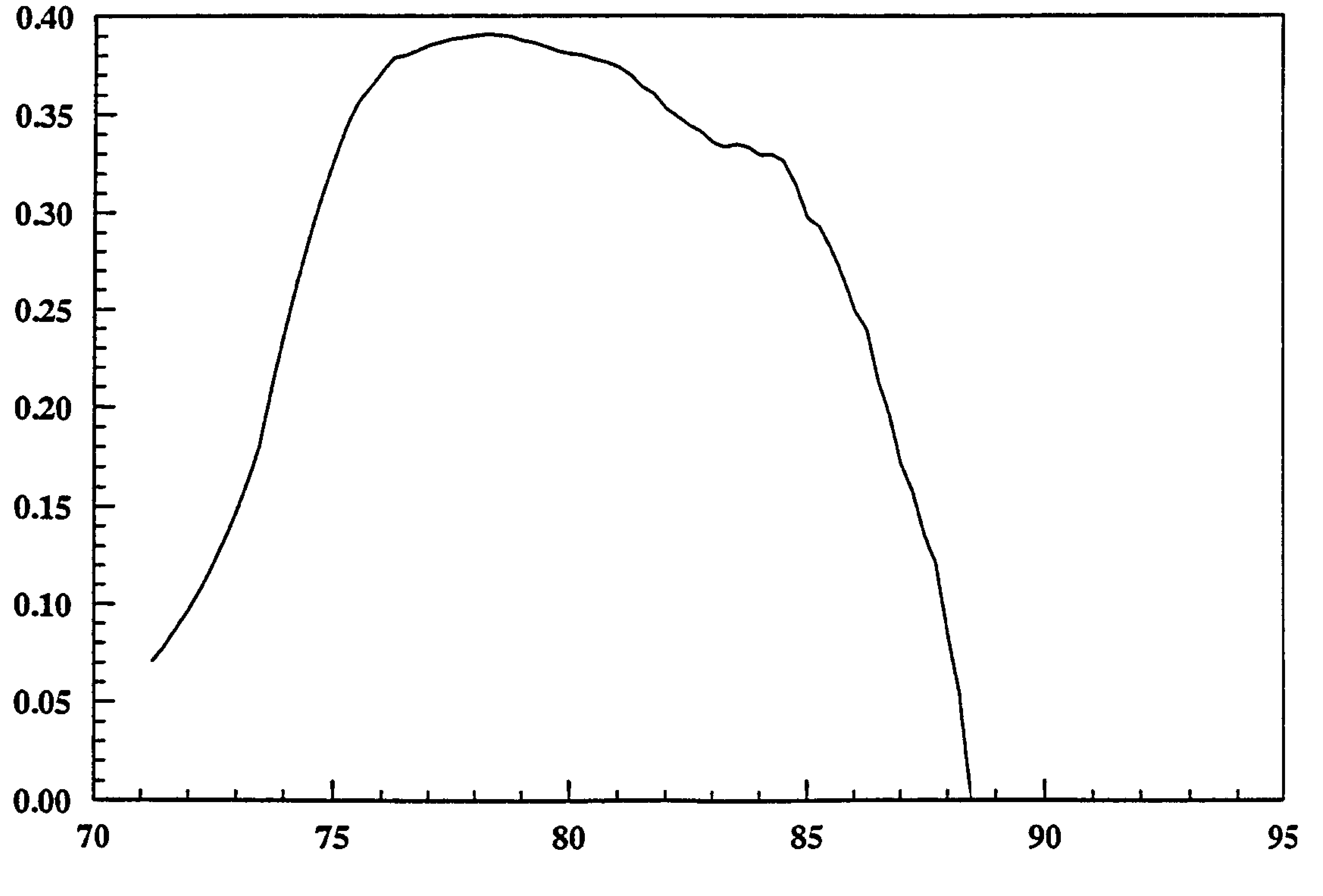


Figure 1.2 Estimated time trajectory of λ_t in France



CHAPTER 2

Capital Markets Integration, European Capital Flows and Saving-Investment Correlations: Transitory versus Permanent

'Feldstein's discovery of the tight link between national saving and investment rates continues to baffle the profession. Ample research over the past few years has failed to reject the basic finding... The Feldstein finding runs counter to the spirit of the open economy literature in which, under conditions of perfect capital mobility, changes in national saving rates are primarily reflected in the current account, *not* in investment.'

(Dornbusch, 1991, p. 220)

2.1 Introduction

In this chapter we look at a range of issues concerning capital flows in Europe. We gauge the extent of international capital mobility through an examination of saving-investment correlations and, in particular, investigate the difference between the short-run and the long-run saving-investment correlation coefficient, in order to shed light both on the validity of the Feldstein-Horioka regression as a means of measuring the degree of capital mobility and on its implications (Feldstein and Horioka, 1980). In addition, we attempt to measure the degree of capital mobility in the UK and its evolution over time. In October 1979, exchange control was abolished in the UK, ending a long period of restrictions on capital flows between the UK and the international economy. Therefore, one would expect a significant reduction in the correlation of savings and investment after this point, if this is a valid measure of international capital mobility.

The abolition of remaining capital controls by France and Italy in 1991 and the creation of a European Financial Area also raise important questions concerning the nature and the extent of capital mobility in Europe. The single most important macroeconomic topic currently being debated in Europe is the analysis of the relative benefits and costs of the adoption of a single European currency. The claimed benefits mainly derive from the reduced exchange rate risk and lower transaction costs between countries, which may well lead to more efficient production. The main loss is the ability to run an independent monetary policy. This is of little concern in the face of disturbances which affect all the countries in the region in a similar manner, when a common policy response will suffice, or when there are fully flexible factor markets. In the case of country-specific disturbances and sluggish factor markets adjustments, however, the loss of an independent monetary policy hampers adjustment of equilibrium. Capital flows provide a means of insuring individuals' incomes against changes in local productivity, since private investors can create a portfolio which is diversified in the returns to both human and financial capital.[1]

We provide further evidence on the extent and effect of European capital mobility through an examination of GNP/GDP ratios across the largest European countries and a comparison of these with similar calculations executed for regions within the UK. In addition, we also examine the role of private financial markets in reducing regional risk by developing a simple empirical model which is then estimated on data for the European Union.

The remainder of the chapter is set out as follows. Section 2.2 provides an overview of the measures of the degree of international capital mobility commonly used by the relevant literature. In addition, Section 2.2 provides a brief overview of the enormous literature on the Feldstein-Horioka regression and also discusses the puzzling difference between the short-run and the long-run saving-investment coefficients. In Section 2.3 we describe the estimation techniques that we adopt in

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[1] For a comprehensive discussion of the issues relating to Economic Monetary Union, see Masson and Taylor (1993) or De Grauwe (1994).

order to decompose saving and investment shares of GDP into temporary and permanent components. In Section 2.4 we describe the data used in order to construct investment and saving shares of GDP for the UK. In Section 2.5 we investigate the relationship between saving and investment shares of GDP in a vector autoregression context; we execute the decomposition of saving and investment shares, and compute correlation coefficients between the temporary and the permanent components. In Section 2.6, further evidence on the extent and effect of European capital mobility is adduced through an examination of European GNP/GDP ratios and a comparison of those with similar calculations executed for different regions within the UK. In Section 2.7, we examine the role of private financial markets in reducing regional risk by developing an empirical model which is then estimated on European Union data. A final section concludes.

2.2 Quantifying international capital mobility

2.2.1 The link between interest parity conditions and the Feldstein-Horioka criterion

Since the late 1970s, it has become common among economists to describe the world financial system as characterised by perfect capital mobility. This mainly follows from the enhanced financial integration induced by the drastic reduction or complete removal of capital controls in the mid 1970s by a relatively small number of countries. Principal among these were the USA, Germany, Canada and Switzerland (Shafer, 1995)[2]. Also contributing to this impression is the ongoing process of widespread financial deregulation, the improvement of information and

[2] Shafer (1995) contains a chronological account of controls of capital movements for OECD countries during the post-war period.

communication technology in financial markets and the recycling of OPEC surpluses to developing countries (Frankel, 1993; Coakley, Kulasi and Smith, 1995a). Therefore, although many of both the less developed and the major industrialised countries retained restrictions on international capital flows for much of the floating rate period, the assumption of perfect capital mobility was routinely used in the literature on the main exchange rate models, with the exception of variants of the Mundell-Fleming model and the portfolio balance model (Taylor, 1995).

Following Frankel (1993), at least two different approaches to quantifying international capital mobility are suggested by the enormous and growing relevant literature: studies using the Feldstein-Horioka definition based on saving-investment correlations, and studies based on interest rate parity conditions.

The Feldstein-Horioka criterion is, however, closely related to measures of international capital mobility based on interest parity conditions. Below we demonstrate how three different interest parity conditions commonly used by researchers correspond to three different criteria of financial integration, under different underlying assumptions[3]. In particular, use of an interest-parity-condition-based measure of international capital mobility always implies a definition of capital mobility according to which the null hypothesis of *perfect* capital mobility is a joint hypothesis that domestic and foreign bonds - identical in all respects but the currency of denomination - are perfect substitutes and that arbitrage ensures the interest parity condition holds continuously. Clearly, this notion of international

[3] See Frankel (1993) or Lemmen and Eijffinger (1995a) for an excellent discussion of these issues.

capital mobility relies on comovements of domestic and foreign *prices* - ie. interest rates - whereas the notion of international capital mobility suggested by Feldstein and Horioka relies on the comovement of domestic *quantities*, namely savings and investment.

The three criteria based on interest parity conditions are the covered nominal interest parity (CIP), the *ex ante* uncovered nominal interest parity (UIP) and the *ex ante* real interest parity (RIP)[4]. The CIP condition states that:

$$i_t - i_t^* = f_t^{t+k} - s_t \tag{2.1}$$

where i_t, f_t^{t+k} and s_t denote the domestic nominal interest rate on a k-period bond held between time t and t+k, the forward exchange rate agreed at time t for the delivery of foreign currency at time t+k and the nominal exchange rate at time t (domestic price of foreign currency); asterisks denote foreign variables[5]. If (2.1) holds, then the forward premium - or forward discount - equals the nominal interest rate differential at the appropriate maturity. Under perfect capital mobility, the covered nominal interest rate differential or country premium $[=(i_t-i_t^*)-(f_t^{t+k}-s_t)]$, which may be seen as capturing the impact of actual and future capital controls, transaction costs and default risks, equals zero (Frankel and MacArthur, 1988). Assuming that CIP holds and that the forward rate is an unbiased predictor of the

[4] See Taylor and Sarno (1997, Chapter 2) for an extensive discussion of the theoretical and empirical literature on interest parity conditions.

[5] Here and throughout this section, lower-case letters represent natural logarithms, except for the interest rate.

expected future spot exchange rate, ie. $E_t(s_{t+k})=f_t^{t+k}$ (where $E_t=E(\cdot|t)$ denotes the mathematical expectation operator conditional on information available at time t), yields the UIP condition:

$$i_t - i_t^* = E_t(s_{t+k} - s_t) \tag{2.2}$$

Using the decomposition method of Frankel and MacArthur (1988), equation (2.2) implies that under perfect capital mobility both the country premium and the exchange rate premium $[=(f_t^{t+k} - s_t) - E_t(s_{t+k} - s_t)]$ must be zero, since the exchange rate premium measures the extent to which the forward rate is a biased predictor of the future spot rate[6].

Then, assuming that UIP holds and also assuming a zero expected real exchange rate change, ie. $E_t(q_{t+k})=q_t=s_t-p_t+p_t^*$ (where q, p and p^* denote the real exchange rate and the domestic and foreign price levels respectively), yields the RIP condition:

$$E_t(r_{t+k}) = E_t(r_{t+k}^*) \tag{2.3}$$

---

[6] In other words, equation (2.2) is reparametrised as:

$$i_t - i_t^* - [E_t(s_{t+k} - s_t)] = [i_t - i_t^* - (f_t^{t+k} - s_t)] + [f_t^{t+k} - s_t - E_t(s_{t+k} - s_t)]$$

where the two terms on the right hand side represent the country premium and the exchange rate premium respectively.

where r_{t+k} denotes the domestic real interest rate at time t on a bond held between time t and time t+k. The RIP condition clearly refers to a relatively stronger notion of integration since it requires both financial and nonfinancial perfect capital mobility, where nonfinancial includes the mobility of goods and services as well as labor and physical capital. For the RIP to hold, the country premium, the exchange rate premium and the expected real exchange rate premium must all be zero. This may be better seen by using the following decomposition of the RIP:

$$
\begin{aligned}
E_t(r_{t+k}-r^*_{t+k}) = {} & [i_t - i^*_{t,t+k} - (f_t^{t+k} - s_t)] + \\
& [(f_t^{t+k} - s_t) - E_t(s_{t+k} - s_t)] + \\
& [E_t(s_{t+k} - s_t) - E_t(p_{t+k} - p_t) + E_t(p^*_{t+k} - p^*_t)]
\end{aligned} \tag{2.4}
$$

where the third term on the right-hand side of the equation is a measure of the real depreciation of the domestic currency or, in other words, a measure of the extent to which *ex ante* purchasing power parity is violated[7]. Also, the last two terms on the right hand side of (2.4) together represent the currency premium. It is clear how the third interest-rate-parity-based criterion implies a much stronger definition of perfect capital mobility relative to the first two.

The alternative measure of international capital mobility adopted by the relevant literature is the Feldstein-Horioka condition. To investigate the extent of international capital mobility, Feldstein and Horioka (1980) examine the cross-

[7] Clearly, it is assumed that the foreign expected rate of interest is determined exogenously. In other words, the country in question is assumed not to be large enough in world financial markets to affect the world interest rate.

sectional correlation between saving and investment across OECD countries. Feldstein and Horioka argue that if capital is mobile internationally, there need not be any correlation between saving and investment. In a world of perfect capital mobility, savings may be expected to flow from countries with a relatively high propensity to save to countries with relatively more favourable investment opportunities. Feldstein and Horioka, therefore, reason that the null hypothesis of perfect capital mobility implies a zero regression coefficient in a cross-section regression of the savings share of GDP on the investment share of GDP, ie.:

$$(I/Y)_i = \alpha + \beta (S/Y)_i + \epsilon_i \tag{2.5}$$

where I represents gross domestic investment, S is gross national saving, Y is GDP, α is a constant term, ϵ is the regression error and the subscript i is a country index. Feldstein-Horioka suggest that the regression coefficient β in (2.5) measures the degree of international capital mobility, with large values indicating low mobility and low values high mobility.[8]

The Feldstein-Horioka condition, however, can be proved to require a larger number of assumptions relative to the RIP condition (and therefore relative to the other measures of international capital mobility based on interest parity conditions), representing therefore the strongest criterion for financial markets' integration. Using a result due to Dooley, Frankel and Mathieson (1987) and then used - among others - by Frankel (1993) and Lemmen and Eijffinger (1995a), four assumptions

[8] In Appendix 2.1 we briefly show how the current account may be written as the difference between saving and investment.

underlie the Feldstein-Horioka definition of capital mobility. Formally, if the investment share of GDP is a linear function of the expected real rate of interest:

$$I_{i,t+k}/Y_{i,t+k} = \zeta_0 - \zeta_1 E_t(r_{i,t+k}) + \gamma_i \tag{2.6}$$

where the stochastic error term γ_i, which captures the remaining determinants of the domestic investment-GDP ratio, is uncorrelated with the national savings-GDP ratio in the same country, ie.

$$Cov[\gamma_i, (S/Y)_{i,t+k}] = 0 \tag{2.7}$$

and also assuming that the savings share of GDP is neither correlated with the expected foreign real interest rate nor with the deviations from real interest parity, ie.

$$Cov[E_t(r^*_{i,t+k}), (S/Y)_{i,t+k}] = 0 \tag{2.8}$$

$$Cov[E_t(r_{i,t+k} - r^*_{i,t+k}), (S/Y)_{i,t+k}] = 0 \tag{2.9}$$

then β in (2.5) should be zero. All the assumptions (2.6)-(2.9) must hold for the Feldstein-Horioka condition to be a valid test of perfect international capital mobility. In fact, most of the attacks led against the Feldstein-Horioka condition are concerned with the meaning of the covariances described above[9]. Dooley, Frankel and Mathieson (1987) decompose the covariance between investment and saving share of GDP as follows:

$$Cov\left(\frac{I_i}{Y_i}, \frac{S_i}{Y_i}\right) = Cov\left(\gamma_i, \frac{S_i}{Y_i}\right) - \zeta_1 Cov\left(E_t(r_i^*), \frac{S_i}{Y_i}\right) - \zeta_1 Cov\left(E_t(r_i - r_i^*), \frac{S_i}{Y_i}\right) \tag{2.10}$$

where the time subscripts t+k - which apply to all variables in (2.10) except for the stochastic error term γ_i - are dropped for clarity. Under the Feldstein-Horioka definition of perfect capital mobility, equation (2.10) equals zero. If (2.6) holds and therefore domestic investment depends only on the domestic interest rate, then the first covariance term on the right hand side of (2.10) is zero[10]. If the foreign expected rate of return $E_t(r^*_{i,t+k})$ is exogenously determined, however, the second covariance term is also zero. Finally, under this definition of perfect capital mobility, the third covariance must be zero because the first variable in the covariance is zero. Note, however, that for the Feldstein-Horioka regression to yield a zero covariance between investment and saving shares of GDP the RIP is *not necessarily* required, contrary to what is argued by some researchers (eg.

[9] For a more detailed discussion of the meaning attached to those covariances and of the criticisms raised by the relevant literature, see Lemmen and Eijffinger (1995a).

[10] The assumption that domestic investment depends linearly on the domestic interest rate is, however, very controversial and the relevant empirical evidence suggests that the statistical relationship between I and r is quite weak. In practice, the error term γ_i in (2.6) may be very large.

Blundell-Wignall and Browne, 1991). For example, equation (2.10) may equal zero simply because the covariances on the right hand side of (2.10) cancel out, even if none of the covariance terms is zero, in which case a zero covariance between saving and investment share of GDP would not imply perfect capital mobility. Nevertheless, the RIP is *not sufficient* for the Feldstein-Horioka regression to yield a zero β, because the covariance between γ_i and the saving share of GDP - and consequently the covariance between investment and savings shares - may be nonzero. Overall, therefore, the Feldstein-Horioka regression must be interpreted with caution, since it depends on three covariances, which must all be zero for yielding $\beta=0$ *and* implying perfect capital mobility.

From the discussion above, the least ambiguous evidence on international capital mobility comes from a comparison of nominal interest rates on onshore and offshore loans of the same currency: under perfect capital mobility, the interest rate offered on a three-month Italian lira deposit in Rome, for example, should be the same as the interest rate on a three-month Italian lira deposit in London. Studies based on interest parity generally suggest relatively high degrees of international capital mobility and especially for France, Italy, Germany, Japan, the US and the UK, the onshore and the offshore money market are extremely linked and, furthermore, the link has increased during the 1980s. This evidence is, however, generally contradicted by studies which accept the Feldstein-Horioka definition[11], to which we now turn.

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[11] For an excellent survey of the studies on interest rate differentials, see Obstfeld (1995).

2.2.2 The saving-investment literature

In their seminal paper, Feldstein and Horioka (1980) examine the cross-sectional correlation between saving and investment across OECD countries and test the null hypothesis of perfect capital mobility by testing for a zero regression coefficient in a cross-section regression of the savings share of GDP on the investment share of GDP of the form (2.5). Under the null hypothesis, the regression coefficient β measures the degree of capital mobility, with large values indicating low mobility and low values high mobility.

Empirically, Feldstein and Horioka (1980) find that saving and investment shares of GDP across OECD countries are very highly correlated. Indeed the estimated coefficients are generally close to and insignificantly different from unity, and significantly different from zero. This leads them to conclude that the level of capital mobility in the world economy is low. Using a variety of techniques, numerous subsequent researchers have also found a high and significant correlation between saving and investment across countries[12]. Studies employing instrumental variables estimators, in order to avoid the possibility of simultaneity bias due to the endogeneity of savings, do not contradict the original finding of a high correlation of domestic savings and investment (Dooley, Frankel and Mathieson, 1987; Bayoumi, 1990; Tesar, 1991).

Nevertheless, for reasons formally explained in the Section 2.2.1, while a number of researchers have confirmed the empirical findings of Feldstein and Horioka, some have argued that these results do not necessarily imply low capital

[12] For a comprehensive survey of the literature, see Coakley, Kulasi and Smith (1995a).

mobility in the international economy. But if savings and investment need not be closely related because capital is mobile, it is difficult to rationalise the Feldstein-Horioka finding - corroborated in many similar studies - that savings and investment are in fact highly correlated.

A number of alternative hypotheses have been used to rationalise the Feldstein-Horioka regression result without concluding that capital is immobile. For example, saving and investment may be highly correlated as a result of endogenous private sector responses to private shocks (eg. population, technology, growth or productivity) which may simultaneously raise savings and investment (Obstfeld, 1986; Baxter and Crucini, 1993; Razin, 1995).

Real business cycle models provide neither unequivocal support nor rejection of the Feldstein-Horioka interpretation. For example, Tesar (1993) calibrates a dynamic model in which consumer preferences over traded and non-traded goods and over the intertemporal allocation of consumption lead to a significant correlation between saving and investment, even under perfect capital mobility. By contrast, van Wincoop and Marriman (1993), after asserting that it is not possible to explain the significant saving-investment correlation by simply referring to shocks to technology in a model assuming perfect capital mobility, argue that the cause of the correlation is the immobility of capital across countries.

Alternatively, the identity which links savings and investment to the current account of the balance of payments may have induced the high saving-investment correlations, for instance because governments target the current account (Summers, 1988; Bayoumi, 1990; Artis and Bayoumi, 1991; Ballabriga, Dolado and Vinals, 1991; Argimon and Roldan, 1994; Coakley, Kulasi and Smith, 1995b; Ghosh, 1995; Glick and Rogoff, 1995). The central argument of this strand of the literature is that governments operationalise the long-run solvency constraint on the current account in order to avoid Ponzi financing, biasing therefore the correlation between saving and investment towards unity, regardless of the degree of capital mobility (Coakley, Kulasi and Smith, 1995a,b). This may also help explain the fact - discussed below - that the short-run saving-investment correlation is generally found to be lower than the long-run correlation, although the converse should be implied by the Feldstein-Horioka interpretation (Coakley, Kulasi and Smith, 1995a).

Other researchers emphasise capital market imperfections to support the Feldstein-Horioka interpretation - eg. Obstfeld (1995), who discusses hysteresis in factor supplies, or Stefani (1994), who explains the association of saving and investment by developing a model of investment under credit rationing.

Finally, evidence against the interpretation of capital mobility suggested by Feldstein and Horioka is provided by some neoclassical growth models, in many of which human capital - second in immobility only to physical capital - plays a central role (Mankiw, Romer and Weil, 1992; Barro, Mankiw and Sala-i-Martin, 1995).

Further support in favour of the Feldstein-Horioka hypothesis is, however, provided by studies which compute saving-investment correlations across regions within a highly financially integrated area such as a single country. Bayoumi and Rose (1993), for example, compute saving-investment correlations across regions of the UK for the period 1971-85. In contrast to the Feldstein-Horioka results across OECD countries, Bayoumi and Rose (1993) find that the regression coefficient of UK regional investment on UK regional saving is generally statistically insignificantly different from zero. The Bayoumi-Rose results thus confirm that saving and investment shares of GDP need not be correlated in a financially integrated economy. Similar results are reported by Bayoumi and Sterne (1993) using regional data for Canada, and by Obstfeld (1995), who uses data for 45 different Japanese prefectures. Moreover, Sinn (1992), who looks at saving-investment correlations for 48 US states and Alaska in both 1953 and 1957, finds a negative but statistically insignificant cross-sectional correlation. Also, in Section 2.6, we provide empirical support to the idea that capital mobility is higher within a national economy than internationally, by making a comparison of GNP/GDP ratios across the major European countries with similar computations executed across different UK regions[13].

Overall, therefore, the debate concerning the "Feldstein-Horioka puzzle" has strongly deepened at a theoretical level and, although a number of alternative solutions proposed by the literature on the saving-investment correlation reject the interpretation of low or zero international capital mobility, a strand of literature still provides support to a modified view of only partial international capital mobility. While saving-investment regressions should be interpreted with caution, they may still shed some light on the issue of international capital mobility, especially when complemented by other statistical evidence.

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[13] Nevertheless, international capital mobility is expected to increase in moving towards a single market, one provision of which was the widespread abolition of exchange controls within the European Union. Bhandari and Mayer (1990), Artis and Bayoumi (1991) and Feldstein and Bacchetta (1991) - among others - argue that the progress in informational and institutional links is expected to increase financial flows within the European Union relative to the whole OECD and, looking at correlations within the European Monetary System, arrive at similar conclusions.

2.2.3 Saving-investment correlations: transitory versus permanent

In order to abstract from cyclical effects Feldstein and Horioka (1980) and others average the data over successive subperiods. In general, the results obtained have been extremely robust to the choice of the length of subperiods. Many researchers have used long-term averages of the data, both to capture the adjustment of investment to a *sustained* change in savings and in order to circumvent the problem that annual data estimations may be subject to some bias even in a cross-section context if business cycles are not synchronised internationally (Bayoumi, 1990). Nevertheless, Sinn (1992) argues that the savings-investment correlation coefficient, estimated using long-term averages, is biased towards non-rejection of the null hypothesis of immobility of capital. Calculations based on annual data do in fact tend to give lower correlation coefficients, although still suggesting an increase in capital mobility over time in the major industrialised countries.

Time-series estimates have also been undertaken in the literature. Although they are perhaps subject to greater bias than cross-section regressions because of the problems of identification and estimation induced by the endogeneity of both savings and investment, time-series estimations are considered very important as a guide to the appropriateness of imposing the same correlation across different countries (Obstfeld, 1986, 1995). The results obtained using time-series estimation, in fact, present a great range of dispersion across countries (Coakley, Kulasi and Smith, 1995a).

Nevertheless, even if initially time-series estimates seemed less supportive of the Feldstein-Horioka findings, Coakley, Kulasi and Smith (1994) show that the average of the long-term estimate of the Feldstein-Horioka coefficient across countries is very close to the standard cross-section estimate. Some authors argue that the time-series estimate of the saving-investment correlation coefficient, especially when the data is expressed in first differences, may be likely to capture the short-run transitory responses of investment to savings, whereas cross-section estimates may pick up mainly the long-run permanent response (Coakley, Kulasi and Smith, 1995a). Following this argument, it is considered a "fairly general result" (Coakley, Kulasi and Smith, 1995a) that the short-run estimates of the savings-investment coefficients are smaller than the long-run ones. This result has been heavily stressed by the literature supporting current account targeting or the solvency constraint argument as explanations of the Feldstein-Horioka puzzle, since these approaches imply that the short-run saving-investment correlation should be lower than the long-run correlation, whereas the converse should be found as an implication of the Feldstein-Horioka interpretation (Coakley, Kulasi and Smith, 1995a).

Nevertheless, Feldstein (1983) provides a theoretical benchmark based on a portfolio balance type model in order to explain why the short-run (temporary) effect of savings on investment may be expected to be smaller than the long-run (permanent) effect. The argument proposed is that the short-run capital flow is only part of a once-for-all adjustment of the international portfolio adjustment and that, once this adjustment is completed, the flow of capital tends to a lower level, determined by the rate of growth of the world capital stock and by the share of international assets held in the equilibrium portfolio (Feldstein, 1983, p. 147).

This - perhaps *ad hoc* - explanation of the underlying dynamics of saving and investment has not, however, received widespread support (Coakley, Kulasi and

Smith, 1995a), but very little work exists on the theoretical interpretation of the difference between the short-run and long-run savings-investment coefficients. In particular, the finding of a relatively lower short-run correlation coefficient seems inconsistent with the Feldstein-Horioka interpretation of the savings-investment correlation coefficient (Coakley, Kulasi and Smith, 1995a). The argument is that a *transitory* increase in savings may be expected to stay at home since it may not be worth incurring the costs of analysing foreign investment opportunities or evading or avoiding exchange controls. In contrast, if a *permanent* increase in savings occurs, then these costs to mobility are more likely to be countenanced by economic agents and a larger proportion of the permanent increase in savings may be expected to flow abroad.

In the following sections we describe and employ a rigorous estimation strategy in order to analyse empirically - using UK data - the difference between short-run and long-run estimates of the savings-investment coefficient.

2.3 Estimation techniques

Blanchard and Quah (1989) suggest an econometric technique which enables researchers to identify temporary and permanent shocks in the context of a multivariate time series model. Using the bivariate vector autoregressive representation (BVAR) of two variables w(t) and z(t), expressed in stationary form, and assuming the long-run restriction that the cumulative impulse response function of w(t) to temporary shocks goes to zero, Blanchard and Quah show how restrictions on the BVAR may be imposed in order to recover the temporary and permanent shocks driving the series - see Appendix 2.2 for a more technical outline of the Blanchard-Quah decomposition.

Blanchard and Quah analyse US data on real output and unemployment, and motivate the method using a small stylised macro model with a long-run vertical supply curve. Hence, they identify temporary shocks with demand innovations and permanent shocks with supply innovations. Such a taxonomy is not, however, necessary in the present context.

In Section 2.5, we employ the BQ decomposition to decompose saving and investment shares of GDP into their temporary and permanent components.

2.4 Data on UK saving, investment and GDP

Quarterly UK data were obtained on gross savings as a ratio of personal disposable income, gross fixed capital formation and GDP from *Economic Trends* (Central Statistical Office, CSO). The data cover the period from 1955Q1 to 1994Q4. From this data set, savings and investment shares of GDP were computed, consistent with much of the literature on saving-investment correlations (Coakley, Kulasi and Smith, 1995a,b).

2.5 Empirical results

2.5.1 Saving and investment in the UK

The first exercise involves the ordinary least squares (OLS) estimation of the regression coefficient and the correlation coefficient in the Feldstein-Horioka equation with saving and investment shares both in levels and first differences. We also examine the effect of the abolition of exchange controls by the UK government in October 1979. In fact, one would expect saving and investment shares of GDP to have been affected by the removal of barriers to capital flows, and perhaps greater international capital mobility to have been induced.

Taylor and Tonks (1989) assess the impact of the abolition of exchange control on the degree of integration of UK and overseas stock markets, providing evidence that a marked increase in the degree to which these markets move together in the long run occurred after 1979. Also, Artis and Taylor (1990) examine the effect generated by the removal of exchange controls in 1979 in the UK, showing that this action contributed to move towards the elimination of the deviations from onshore-offshore interest rate parity: a greater degree of financial integration appears, therefore, to have been induced by the abolition of exchange rate controls.

In order to test for a structural break in the Feldstein-Horioka coefficient, we employ a Chow test for structural stability. The result - reported in Table 2.1 - enables us to reject the null hypothesis of no structural break in the estimated coefficient in the regression of investment on savings shares. It seems therefore reasonable to estimate the regressions not only for the full sample period, but also for the two subperiods 1955Q1-1979Q4 and 1980Q1-1994Q4, in order to take account of the structural break.

The estimated Feldstein-Horioka coefficients and the correlation coefficients on both regressions in levels and in changes and for the three sample periods considered are reported in Table 2.1. For all sample periods considered, the estimated Feldstein-Horioka coefficients are - as expected - rather lower than the ones generally reported from cross-section estimates. They are, however, always significantly different from zero at conventional nominal levels of significance in the levels regression. Also, the estimated coefficient is much higher before than after the abolition of capital controls, and, further, the post-abolition estimated coefficient is significantly *negative*. Conversely, in the regression of the variables in first differences, the estimated coefficient is insignificantly different from zero apart from the pre-abolition sample period. All the point estimates using first differences are, however, negative.

The correlation coefficients are generally higher than the Feldstein-Horioka coefficients and, for savings and investment shares in levels, they are statistically significant at conventional nominal levels of significance. Most tellingly, the correlation in the subperiod 1955Q1-1979Q4 is relatively high (0.755), whereas the correlation in the second subperiod is statistically significant and negative (-0.574).

The results are consistent with the literature in that the estimated coefficients may largely vary from one period to another and it is interesting that they may be relatively high also in time series estimation, as it is the case in both subperiods. Following the interpretation of Feldstein and Horioka, the results indicate a strong increase in the degree of capital mobility after 1979.

The latter result is consistent with recent work by Lemmen and Eijffinger (1995a), who provide evidence that savings and investment shares of GDP are in fact cointegrated before 1979, not after 1979, using the augmented Dickey-Fuller test statistic on both annual and quarterly UK data. This evidence may suggest in fact that savings and investment contain common trends before the abolition of capital controls, while their abolition made it more feasible to run large, persistent current account imbalances after this date. After the abolition of capital controls, therefore, targeting the current account by government has perhaps become more cumbersome. In general, in a world of limited capital mobility, the government can adjust its own saving-investment balance to match changes in the private sector balance, using - for example - appropriate fiscal policy. In internationally financially integrated economies, however, the possibility of targeting the current account seems less feasible (Artis and Bayoumi, 1991; Summers, 1988). This result is also consistent with the implications of the intertemporal approach to the current account balance, suggesting that the abolition of barriers and impediments to capital movements, enhancing greater capital market integration, *ceteris paribus*, make current account imbalances more likely and the predictions of the intertemporal approach more applicable (Sachs, 1981, 1982; Frenkel and Razin, 1987; Obstfeld and Rogoff, 1995a; Razin, 1995).

A general point, however, needs to be made against the widely used technique of using the Feldstein-Horioka regression for quantifying the degree of international capital mobility. The two variables in the estimated regression are generally expected to be and - in the next section shown to be - integrated of order one, I(1) (Coakley, Kulasi and Smith, 1995a). This implies that both the Feldstein-Horioka coefficient and the correlation between the two variables in levels are - unless the variables in question are cointegrated - spurious and unreliable, both in time series and cross section estimates (Granger and Newbold, 1974; Pesaran and Smith, 1995). It might, therefore, seem more appropriate to look at the regression in first differences rather than in levels. As Table 2.1 (Panel B) shows, such an analysis would suggest a complete lack of correlation between savings and investment. Upon reflection, however, it is clear that those regressions should also be treated with caution, since they do not adequately distinguish between temporary and permanent movements in the underlying series.

2.5.2 *Decomposition of savings and investment shares*

In order to identify temporary and permanent shocks in the variables in question, we employ the Blanchard-Quah decomposition in BVARs with the first difference of the logarithm of nominal GDP as a second variable. As noted by Blanchard and Quah (1989), the decomposition can be applied to any two series whose time series behaviour is jointly determined.

Modelling saving poses both practical and theoretical challenges to economists. Differences in saving rates are often considered responsible for differences in output growth across countries and in this context the decline in saving rates recently experienced by many industrialised countries is particularly worrying (Modigliani, 1993)[14]. Moreover, where the old "exogenous growth" literature attributed to saving a permanent effect on the level of per capita income, but only a transitory effect on the rate of growth of per capita income, post-neoclassical endogenous growth models usually assume that saving rates may generate permanent increases in both the growth and the level of income in an economic system (eg. Romer, 1986, 1990; Lucas, 1988; Barro, 1990; Barro and Sala-i-Martin, 1990; Sala-i-Martin, 1990a, 1990b; Grossman and Helpman, 1991; Helpman, 1991; Buiter, 1993). Further, a large amount of recent work on microeconomic data has been done to identify the determinants of saving and some degree of consensus is emerging in that the most commonly suggested determinant of the saving rate of a country is the growth rate of output (eg. Lau, 1993; Fitoussi

[14] In the growth model of Harrod and Domar, for example, the rate of economic growth is governed by the quotient of the savings ratio and the capital output ratio, so that there is a direct relationship between saving and economic growth.

and Le Cacheux, 1993; Ickes, 1993; Horioka, 1993).

The relationship between investment and output is also heavily discussed in the literature, going back to accelerator type models where the level of investment in an economy is modelled as a function of the rate of change of output (Samuelson, 1939). Post-neoclassical endogenous growth models also predict a positive and dual relationship between capital accumulation and output. Therefore shocks which impinge upon output and output growth are likely to have much in common with those that affect investment.

Overall, therefore, nominal GDP appears to be a variable which is likely to be determined jointly with both savings and investment. It seems therefore appropriate to apply the BQ decomposition to investment and savings using GDP as the second variable.

Preliminary unit root tests are reported in Table 2.2. In no case can we reject the hypothesis that each of the series in a realization of a stochastic process containing one unit root, at the five percent nominal level of significance.

In Table 2.3 we also report the results of testing for cointegration between investment and savings ratios in the full sample period and in the two subperiods before and after the abolition of UK exchange controls (Engle and Granger, 1987). At the five percent nominal significance level the ADF test statistic fails to reject the null hypothesis of no cointegration between savings and investment ratios in all sample periods considered.

Figures 2.1 and 2.2 show the cumulative impulse response functions of saving and investment shares respectively, obtained from the execution of the BQ decomposition employed in a BVAR with the logarithm of GDP as a second variable. A positive temporary shock to the savings ratio (Figure 2.1), in both subperiods, increases the savings ratio in the short run, while producing a long-run cumulative effect of zero. As would be expected, a positive permanent shock to savings raises the savings ratio. The effect of a temporary savings innovation is to cause a short-term variation in the output level, which dissipates and settles down to a steady-state long-run positive cumulative effect. A positive permanent saving shock also generates a permanent increase in the level of output.

The impact effect on investment of a temporary shock to the investment ratio (Figure 2.2), as one might expect, declines to zero, but only after a quite long period of time, reflecting the identifying restriction imposed but also suggesting that temporary shocks may be quite persistent. The effect of a positive permanent investment shock is, conversely, a permanent increase in investment. For both sub-periods, both temporary and permanent innovations to the investment ratio lead to a permanent increase in output and, as one would expect, the output effect is greater for the permanent investment innovations.

Overall, therefore, the decomposition of saving and investment shares suggest that strong similarities exist in the dynamics of the two processes responding to both the permanent and the temporary shocks impinging upon them.

2.5.3 Correlations between temporary and permanent components of the change in savings and investment shares

Having identified the temporary and permanent shocks to savings and investment shares, we can now construct the paths of those variables which would have been obtained in the absence of temporary or permanent shocks. Effectively, this involves setting either the transitory or the permanent shocks to zero in the moving average representation.

We can then compute cross correlation coefficients between the resulting temporary and permanent components of the change in saving and investment shares in order to shed light on the issue of capital mobility, on the implications which we should expect from the analysis of saving-investment correlations and also on the validity of this measure for quantifying the degree of international capital mobility. As already stressed in Section 2.5.1, correlations of the levels of the shares of GDP or on their permanent components, which are I(1), are spurious. Only the correlation of the stationary, temporary components is meaningful and should be considered. Conversely, the correlation of the temporary and permanent shocks to movements in savings and investment shares, ie. of the temporary and permanent components of the change in the variables in question, is always meaningful, since we are dealing with stationary processes.

Table 2.4 reports the correlations between both the temporary and the permanent components of the change in saving and investment shares for the three sample periods considered. The results are extremely interesting. Correlations for the full sample period are relatively high for both components, suggesting quite a strong and immediate association of the two processes. Nevertheless, perhaps the most interesting result is that the temporary components are more highly correlated ($r_{is}=0.43$) than the permanent components ($r_{is}=0.29$). This finding provides *prima facie* support to the view that if the Feldstein-Horioka interpretation of the relationship between savings and investment is correct, then temporary components of the series should be more highly correlated than the permanent ones, ie. the short-run correlation coefficient should be higher than the long-run coefficient.

Table 2.4 also reports the correlations between savings and investment components for the two subperiods 1955Q1-1979Q4 and 1980Q1-1994Q4. If the savings-investment association has some power in highlighting the degree of international capital mobility, then one would expect the correlation to be lower in the second subperiod, as an effect of the abolition of UK exchange rate control. The results do indeed display a very large and significant difference between the two correlation coefficients. The correlation between the temporary components of the change in savings and investment shares declines from 0.72 in the first subperiod to 0.32 in the second subperiod, while the correlation between the permanent components is lowered from 0.54 to 0.18 from one subperiod to another. The permanent components are in all cases less correlated than are the temporary components. Most tellingly, the correlation coefficient for the permanent components of investment and saving is, on the basis of the t-statistic, significantly different from zero before 1980 (t-ratio of 6.12) but is insignificantly different from zero after the abolition of exchange control (t-ratio of 1.39). Correlation coefficients between the temporary components are strongly significantly different from zero in both cases.

As a matter of statistical precision, we also compute a test statistic for equality of the correlation coefficients across the two subperiods. The test statistic employed is constructed as follows. If r_{is} is the sample correlation coefficient of the change in saving and investment shares, then the statistic:

$$\zeta = \frac{1}{2} \log \left(\frac{1 + r_{is}}{1 - r_{is}} \right) \tag{2.11}$$

is approximately normally distributed with mean and variance equal to $(1/2)\log[(1+\rho_{is})/(1-\rho_{is})]$ and $1/(n-3)$ respectively, where ρ_{is} represents the population correlation coefficient between savings and investment and n is the sample size. The approximation is close for sample sizes larger than fifty (Kendall and Stuart, 1967). Therefore, under the null hypothesis of equality of the two population correlation coefficients, the test statistic:

$$\xi = 0.5 \frac{\log\left(\frac{1+r_{is1}}{1-r_{is1}}\right) - \log\left(\frac{1+r_{is2}}{1-r_{is2}}\right)}{\left(\frac{1}{n_1-3}\right) + \left(\frac{1}{n_2-3}\right)} \tag{2.12}$$

(where the subscripts 1 and 2 attached to r and n refer to the first and second subperiod respectively) is distributed approximately standard normal.

For both temporary and permanent components, the test statistic for equality of the correlations, reported in Table 2.4, is very high (a marginal significance level of virtually zero), implying a very marked decrease in the correlation between the two periods.

Overall, therefore, the results of this section demonstrate that the correlation coefficients between the temporary components of changes in saving and investment ratios is high relative to that between the permanent components. There was, however, a significant reduction in both of these correlations consequent upon the abolition of UK capital controls in 1979 and, indeed, the correlation between the permanent components of movements of UK savings and investment ratios is statistically insignificantly different from zero after 1979.

2.6 GNP/GDP ratios for Europe and the UK

In order to supplement the Feldstein-Horioka regressions and gain further insight into the nature of capital mobility within Europe, we perform some calculations on GNP/GDP ratios for seven European Union (EU) countries, using annual data on GNP and GDP taken from the *OECD National Accounts* for the sample period 1964-1992. This ratio gives an indication of whether a country is a net receiver or remitter of factor income from abroad. It is, thus, a measure of the extent to which net capital positions diverge across EU countries.

Table 2.5 lists the percentage deviation from the group average of the GNP/GDP ratio for the seven largest European countries for four-year subperiods from 1964 through 1992.

The data illustrate some interesting regional patterns. Italy and, more significantly, Spain, two fast-growing European economies, have negative ratios throughout the period, indicating that they were net borrowers of capital from other European economies. By contrast, the most mature and slowest growing economy, the UK, has the largest positive ratio in the group considered, indicating that it has been a net creditor. Also France, to a less extent than the UK, displays a positive ratio. Germany has switched position, from debtor in the 1960s to creditor in the 1980s, presumably reflecting the slowdown in its growth relative to the rest of Europe, displaying a significantly large ratio for the whole sample period. In

Belgium, the data appear dominated by the high level of government debt built up in the late 1970s and 1980s, which have made it a debtor over the last two decades. Finally, the Dutch have overall maintained a very balanced GNP/GDP ratio, switching from a position of net debtor in the 1960s and 1970s to one of net creditor in the 1980s.

It seems instructive to compare the results obtained for different European countries with similar calculations computed for different regions within a single European country. Table 2.6 shows the results obtained from similar calculations which we carried out for geographical regions within the UK for five-year subperiods from 1972 through 1990.[15] Such a comparison seems interesting since, like the Bayoumi-Rose result on saving-investment correlations for the UK, it provides a benchmark comparison of the results for European countries with a completely financially integrated area.

Although data on GDP are available at regional level, figures on GNP are not. As a proxy for GNP we use total personal income (TPI) adjusted for government transfer payments. The main difference between TPI and GNP is retained corporate profits, which appear unlikely to have a significant regional pattern.

Table 2.6 shows ratios of TPI (net of government transfer payments) to regional product, calculated as percentage deviations from the UK average. The first thing to note about the data is that the ratios within the UK are considerably larger than those across European countries. The data generally diverge by at least

[15] Data taken from the UK Central Statistical Office *Regional Trends*, various issues.

one percent from the average. By contrast, in the European data only exceptionally does a country present a GNP/GDP ratio which diverges from the average by more than one percent. In particular, the South West and, to a smaller extent, East Anglia and the South East, display very large deviations from the UK average. Overall, therefore, it appears that net capital positions are considerably more diverse across the UK monetary union than they are across European countries.

Regional differences in the UK appear largely to reflect economic specialization. Interestingly, heavily industrialised areas, such as the North and North West are net debtors, while regions which are more service oriented and with larger retirement communities, such as the South East, South West and East Anglia, are the most significant net creditors. Wales has shown a significant change in its position over the sample period, seeing its percentage deviation from the UK average switching from a mildly positive value in the 1970s and a heavily positive value in the early 1980s (presumably reflecting the decline of much of its traditional heavy industry, in particular coal mining) to a significantly negative value in the second half of the 1990s, displaying an overall position of net debtor throughout the whole period. Scotland and Northern Ireland also display a negative percentage deviation (probably largely due to the remittance of capital earned from emigrants) while England is the only nation with a positive deviation from the UK average. Overall, it appears that, within the UK, industrial regions have widely used open capital markets to borrow capital from outside of that particular region.

Some simple calculations may underline this point. The data in Table 2.6 indicate that in the UK the average deviation of the ratio of total personal income to regional product from the national average is 1.3 percent. Assuming that this reflects capital income flows and that the interest rate is 10 percent, this implies that the average net capital position of a UK region represents 13 percent of its regional product. Put another way, the average UK region has a net capital position of plus or minus 13 percent of its regional output. By contrast, on the same assumptions, the average European country has a net capital position equal to only one fortieth of its output.

If the UK data is a useful guide, the continued financial integration of the European economy should be expected to make capital ownership less insular. This has at least three implications. First, such capital flows imply significant regional current account imbalances, as implied by the intertemporal approach to balance of payments theory (Sachs, 1981; Frenkel and Razin, 1987; Obstfeld and Rogoff, 1995a). The removal of liquidity constraints due to financial integration should make the intertemporal approach more fully applicable and reduce current account imbalance to a minor role, in the same way in which, in the domestic economy, financial deregulation should make more applicable theories based on consumption-smoothing rather than liquidity-constrained consumers (Mathieson and Rojas-Suarez, 1990; Artis and Bayoumi, 1991). Second, as net factor flows become more important, there will be an increasing divergence between the output of a country and its level of income; in the UK, over four percent of the income in the South West is generated outside the region, while over one percent of the output in the North and North West flows to other regions. Finally, open capital markets may make it easier for individuals to use private capital markets to insure themselves against regional risk. It is to this topic that we now turn.

2.7 Private capital flows and regional risk

In addition to producing net capital movements between economies, transfer payments could also be used to reduce regional income fluctuations. Most of the discussion of transfer payments in this context has focused on the role of government transfers. For example, Sachs and Sala-i-Martin (1992) examine the contribution of a federal fiscal system to the stabilization of incomes in individual states or regions. They are able to show that, in the US, about 40 percent of any income shock is offset by autonomous changes in tax payments to the federal governments and endogenous changes in transfers from it.

Nevertheless, private capital markets may also play a role in reducing income fluctuations. In order to investigate this role, we first derive an empirical model which we then estimate using data for EU countries.

Consider a simple model with three types of individuals - capitalists, workers and capital investors. In what follows we shall refer to the *area* as the geographical area made up by the various *regions*. Capitalists derive all their income from the ownership of capital; since they have no income from labour they will invest in the average area portfolio, and hence their capital income will vary with area capital income. Workers, on the other hand, attempt to insure themselves against movements in their own local or regional labour income. They will choose a portfolio whose returns are negatively correlated with regional labour income and hence their capital income will be negatively correlated with regional labour income. Finally there are some capital investors who invest in the regional economy, due to factors such as superior information about local investment options. Their capital income is assumed to be correlated with that of the region.

Formally, indexing these three types of individuals by superscript 1, 2 and 3 respectively, the following equations can be derived:

$$\begin{aligned} \Delta I^1_{kit} &= \Delta I_{kat} \\ \Delta I^2_{kit} &= -\beta \Delta I_{lit} \\ \Delta I^3_{kit} &= \Delta P_{kit} \end{aligned} \tag{2.13}$$

where I_{kit} represents per capita income from capital in region i in period t, I_{kat} is area per capita income from capital in period t, I_{lit} is per capita income from labour in region i in period t, and P_{kit} is capital product per capita from region i in period t; Δ is the first difference operator. Thus, I represents personal income, while P represents income from the production side of the accounts.

If it is assumed that these three types of individuals each receive a fixed proportion of total capital income, then, writing these proportions as λ^{1i}, λ^{2i}, and $(1-\lambda^{li}-\lambda^{2i})$ for capitalists, workers and local investors respectively in region i, then from equations (2.13), the data generation process for the change in total capital income in region i becomes:

$$\Delta I_{kit} = \lambda^{1i}\Delta I_{kat} - \beta\lambda^{2i}\Delta I_{lit} + (1-\lambda^{1i}-\lambda^{2i})\Delta P_{kit} \tag{2.14}$$

Including a constant term in each equation, α^i, to account of unmeasured factors, produces estimating equations for regions i=b, . . ., h:

$$\begin{aligned}\Delta I_{kbt} &= \alpha^b + \lambda^{1b}\Delta I_{kat} - \beta\lambda^{2b}\Delta I_{1bt} + (1-\lambda^{1b}-\lambda^{2b})\Delta P_{kbt}\\ \Delta I_{kct} &= \alpha^c + \lambda^{1c}\Delta I_{kat} - \beta\lambda^{2c}\Delta I_{1ct} + (1-\lambda^{1c}-\lambda^{2c})\Delta P_{kct}\\ &\vdots \qquad\qquad\qquad \vdots\\ \Delta I_{kht} &= \alpha^h + \lambda^{1h}\Delta I_{kat} - \beta\lambda^{2h}\Delta I_{1ht} + (1-\lambda^{1h}-\lambda^{2h})\Delta P_{kht}\end{aligned} \tag{2.15}$$

These are the basic estimating equations used in this section of the chapter. Annual data on household income derived from capital and from labour and the operating surpluses of incorporated enterprises were obtained over the period 1964-92 from the OECD *National Accounts*. The data cover eight members of the EU: Germany, France, Italy, the United Kingdom, Belgium, the Netherlands, Spain and Greece. In the above terminology, each country is an individual region. The series were converted into real per capita terms using the respective GDP deflators, 1985 bilateral exchange rates (national currency per US dollars) and OECD population estimates. Area per capita income from capital, I_{kat}, is computed as the sum of per capita income from capital over the eight EU countries examined.

First, we estimate equations (2.15) jointly as a set of seemingly unrelated regressions (SUR) for all countries considered since the exogenous disturbance terms in these equations are likely to covary across countries (Zellner, 1962). It is a well known result that generalized least squares estimation of seemingly unrelated regressions is an efficient technique for analysing the effect of a common set of global factors across a group of countries where the disturbances in the individual regressions are expected to be correlated since they include some factors that are common to all of the countries as well as some factors that are specific to a particular country (Dwivedi and Srivastava, 1978).[16]

Table 2.7 shows the results from estimating simultaneously equations (2.15) as seemingly unrelated regressions across the European countries listed above.[17] Since the cross-equation restriction of equality of all the coefficients except the constant could not be rejected at conventional nominal levels of significance, we report only the results with these restrictions imposed. As Table 2.7 shows, the coefficient estimates are correctly signed and are all very strongly significantly different from zero at conventional nominal levels of significance.[18] Nevertheless, they are rather small in magnitude, and suggest that private capital markets provide a very limited degree of insurance against fluctuations in labour income. The estimate of the coefficient on the change in household labour income is -0.201, indicating that a one (constant 1985) dollar fall in labour income is associated with a rise in income from capital of twenty cents. Overall, it appears that, at least at present, private capital markets have a relatively small role in reducing regional income fluctuations across Europe. The estimated coefficient associated with the

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[16] See Appendix 2.3 for a technical discussion of SUR and GLS.

[17] A consistent estimate of the contemporaneous covariance matrix was obtained by first-stage ordinary least squares. This was then used to construct the SUR estimator.

[18] A battery of diagnostics (not reported to conserve space, but available from the author) was also executed on the eight regressions in the system - namely Durbin-Watson statistics, Ljung-Box test statistics for serial correlation, Jarque-Bera (1980) tests for normality of the residuals, Breusch-Pagan (1979) tests for heteroskedasticity, and tests for first order autoregressive conditional heteroskedasticity. Model adequacy was suggested in every case (none of the diagnostics was significant at the five percent level and high coefficients of determination - in the range between 0.82 and 0.95 - suggest that the model fits the data very well). Also, using a simple Chow test we found no evidence of a structural break at the mid-point of the sample period at conventional nominal levels of significance.

change in aggregate capital income is 0.006, but it is very significantly different from zero at conventional nominal significance levels. The coefficient on the change in operating surpluses (profits) is estimated at 0.205, perhaps smaller than one would expect. Fluctuations in income from capital within Europe do not appear to be strongly correlated with fluctuations in capital product, possibly because corporations smooth capital income flows.

In order to take account of potential country-specific random shocks across the panel, we also estimated an error components model formulation of the system. In the error components model, the country-specific intercept is assumed to be a random variable with constant mean.[19]

In Table 2.8, we report the results from estimating equations (2.15) as an error components model for the eight European countries examined. The coefficient estimates are correctly signed, very strongly significantly different from zero at conventional nominal levels of significance and close to the estimates reported in Table 2.7. A Hausman (1978) test for the applicability of the error components model (see eg. Judge, Hill, Griffiths, Lutkepohl and Lee, 1982, p. 498) suggests, in fact, that it is the more appropriate model. The estimate of the coefficient on the change in household labour income and on the change in aggregate capital income are equal to -0.201 and 0.006 respectively, identical to the estimates given in Table 2.7 from the estimation of equations (2.15) as seemingly unrelated regressions. The coefficient on the change in operating surpluses is

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[19] For an excellent description of SUR, fixed and random effects models and the underlying assumptions about the structure of the covariance matrix made in those models, see Greene (1993, Chapters 16,17) or Judge, Hill, Griffiths, Lutkepohl and Lee (1982, Chapter 16).

estimated, however, at 0.131. This is even smaller than the point estimate obtained from the SUR estimation, reinforcing the finding that fluctuations in income from capital within Europe are not strongly correlated with fluctuations in capital product.[20]

The implications of these results is that private capital markets currently provide relatively little insurance across EU countries. Evidence from the US (Atkeson and Bayoumi, 1993) indicates that financial integration may well produce more geographic diversification of investment portfolios within the EU, in turn providing some insurance against fluctuations in regional output. However, such private markets will probably continue to provide a partial degree of insurance against regional fluctuations.

Nevertheless, there remain some ambiguities affecting the comparison between regional risk-sharing within nations with risk-sharing among nations. As noted by Obstfeld (1995), given incompleteness of asset markets, the appearance of relatively greater risk sharing within nations may be due to the predominance of non-insurable idiosyncratic shocks. Moreover, redistributive domestic fiscal policies may be effective in pooling risks within national borders (Obstfeld, 1995). Despite these drawbacks, however, approaches based on comparisons of regional and international differences provide a better understanding of how international and intranational financial linkages differ and throw light on the degree of capital mobility present in the economy.

[20] Footnote 18 also applies here, except that the coefficients of determination were in excess of 0.9.

2.8 Conclusions

In this chapter we have investigated a number of related issues concerning capital markets' integration in Europe. In particular, we examined the difference between the short-run and the long-run saving-investment correlation coefficient, in order to shed light both on the validity of the Feldstein-Horioka regression as a means of measuring the degree of international capital mobility and on its implications. This task was accomplished through a statistical investigation of the relationship between transitory and permanent components of movements in savings and investment shares of GDP on UK data. In fact, estimation of the regression of investment shares on savings shares - both non-stationary - produce spurious estimates and are therefore unreliable, unless a cointegrating relationship between the two variables is established. Our results indicate that the short-run correlation is statistically significantly higher than the long-run correlation, in the sense that the temporary components are relatively more highly correlated. This suggests that transitory movements in savings, given the frictions existing in international transactions, are more likely to remain in the UK, since it may not be worthwhile facing the cost of analysing foreign investment opportunities or evading or avoiding exchange controls. Conversely, when an increase in savings occurs which is perceived to be permanent, then it is more likely to flow abroad.

Overall, the change in savings and investment shares, which are not found to be cointegrated in any sample period considered, display correlations which strongly suggest that the Feldstein-Horioka definition of international capital mobility, correctly interpreted, provides a valid way to shed light on the degree of international capital mobility. In fact, the correlation, both between temporary (short-run) and permanent (long-run) components of the change in savings and investment shares are significantly lower in the period after the abolition of UK exchange controls. Moreover, the correlation between the permanent components in the post-1979 period is not significantly different from zero at the one percent nominal level of significance, suggesting a very high degree of capital mobility when the change in savings is perceived to be permanent[21].

In addition, we gauge the extent of financial markets' integration with regard to the past, present and prospective nature of capital flows in Europe. A comparison of GNP/GDP ratios across Europe with TPI/GDP ratios across regions of the United Kingdom indicated that, as European financial integration continues, one may expect an increasingly diverse pattern of capital ownership across the EU to emerge. In turn, this may be expected to enhance higher intra-EU current account imbalances, generating a greater possibility of European output and income measures diverging and greater opportunities for individuals to insure themselves against region-specific risk. At the present time, however, our empirical work suggest that private capital markets currently provide relatively little insurance across EU countries.

[21] This finding is also consistent with some recent evidence from Euler equation tests, which suggest that the UK is one of the most highly financially integrated countries in Europe (Lemmen and Eijffinger, 1995b).

Appendix 2.1 Savings, investment and the current account

The Feldstein-Horioka regression is based on the parametrisation of the current account as the difference between national savings and investment. This parametrisation may de derived using elementary manipulation of the balance of payments identities:

$$S = GNP - C + NCT \tag{A2.1.1}$$

$$GNP = C + I + X - M + NFI \tag{A2.1.2}$$

Substitution of (A2.1.2) onto (A2.1.1) yields:

$$S = I + X - M + NFI + NCT \tag{A2.1.3}$$

with

$$I = FCF + ST \tag{A2.1.4}$$

where S denotes gross national savings, I is gross domestic investment, C is total private and government final consumption expenditure, M and X denote import and export of goods and services respectively, GNP is gross national product, NCT denotes net current transfers from the rest of the world, NFI is net factor income from the rest of the world, FCF is gross fixed capital formation and ST denotes increase in stocks.

Using (A2.1.1)-(A2.1.4) yields:

$$(X-M+NFI+NCT) = (I+X-M+NFI+NCT) - (FCF+ST) \tag{A2.1.5}$$

or

$$CA = S - I \tag{A2.1.6}$$

where CA denotes the current account of the balance of payments.

Also, Artis and Bayoumi (1991) demonstrate also how the current account balance may be written as the sum of private and public sector savings-investment balances:

$$CA = (S_p - I_p) + (S_g - I_g) = S - I \tag{A2.1.7}$$

Then, splitting the statistical discrepancy equally between savings and investment and dividing both regressors in (A2.1.6) by GDP[=GNP+NCT] yields exactly the Feldstein-Horioka regression.

Appendix 2.2 The Blanchard-Quah decomposition

Consider a 2x1 vector of macroeconomic variables, X(t)=(1-L)[w(t) z(t)]' where L denotes the lag operator. Both (1-L)w(t) and (1-L)z(t) are assumed to be realisations at time t from stationary stochastic processes with - for the purposes of exposition - their deterministic components removed. Both variables w(t) and z(t) are therefore assumed to be integrated processes of order one, I(1) and not cointegrated[22]. By employing the multivariate form of Wold's decomposition, X(t) will have an infinite order moving average representation. Blanchard and Quah consider a transformation of the Wold representation of the following form:

[22] Blanchard and Quah consider a system in which z(t) is stationary. For our purposes, however, it is more convenient to define z(t) as an integrated process.

$$(1-L)\begin{bmatrix} w(t) \\ z(t) \end{bmatrix} = \begin{bmatrix} \Psi_{11}(L) & \Psi_{12}(L) \\ \Psi_{21}(L) & \Psi_{22}(L) \end{bmatrix} \begin{bmatrix} \eta_1(t) \\ \eta_2(t) \end{bmatrix} \qquad \text{(A2.2.1)}$$

or

$$X(t) = \Psi(L)\eta(t) \qquad \text{(A2.2.2)}$$

where $\eta(t)=[\eta_1(t)\ \eta_2(t)]'$ is a 2x1 vector of innovations occurring at time t, and

$$\Psi_{ij}(L) = \sum_{k=1}^{\infty} \psi_{ij}(k) L^k \qquad \text{(A2.2.3)}$$

so that $\psi_{ij}(k)$ (i,j=1,2) denotes the impulse response of the i-th element of X(t) to the j-th element of $\eta(t)$ after k periods. If the innovation $\eta_1(t)$ is to have only a temporary effect on w, then the following restriction must hold:

$$\Psi_{11}(1) = 0 \qquad \text{(A2.2.4)}$$

The p-th order vector autoregressive representation for X(t) yields a vector of innovations $\nu(t)$:

$$[I - \Theta(L)]X(t) = \nu(t) \qquad \text{(A2.2.5)}$$

where $\Theta(L)$ is the 2x2 matrix of estimated polynomials in the lag operator. Since X(t) is stationary, this can be inverted to obtain the estimated moving average representation:

$$X(t) = [I - \Theta(L)]^{-1}\nu(t) \tag{A2.2.6}$$

or

$$X(t) = C(L)\nu(t) \tag{A2.2.7}$$

From (A2.2.6) and (A2.2.7) it is clear that C(0) = I. Equating coefficients in (A2.2.2) and (A2.2.7) implies that the VAR innovations will be linear combinations of the underlying temporary and permanent shocks:

$$\nu(t) = \Psi(0)\eta(t) \tag{A2.2.8}$$

Since $\Psi(0)$ is a (2x2) matrix, Blanchard and Quah (1989) derive four restrictions in an attempt to achieve identification.

Three restrictions can be obtained by normalizing the variances of the temporary and permanent shocks, $\eta_1(t)$ and $\eta_2(t)$, to unity and requiring them to be orthogonal. Let Ω be the variance-covariance matrix of ν_t, then, using (A2.2.8), this implies the following three restrictions:

$$\Psi(0)\Psi(0)' = \Omega \tag{A2.2.9}$$

From (A2.2.2), (A2.2.7) and (A2.2.8) we can deduce the impulse response functions in terms of C(j) and $\Psi(0)$:

$$\Psi(j) = C(j)\Psi(0) \tag{A2.2.10}$$

Finally, using (A2.2.4) and (A2.2.10) we can deduce a fourth restriction on $\Psi(0)$:

$$\kappa' \, C(1)\, \Psi(0)\kappa = 0 \tag{A2.2.11}$$

where κ = (1 0)'.

It is also instructive to note that the bivariate vector autoregression model (A2.2.1) may also be written in a common trends representation as follows:

$$\begin{bmatrix} w(t) \\ z(t) \end{bmatrix} = \begin{bmatrix} \Psi_{11}(1) & \Psi_{12}(1) \\ \Psi_{21}(1) & \Psi_{22}(1) \end{bmatrix} \begin{bmatrix} \tau_1(t) \\ \tau_2(t) \end{bmatrix} + \begin{bmatrix} \eta_1^*(t) \\ \eta_2^*(t) \end{bmatrix} \tag{A2.2.12}$$

where

$$\begin{bmatrix} \eta_1^*(t) \\ \eta_1^*(t) \end{bmatrix} = \frac{1}{(1-L)} \begin{bmatrix} \Psi_{11}(L)-\Psi_{11}(1) & \Psi_{12}(L)-\Psi_{12}(1) \\ \Psi_{21}(L)-\Psi_{21}(1) & \Psi_{22}(L)-\Psi_{22}(1) \end{bmatrix} \begin{bmatrix} \eta_1(t) \\ \eta_2(t) \end{bmatrix} \tag{A2.2.13}$$

and

$$\tau_i(t) = \tau_i(t-1) + \eta_i(t), \qquad i=1,2 \tag{A2.2.14}$$

Clearly, $[\eta_1^*(t)\ \eta_2^*(t)]' \sim I(0)$, while τ_1(t) and τ_2(t) are pure random walks, so that both w(t) and z(t) are shown to be the sum of two common stochastic trends and an I(0) component.

Then, the Blanchard-Quah restriction (A2.2.4) that η_1(t) has only a temporary effect on w may be written as:

$$\begin{bmatrix} w(t) \\ z(t) \end{bmatrix} = \begin{bmatrix} 0 & \Psi_{12}(1) \\ \Psi_{21}(1) & \Psi_{22}(1) \end{bmatrix} \begin{bmatrix} \tau_1(t) \\ \tau_2(t) \end{bmatrix} + \begin{bmatrix} \eta_1^*(t) \\ \eta_2^*(t) \end{bmatrix} \tag{A2.2.15}$$

that is to say (A2.2.4) is equivalent to imposing the restriction that the stochastic trend in which η_1(t) is the innovation, ie. τ_1(t), is suppressed from the time series representation for w(t).

Appendix 2.3 Seemingly unrelated regressions and generalised least squares

Consider the system of equations:

$$\begin{aligned} y_1 &= X_1\beta_1 + \epsilon_1 \\ y_2 &= X_2\beta_2 + \epsilon_2 \\ &\vdots \\ y_N &= X_N\beta_N + \epsilon_N \end{aligned} \tag{A2.3.1}$$

or, in compact form:

$$y = X\beta + \epsilon \tag{A2.3.2}$$

where the vector of disturbances $\epsilon=[\epsilon_1',\ldots,\epsilon_N']$ satisfies $E(\epsilon)=0$, $E(\epsilon\epsilon')=V$ and $E(\epsilon_{it}\epsilon_{js})=\sigma_{i\,i+1}$ if t=s and 0 otherwise with i and j denoting the i-th and j-th equation in the system - ie. the disturbances are uncorrelated across observations. Defining the total number of observations as T and the number of regressors in each individual regression as K_i (i=1,...,N), so that $K=\Sigma_i K_i$, the disturbance formulation implies:

$$E(\epsilon_i\epsilon_j') = \sigma_{ij}I_T \tag{A2.3.3}$$

or

$$E(\epsilon\epsilon') = V = \begin{bmatrix} \sigma_{11}I & \sigma_{12}I & \cdots & \sigma_{1N}I \\ \sigma_{21}I & \sigma_{22}I & \cdots & \sigma_{2N}I \\ & & \vdots & \\ \sigma_{N1}I & \sigma_{N2}I & \cdots & \sigma_{NN}I \end{bmatrix} \tag{A2.3.4}$$

where I is the identity matrix.

The set of assumptions adopted above implies that each of the regressions in the system considered satisfies the classical assumptions and, therefore, ordinary least squares (OLS) estimation provides consistent, if not efficient, estimates of the parameters. Generalised least squares (GLS) estimation, applied to the system (A2.3.2), provides, however, efficient estimates. Noting that the NxN covariance matrix of the disturbances for the t-th observation is:

$$\Sigma = \begin{bmatrix} \sigma_{11} & \sigma_{12} & \cdots & \sigma_{1N} \\ \sigma_{21} & \sigma_{22} & \cdots & \sigma_{2N} \\ & & \vdots & \\ \sigma_{N1} & \sigma_{N2} & \cdots & \sigma_{NN} \end{bmatrix} \tag{A2.3.5}$$

then, clearly, (A2.3.4) may be rewritten as:

$$V = \Sigma \otimes I \tag{A2.3.6}$$

or

$$V^{-1} = \Sigma^{-1} \otimes I \tag{A2.3.7}$$

Defining the fg-th element of Σ^{-1} as σ^{fg}, the GLS estimator is given by:

$$\hat{\beta} = [X'(\Sigma^{-1}\otimes I)X]^{-1}X'(\Sigma^{-1}\otimes I)\,y$$

$$= \begin{bmatrix} \sigma^{11}X_1'X_1 & \sigma^{12}X_1'X_2 & \cdots & \sigma^{1N}X_1'X_N \\ \sigma^{21}X_2'X_1 & \sigma^{22}X_2'X_2 & \cdots & \sigma^{2N}X_2'X_N \\ & & \vdots & \\ \sigma^{N1}X_N'X_1 & \sigma^{N2}X_N'X_2 & \cdots & \sigma^{NN}X_N'X_N \end{bmatrix}^{-1} \begin{bmatrix} \Sigma_i \sigma^{1i}X_1'y_i \\ \Sigma_i \sigma^{2i}X_2'y_i \\ \vdots \\ \Sigma_i \sigma^{Ni}X_N'y_i \end{bmatrix} \qquad \text{(A2.3.8)}$$

where the inverse on the right hand side is the asymptotic covariance matrix for the GLS estimator. At this stage, clearly, the seemingly unrelated regressions in the system are linked only by their disturbances. As demonstrated by Dwivedi and Srivastava (1978), if *either* the regressions in the system are in fact unrelated and therefore $\sigma_{ij}=0$ *or* the explanatory variables are the same in all equations and therefore $X_i=X_j$, GLS does not provide any efficiency gain relative to OLS. Also, if the explanatory variables in one set of equations are simply a subset of the explanatory variables in another equation, GLS does not provide any efficiency gain in estimating the smaller equations relative to OLS. Given different explanatory variables in the regressions, however, the efficiency gain from using the GLS estimator relative to OLS is greater the greater the correlation of the disturbances and the lower the correlation between the X matrices.

Nevertheless, given that the true covariance matrix of the disturbances is generally unknown, feasible GLS is used in place of GLS. In practice, the elements of Σ can be estimated consistently by OLS as:

$$\hat{\sigma}_{ij} = s_{ij} = (e_i'e_j)/T \qquad \text{(A2.3.9)}$$

where e_i (e_j) denote the OLS residuals from the i-th (j-th) regression, and therefore the estimated Σ

$$S = \begin{bmatrix} S_{11} & S_{12} & \cdots & S_{1N} \\ S_{21} & S_{22} & \cdots & S_{2N} \\ & & \vdots & \\ S_{N1} & S_{N2} & \cdots & S_{NN} \end{bmatrix} \qquad \text{(A2.3.10)}$$

is used in place of the true Σ.

Table 2.1 Savings-investment correlations

Panel A: The Feldstein-Horioka regression

	Estimated β	Estimated correlation of I/Y and S/Y	Chow test for structural break in 1979Q4
1955Q1-1994Q4	0.150 (0.063)	0.185 [2.359]	4.0549 {0.044}
1955Q1-1979Q4	0.587 (0.051)	0.755 [4.875]	
1980Q1-1994Q4	-0.488 (0.091)	-0.574 [-5.292]	

Notes: Standard errors are reported in parentheses, next to the estimated value of β; t-statistics are reported in square brackets, next to the estimated correlation of the savings and investment shares of GDP. P-values are reported in braces, next to the Chow test statistic for a structural break in the regression.

Panel B: The Feldstein-Horioka regression in first difference

	Estimated coefficient	Estimated correlation of Δ(I/Y) and Δ(S/Y)	Chow test for structural break
1955Q1-1994Q4	-0.006 (0.034)	-0.014 [-0.175]	0.3116 {0.577}
1955Q1-1979Q4	-0.288 (0.037)	-0.008 [-0.079]	
1980Q1-1994Q4	-0.105 (0.083)	-0.164 [-1.250]	

Notes: Standard errors are reported in parentheses, next to the estimated value of the coefficient on the change in the savings share of GDP; t-statistics are reported in square brackets, next to the estimated correlation of the changes in savings and investment shares of GDP. P-values are reported in braces, next to the Chow test statistic for a structural break in the regression. Δ denotes the first difference operator.

Table 2.2 Unit root tests: augmented Dickey-Fuller (ADF) test statistics

	ADF test statistic for the sample period 1955-1994	ADF test statistic for the sample period 1955-1979	ADF test statistic for the sample period 1980-1994
log(Y)	-1.2839 (-3.4389)	-1.3043 (-3.4561)	-0.4391 (-3.4849)
I/Y	-1.4727 (-3.4391)	-1.6506 (-3.4561)	-1.4556 (-3.4849)
S/Y	-2.0980 (-3.4394)	-2.2994 (-3.4571)	-1.5515 (-3.4849)
Δlog(Y)	-5.2746 (-2.8799)	-7.6315 (-2.8909)	-5.7025 (-2.9101)
ΔI/Y	-6.0423 (-2.8800)	-8.9258 (-2.8912)	-4.4084 (-2.9101)
ΔS/Y	-11.0446 (-2.8800)	-7.0664 (-2.8912)	-6.0922 (-2.8800)

Notes: The number of lags included is chosen such that the estimated disturbance is approximately white noise. A trend is also included in executing the test statistic, but is dropped if found statistically insignificantly different from zero at conventional nominal levels of significance; for the variables considered, however, the results are qualitatively the same no matter whether a time trend is included or not. Δ denotes the first difference operator. Five percent critical values are reported in parentheses.

Table 2.3 Cointegration tests between I/Y and S/Y

	ADF test statistic on the residuals
Sample 1955Q1-1994Q4	-2.3235 (-3.3766)
Sample 1955Q1-1979Q4	-1.5646 (-3.3995)
Sample 1980Q1-1994Q4	-1.5288 (-3.4687)

Notes: Five percent critical values are reported in parentheses.

Table 2.4 Correlation of components of $\Delta(I/Y)$ and $\Delta(S/Y)$

	Correlation between temporary components	Correlations between permanent components
Sample 1955Q1-1994Q4	0.43 (5.85)	0.29 (3.72)
Sample 1955Q1-1979Q4	0.72 (9.86)	0.54 (6.12)
Sample 1980Q1-1994Q4	0.32 (2.57)	0.18 (1.39)
Test of the equality of the correlations before and after abolition	24.94 [0.000]	14.73 [0.001E-01]

Notes: In parentheses and square brackets we report the t-statistic for the null hypothesis that the correlation is not significantly different from zero and p-values respectively.

Table 2.5 Ratio of GNP to GDP (percentage deviation from group average)

	1964-1968	1969-1972	1973-1976	1977-1980	1981-1984	1985-1988	1989-1992	*1964-1992*
France	0.28	0.10	0.07	0.28	0.32	0.10	-0.3E-3	*0.17*
Germ.	-0.38	-0.12	-0.4E-2	0.28	0.53	1.03	1.26	*0.35*
Italy	0.15	0.24	-0.24	0.07	-0.16	-0.27	-0.76	*-0.13*
Spain	-0.59	-0.80	-0.40	-0.73	-1.06	-0.56	-0.50	*-0.66*
UK	0.45	0.40	0.44	0.28	0.80	0.42	-0.06	*0.39*
Belg.	0.41	0.42	0.31	-0.16	-0.87	-0.86	-0.30	*-0.13*
Neth.	-0.32	-0.26	-0.18	-0.02	0.44	0.13	0.36	*0.01*

Table 2.6 Ratio of total personal income net of transfer payments to regional product (percentage deviation from UK average)

	1972-75	1976-80	1981-85	1986-90	*1972-90*
North	-1.46	-1.74	-1.38	-0.87	*-1.36*
Yorkshire-Humbers.	0.50	0.53	-0.07	1.19	*0.54*
East Midlands	1.21	1.05	0.36	0.84	*0.85*
East Anglia	2.12	2.71	0.40	2.12	*1.82*
South East	2.59	2.41	0.21	2.57	*1.91*
South West	4.10	4.25	3.82	5.19	*4.35*
West Midlands	1.70	1.95	0.41	0.49	*1.11*
North West	-0.51	-1.18	-1.01	-0.63	*-0.85*
Wales	-0.15	0.30	1.62	-3.11	*-0.35*
Scotland	-0.74	-0.92	-1.25	0.81	*-0.52*
Northern Ireland	-0.73	-0.88	-0.66	0.56	*-0.41*
England	1.62	1.50	0.29	1.74	*1.27*

Notes: Transfer payments allowed for include all National Insurance benefits, supplementary benefits, child benefits and war pensions.

Table 2.7 SUR Estimation Results

Aggregate Income from Capital λ^1	Income from Labor $-\beta\lambda^2$	Capital Product $(1-\lambda^1-\lambda^2)$
0.006 (0.008E-01)	-0.201 (0.004)	0.205 (0.007)

Note: The equations were estimated using generalized least squares on data for eight EU countries with the intercept and error variances allowed to vary but with slope coefficients constrained to be equal across countrires. Estimated standard errors for coefficients are reported in parentheses.

Table 2.8 Error Components Estimation Results

Aggregate Income from Capital λ^1	Income from Labor $-\beta\lambda^2$	Capital Product $(1-\lambda^1-\lambda^2)$
0.006 (0.002)	-0.201 (0.006)	0.131 (0.005)

Note: The equations were estimated using generalized least squares and assuming an error components model where, while the slope coefficients were constrained to be equal across countries, error variances were allowed to vary and the intercept terms were assumed to be random variables with constant mean. Estimated standard errors for coefficients are reported in parentheses.

Figure 2.1 Cumulative Impulse Response Functions

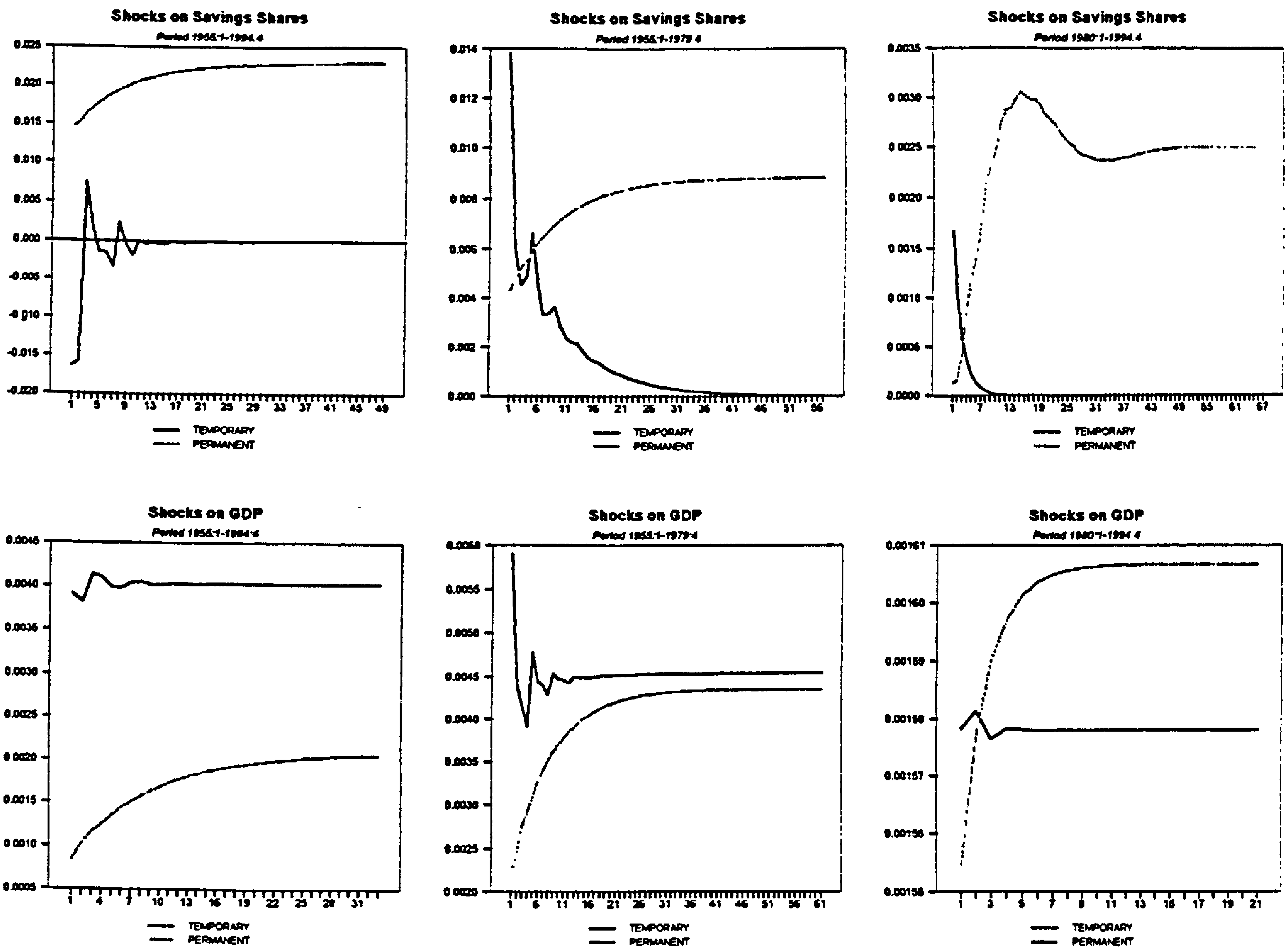
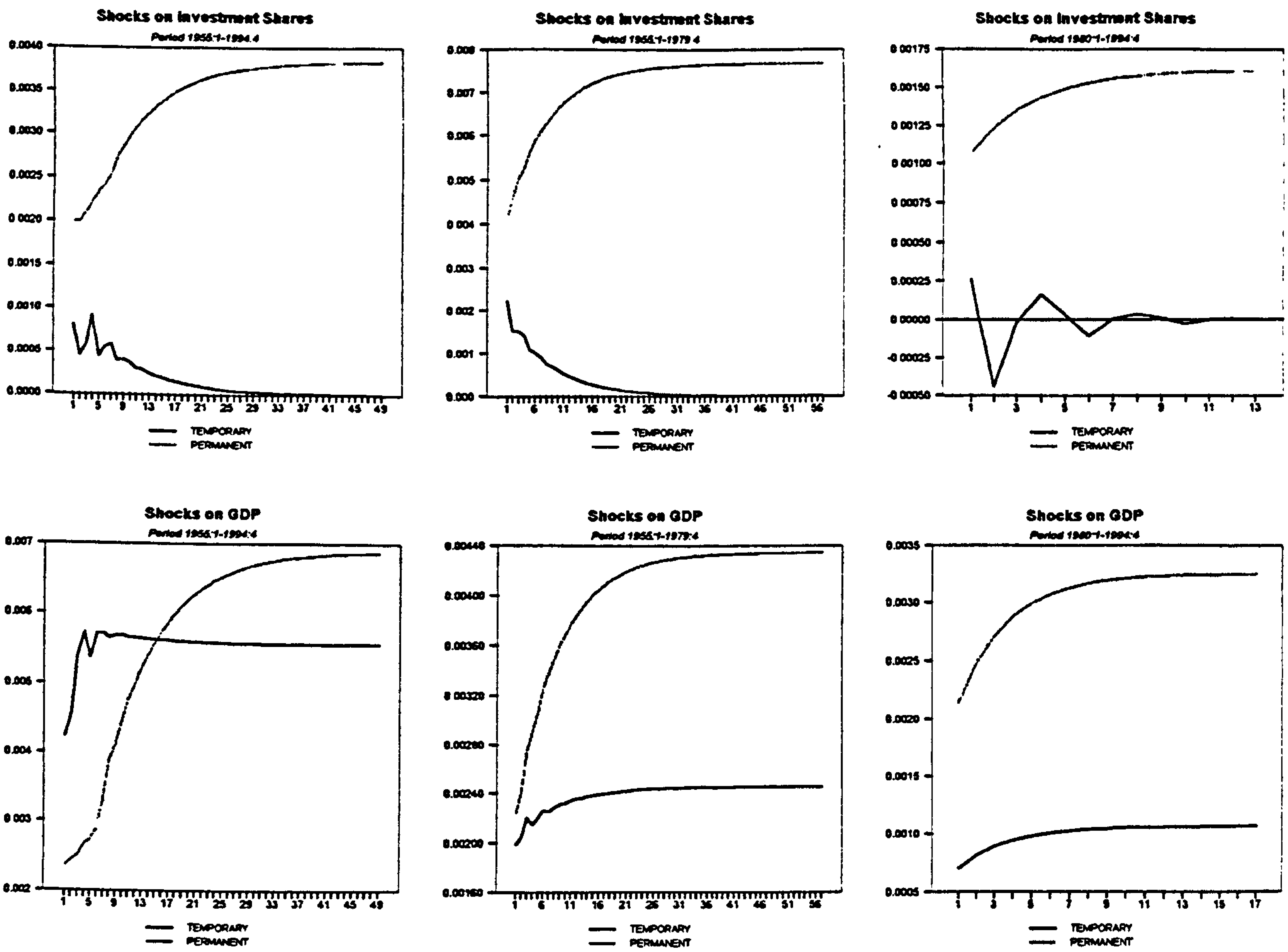


Figure 2.2 Cumulative Impulse Response Functions



CHAPTER 3

The Recent Surge of Capital Flows to Developing Countries

'The surge of capital inflows to Asia and Latin America in the first half of the 1990s is an enormous and multidimensional natural experiment, one that economists will be studying for years to come .... Attempting to accomplish this task is essential to a better understanding of the forces driving the cycles in foreign lending to developing countries, and to gauging the vulnerability of developing countries to a reversal of such capital flows.'

(Calvo, Leiderman and Reinhart, 1996, pp.137-138)

3.1 Introduction

A sharp expansion of net and gross capital flows together with a marked increase in the participation of foreign investors and foreign financial institutions in the financial markets of developing countries is one of the main features of the recent development of international capital flows (World Bank, 1997)[1]. This expansion has been significantly greater than that of international trade flows (Goldstein, Mathieson and Lane, 1991; Montiel, 1993), and has been reinforced by the ongoing process of abolition of various sorts of impediments and capital controls and the broader liberalization of financial markets in developing countries from the late 1980s. In fact, even if the Feldstein-Horioka finding of low international capital mobility based on the correlation between savings and investment shares of GDP (Feldstein and Horioka, 1980) still remains a puzzle, studies based on interest-rate differentials generally provide evidence that a high and increasing degree of international capital mobility already exists not only between the major industrialised countries, but also including the developing countries (Montiel, 1993).

[1] For a comprehensive review of some recent prospects and developments concerning capital flows to developing countries, see *The Road to Financial integration: Private Capital Flows to Developing Countries* (World Bank, 1997), *World Economic Outlook* (IMF, 1994a), or *Trends in Developing Economies* (World Bank, 1995), *International Capital Markets: Developments, Prospects, and policy Issues* (IMF, 1994b, pp. 83-91), *Private Market Financing for Developing Countries* (IMF, 1995), and - for Latin American countries - *Economic Panorama of Latin America, 1994* (UN, 1994).

Another main feature of the recent trend of capital flows to developing countries is that private (bond and equity) as opposed to official capital flows are increasingly a crucial source of financing of large current account imbalances: 'This surge of portfolio investment combined with the large amounts of foreign direct investment has meant that in the early 1990s, close to half of all aggregate external financing of development economies comes from private sources and goes to private destinations' (Bruno, 1993).

These trends in the pattern of net and gross capital flows raise important issues concerning which factors motivate the flows and how they affect and are expected to affect the performance of developing countries. The first part of this chapter focuses on the first of these issues, in an attempt to identify the main determinants of US capital flows to developing countries and to shed light on the relative importance - on the one hand - of the improved economic performances of these countries (country-specific or "pull" factors) and of the stimulus - on the other hand - provided by the decline of US interest rates and the slowdown of the US economy in the early 1990s (global or "push" factors) (Chuhan, Claessens and Mamingi, 1993; Fernandez-Arias, 1996; Agenor, 1997).

The reason why the identification of the relative importance of push and pull factors is important for the effective design of policy - and therefore worthy of investigation - has recently been made forcefully by Fernandez-Arias and Montiel (1996). Fernandez-Arias and Montiel (1996) first summarise a number of arguments describing why large capital flows may, under various circumstances, have adverse effects on developing countries unless proper policies designed to neutralise such effects are adopted. They then point out that, if the causes of capital flows are deemed largely external and exogenous to the developing country in question, then compensatory policies are appropriate. If, on the other hand, the causes of the large capital flows are deemed largely domestic, then a more direct policy design may be more appropriate and effective.

Also, given that much of the recent capital inflow consists of flows labelled short term - primarily bonds, equities and other money market instruments which may be included in the World Bank's definition of portfolio flows - a crucial issue for policy makers is to quantify the degree of persistence characterising capital flows received by developing countries (Claessens, Dooley and Warner, 1995; Corbo and Hernandez, 1993; Gooptu, 1993). Hence, in the second part of this study we address this question in that we measure the relative size and statistical significance of the permanent and temporary components of capital flows to a large group of Latin American and Asian developing countries, on the basis that accounting labels may not in themselves be a reliable guide as to the "coolness" or "hotness" - persistence or temporariness - of capital flows (Claessens, Dooley and Warner, 1995).

The remainder of the chapter is set out as follows. In Section 3.2 we discuss the motivation of international capital flows, briefly review the findings of the relevant literature on the assessment of the determinants of capital flows to developing countries, and develop a simple theoretical model of capital flows. In Section 3.3 we discuss the data set used in this chapter. In Section 3.4 we describe the estimation techniques employed to model capital flows to developing countries in the long-run and in the short-run. In Section 3.5 we report and discuss results from applying cointegration techniques to capital flows to a set of Asian and Latin

American countries in order to identify their long-run determinants. We also discuss in Section 3.5 the results of estimating seemingly unrelated error-correction models for capital flows to those countries in an attempt to understand the short-run dynamics underlying them. In Section 3.6 we briefly discuss the importance of the degree of persistence of capital flows to developing countries from the policy-making point of view. In Section 3.7 we describe the estimation techniques employed in order to model portfolio flows to developing countries and measure their persistence, while in Section 3.8 we report and discuss the empirical results from employing those techniques on data for portfolio flows to a number of Asian and Latin American developing countries. A final section concludes.

3.2 Determinants of international capital flows

Net capital flows arise when savings and investment are unbalanced across countries, and therefore a transfer of real resources is generated through a trade or current account imbalance. *Gross* capital flows, on the other hand, need not involve any transfer of real resources, since they may be offsetting across countries. Nevertheless, they allow individuals and firms to adjust the composition of their financial portfolios and are therefore important in improving the liquidity and diversification of portfolios.

Both net and gross capital flows generally respond to economic fundamentals, official policies and financial market imperfections. International capital flows play an important role in increasing economic efficiency, provided that the international financial markets can correctly evaluate the portfolio preferences of savers, identify and fund investments with a relatively higher expected rate of return, appropriately price financial assets on the basis of their underlying risks and returns and provide information in order to reduce uncertainty.

Following the traditional literature in financial economics, assets are priced, in the absence of distortions, so that the more risky assets offer higher rates of return. Also, while the unsystematic risk can be reduced to a minimum level by appropriately diversifying portfolios, the systematic risk is nondiversifiable and should be reflected in the assets' price, so that investors should be compensated for a relatively higher covariance between the return of an asset and the other assets in the market portfolio (see Taylor, 1991 for a survey of this literature). These arguments therefore imply that among the fundamental determinants of international capital flows one should consider - in the absence of distortions in the international financial markets - factors such as the investment opportunity set available in the global economy, the covariances between the expected returns on various investment projects and the preferences of individuals for present and future consumption as well as their attitudes towards risk.

A major problem, however, exists in measuring empirically the effect of those factors in determining capital flows: international capital markets may react to a shock in one country either through capital flows or through changing the prices of the country's financial claims, or through a mix of the two. Moreover, as the international financial system becomes more integrated and portfolios more diversified, it seems likely that asset price changes may substitute for net capital flows to restore market equilibrium. Therefore, most econometric models try to express financial linkages across countries in terms of interest rate parity conditions. That is to say they specify the asset price linkages that are the outcome of arbitrage between financial markets rather than the capital flows that are part of the arbitrage process (Goldstein, Mathieson and Lane, 1991).

In addition to these economic fundamentals, government policies and capital market imperfections also represent crucial determinants of international capital movements. It is, however, extremely difficult to assess the impact of these policies and distortions on the level and the change of capital flows, since these effects generally overlap one another, creating a mix of impediments and stimuli to capital flows.

Nevertheless, a widely accepted fact is that the processes of deregulation, globalisation and innovation in financial markets have induced a double effect. On one side, financial market efficiency has increased; on the other side, however, volatility in financial markets has also been increased, generating an additional source of risk, not only making the pricing of financial assets more difficult, but also creating potentially more unstable portfolio flows (Corrigan, 1989; Claessens, Dooley and Warner, 1995; Grabel, 1995; Clarke, 1996)[2].

3.2.1 Push and pull factors

The recent literature usually distinguishes between two different sets of factors affecting capital movements (eg. Claessens, Dooley and Warner, 1995; Chuhan, Claessens and Mamingi, 1993; Fernandez-Arias, 1996; Fernandez-Arias and Montiel, 1996; Agenor, 1997). First, country-specific (pull) factors reflecting

[2] Nevertheless, some evidence exists which suggests that volatility is not correlated at all with any measure of financial integration and that no significant increase in volatility arises because of financial liberalisation (eg. Tesar and Werner, 1995; Bekaert, 1995).

domestic opportunity and risk are potential determinants of international capital flows. As developing countries' creditworthiness is restored, capital (bond and equity) flows are likely to become an increasingly prominent source of external finance. For example, equity-related capital flows may potentially be very large and in the form of either foreign direct investment or portfolio investment in equities. Foreign direct investment may for example be attracted by the opportunity to utilise local raw materials or employ a local labour force. Portfolio equity flows to developing countries, although they have increased sharply in recent years, are expected to be extremely sensitive to the degree of openness of the country considered, and in particular to the rules concerning the repatriation of capital and income (Williamson, 1993). The right to repatriate dividends and capital is perhaps the most important factor in attracting significant amounts of foreign equity flows (Goldstein, Mathieson and Lane, 1991)[3].

Rates of return, an obviously crucial determinant of capital flows, are often very high in developing countries' financial markets in comparison with many major markets in the industrial countries, also reflecting the high risk stemming from the usually high volatility of those markets. In particular, rates of return significantly rose in developing countries in the late 1980s relative to those available in the major industrialised economies (Calvo, Leiderman and Reinhart, 1993; Chuhan, Claessens and Mamingi, 1993).

[3] The International Finance Corporation (IFC) differentiates between countries which give foreign investors free and unrestricted repatriation of capital and income from shares and countries, defined "relatively open", which apply some restrictions on the repatriation of capital and income, and some other countries, defined "relatively closed", which still apply very strict restrictions to the way in which capital can be repatriated.

Credit ratings and secondary market prices of sovereign debt, reflecting the opportunities and risks of investing in the country, are also likely to be important factors in determining capital flows (Bekaert, 1995). Those indicators also experienced a rising trend in the late 1980s (Mathieson and Rojas-Suarez, 1992; Reisen and Fischer, 1993; Chuhan, Claessens and Mamingi, 1993).

The second set of determinants of capital flows to developing countries identified by the literature are global (push) factors. For example the sharp increase in US capital flows, which represent a significant share of the portfolio flows received by emerging markets, may also have been induced to some extent by the fast and marked fall in US interest rates (short, medium and long-term) which occurred in the late 1980s. Moreover, the slowdown of the US economy may also have been a significant explanatory variable in attracting US capital flows in the late 1980s, especially because in that period macroeconomic policies, labour market conditions and exchange rate policies in many developing countries were starting to appear noticeably more stable (Calvo, Leiderman and Reinhardt, 1993, 1996). One would expect that, as the governments of the developing countries establish macroeconomic and institutional reforms, international investors gain confidence and may be more willing to direct capital flows towards the new markets (Papaioannu and Duke, 1993).

In an analysis of the motivations of portfolio flows to developing countries, Chuhan, Claessens and Mamingi (1993) find - using a panel data approach for the period 1988-1992 - that a sample of Latin American and Asian countries are about equally sensitive to global factors and country-specific factors. They also find that equity flows, relative to bond flows, are more responsive to global factors; bond flows, however, are more responsive to a country's credit rating and to the secondary market price of debt.

Using a model with partial irreversibility of investment, Daveri (1995) derives a negative relationship between costs of entry and exit from financial markets and foreign investment, generating a theoretical framework which is consistent with the results of Chuhan, Claessens and Mamingi (1993).

3.2.2 A simple theoretical framework

A useful analytical framework which incorporates the effect of domestic and global factors on capital flows is due to Fernandez-Arias and Montiel (1996)[4]. In particular, they separate potential domestic causes into those that operate at the project level and those which operate at the country level. Assuming capital flows may occur as transactions in n different types of assets, indexed by s ($s=1,\ldots,n$), the domestic return on an asset of type s is decomposed into two components: a project expected return (G_s) and an adjustment factor depending upon the creditworthiness of the country (C_s). The project return is assumed to be a function of a vector of net flows (F) going to projects of all types, while the creditworthiness factor is assumed to be a function of the vector of the end-of-period stocks of liabilities of all types, S: $S=S_{-1}+F$, where S_{-1} denotes the initial stocks of liabilities. Given that external creditors will diversify their portfolios, the opportunity cost of assets of type s, V_s is a function of S. Fernandez-Arias and Montiel (1996) then write down an arbitrage condition - from which F may be

[4] An alternative way to model capital flows would be to apply the international capital asset pricing model (Bohn and Tesar, 1996).

solved for - of the form:

$$G_s(g, F)\, C_s(c, S_{-1}+F) = V_s(v, S_{-1}+F) \tag{3.1}$$

where g, c and v represent shift factors associated, respectively, with the domestic economic environment and domestic creditworthiness (pull factors) and the financial conditions of the creditor country (push factors). G_s, C_s and V_s are assumed to be increasing functions of g, c and v respectively. The equilibrium or 'desired' value of the vector of net flows **F**, F^* say, determined implicitly by equation (3.1), may be expressed as:

$$F^* = F^*(g, c, v, S_{-1}) \tag{3.2}$$

where F^* will be increasing in g and c but decreasing in v and S_{-1}. Holding S_{-1} constant, totally differentiating equation (3.2) and approximating total derivatives by first differences yields:

$$\Delta F^* = F_1^* \Delta g + F_2^* \Delta c + F_3^* \Delta v \tag{3.3}$$

where subscripts denote partial derivatives. Equation (3.3) describes the pattern of changes in desired capital flows, determined by a combination of the changes in the pull factors g and c and the push factors v and the initial value of S. Increases in g and c and decreases in v may therefore induce a prolonged surge in capital flows to developing countries. This simple model is clearly consistent with both the push and the pull view of the surge in capital inflows, while the relative importance of the two factors will depend upon the relative magnitudes of the partial derivatives of F^* as well as on the relative magnitudes of changes in the factors themselves.

Differences in short-run and long-run capital movements might arise according to equation (3.3) by considering the decomposition of changes in g, c and v into permanent and transitory components: permanent changes in g, c and v may cause long-run, permanent changes in the net flows pattern, whereas transitory changes in these factors may be expected to generate transitory, short-term changes in net flows, which may be reversed over time. For example, the gradual permanent removal of capital controls and liberalization of restrictions on foreign direct investment may be expected to reduce the adjustment costs faced by foreign investors in diversifying their portfolio and cause a gradual stock adjustment (flow) over time. This gradual adjustment also implies a complex dynamic pattern of net flows towards their long-run equilibrium value and is consistent with an estimation methodology which distinguishes the short-run determinants from the long-run determinants of capital flows.

Then, dynamic adjustment can be formally introduced into the Fernandez-Arias - Montiel framework by assuming a simple costs-of-adjustment model in which factors such as market imperfections, informational asymmetry (Stiglitz and Weiss, 1981), entry and exit costs to emerging financial markets (Daveri, 1995) and other factors are encapsulated in the assumption that creditors face costs in adjusting their portfolio, which we assume are increasing in the size of the adjustment. The desired vector of capital flows is given by equation (3.2): assume that agents wish to minimise the distance between desired and actual flows, subject to costs of adjustment. A simple way of modelling this is to assume a simple quadratic loss function for investors of the form:

$$\pounds = (F-F^{*})'M_1(F-F^{*}) + (F-F_{-1})'M_2(F-F_{-1}) \qquad (3.4)$$

where M_1 and M_2 represent positive definite weighting matrices. From the first-order conditions for minimisation of £ we can derive a simple equation for changes in F:[5]

$$\Delta F = (M_1+M_2)^{-1}M_1(F^{*}-F_{-1}) \qquad (3.5)$$

which, rearranging and using (3.3), can be equivalently expressed in the error-correction form:

$$\Delta F = A_0(F^{*}-F)_{-1} + A_1\Delta g + A_2\Delta c + A_3\Delta v \qquad (3.6)$$

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[5] In order to derive equation (3.5), note that the first-order condition of the loss function (3.4) implies:

$$2M_1(F - F^{*}) + 2M_2(F - F_{-1}) = 0$$

Subtracting M_1F_{-1} on both sides of the equation yields:

$$(M_1 + M_2)F - M_1F_{-1} - M_2F_{-1} = M_1F^{*} - M_1F_{-1}$$

from which equation (3.5) follows directly.

where $A_0=(M_1+M_2)^{-1}M_1$ and $A_i=(M_1+M_2)^{-1}M_1F_i^*$ $(i=1,2,3)$.

Intuitively, according to (3.6), changes in current capital flows are determined partly by the distance between desired and actual capital flows in the previous period and partly by changes in the factors determining the desired level of capital flows. Again, changes in the push and pull factors might be decomposed into permanent and transitory components, with only the former affecting the long-run level of **F** (by causing persistent changes in **F**) while transitory movements, which are reversed over time, will entail transitory movements in **F** which are also reversed over time. For example, a temporary reduction in US interest rates, which might be interpreted as a downward movement in v, will, other things equal, generate an upward movement in the level of capital flows to the developing country equal to $A_3\Delta v$ (which is positive since F^* is decreasing in v and Δv is negative). If this change persists over time, then the long-run level of **F** will be raised because of the permanent effect on F^* operating through (3.2). If on the other hand, the change in v is reversed over subsequent periods, then although ΔF will be affected over several periods, the net long-run effect on both desired and actual capital flows will be zero.

While the simple theoretical framework developed in this section obviously should not be taken too literally, it does suggest, together with a reading of the literature also discussed in this section, that shifts in capital flows may be determined by both push and pull factors and by factors which are permanent as well as by those which are transitory. However, the issue as to which of these factors - push or pull, short-run or long-run - is relatively more important is hard to determine theoretically and therefore remains largely an empirical matter.

3.3 Data

The data set used in the present study is identical to that employed by Chuhan, Claessens and Mamingi (1993). We use monthly data on US portfolio flows, defined as gross and net purchases of foreign long-term securities for a group of nine Latin American countries and nine Asian countries[6]. The capital flow data are taken from the *International Capital Reports* of the US Treasury Department and, according to the computations of Chuhan, Claessens and Mamingi (1993), cover a very substantial amount of the US portfolio flows to developing countries[7]. Also, following Chuhan, Claessens and Mamingi (1993), we use net equity flows (ef) and gross bond flows (bf) to developing countries, which cover a substantial share of portfolio flows to those countries. Even if in principle we are concerned with modelling net capital flows, it seems preferable to use gross measures for bond flows in order to abstract from the effect of sterilization policy actions and other types of reserve operations by central banks[8].

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[6] The Latin American countries are Argentina, Brazil, Chile, Colombia, Ecuador, Jamaica, Mexico, Uruguay and Venezuela. The Asian countries examined are China, India, Indonesia, Korea, Malaysia, Pakistan, Philippines, Taiwan and Thailand. For a full statistical description of the data set employed in this study, see Chuhan, Claessens and Mamingi (1993).

[7] Note, however, that the US share in total portfolio flows is far larger for the Latin American countries relative to the Asian countries.

[8] These data are collected by the Treasury from financial intermediaries in the US through the *International Capital Form S* reports. Hence, the data do not include direct dealings of US investors with foreign intermediaries as these transactions bypass the system. Note also that the data on bonds cover transactions of foreign securities in the US from and to developing countries; transactions in bonds not issued by the developing country concerned nor by US parties - which would be covered in the data set build in this fashion - is expected to be quite insignificant (see Chuhan Claessens and Mamingi, 1993, pp. 6-7).

The explanatory variables used consist of two different sets: a set of country-specific factors and a set of global factors. As country-specific factors we use the country credit rating (cr) and the black market exchange rate premium (bm) which are available for both groups of countries considered. The credit rating variable is constructed on the basis of the *Institutional Investor*'s semi-annual country credit rating, while the black market exchange rate premia are calculated from data taken from the *World Currency Yearbook* and the official exchange rates of the International Monetary Fund's *International Financial Statistics* (IFS).

Finally, we also take from the IFS data base a short-term and a long-term US nominal interest rate - the Treasury Bill Rate (i) and the Government Bond Yield (r) respectively - and the level of real US industrial production (y), which we consider as the global factors potentially determining US capital flows to developing countries[9].

The sample period examined runs from January 1988 to September 1992, corresponding exactly to the sample period considered by Chuhan, Claessens and Mamingi (1993)[10] [11].

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[9] Another possible candidate as a country-specific variable is secondary market debt prices, which is a continuous random variable likely to be I(1). Chuhan, Claessens and Mamingi (1993, pp. 10-11) use secondary market prices only for the countries for which *Salomon Brothers* provide data. Since secondary market prices are not available for all the countries we consider in this study, for consistency, we do not use secondary market prices in this study.

[10] An immediate avenue for research is to extend the sample period employed in this study. Some of the global factors considered here as well as in Chuhan, Claessens and Mamingi (1993) have changed over the last two years or so; in particular, the US interest rates have started to rise after a prolonged period of decline and the US economy is in recovery after the slow-down in the early 1990s. While structural breaks may have been caused by these changes, therefore generating additional estimation difficulties, it will be interesting to investigate

3.4 Estimation strategy

Consider a panel of N countries, indexed by i (i=1,...,N), with portfolio flows at time t denoted f_{it}, assumed to be integrated processes of order one, I(1). Also define a vector of country-specific factors as $\mathbf{x}_{it}$ and a vector of global factors as $\mathbf{w}_{it}$, and assume that both $\mathbf{x}_{it}$ and $\mathbf{w}_{it}$ contain at least one I(1) variable and no higher-order integrated processes. Then we can analyse the long-run behaviour of portfolio flows by investigating cointegrating relationships of the kind:

$$f_{it} = \boldsymbol{\beta}'\boldsymbol{x}_{it} + \boldsymbol{\gamma}'\boldsymbol{w}_{it} + e_{it} \qquad i = 1, \ldots, N \tag{3.7}$$

where f_{it} may be either equity flows - say ef_{it} - or bond flows - say bf_{it}.

If cointegration is established in equations (3.7) - ie. the error term e_{it} is approximately stationary - then the I(1) variables in $\mathbf{x}_{it}$ and $\mathbf{w}_{it}$ may be considered to capture the long-run or permanent components in f_{it}, while e_{it} captures the short-run or temporary movements. Given that the dependent variable is I(1), then there must be at least one I(1) variable among the explanatory variables; if all of the explanatory variables are I(0), then equation (7) will be mis-specified (Pagan and

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whether and how portfolio flows to developing countries have been affected by these changes.

[11] Chuhan, Claessens and Mamingi (1993, pp.13-15) provide a detailed description of the data employed here in Section 3 of their paper, and also report illustrative graphs of all the data series. In addition, they report a panel correlations table and find that most correlations are of the expected sign: negative between US interest rates and flows, negative between black market exchange rate premia and flows as well as between US industrial production and flows, positive between credit ratings and flows. These preliminary results are in fact encouraging in the sense that they give strong support to the underlying rationale of empirical models of the type employed in the present study.

Wickens, 1989, p. 1002; Banerjee, Dolado, Galbraith and Hendry, 1993; Baffes, 1997). In fact, after we execute preliminary unit root tests on the series in question in order to identify their order of integration, the first exercise we address is testing for non-stationarity of the residuals of the cointegrating regressions described by (3.7) using the two-step procedure due to Engle and Granger (1987).

In order to ensure that (3.7) is a cointegrating relationship we also employ the relatively more powerful technique described by Johansen (1988) in a vector autoregression (VAR) context (see Kremers, Ericsson and Dolado, 1992). The Johansen cointegration technique is based on an error correction representation of the p-th order vector autoregression (VAR(p)) model with gaussian errors of the form:

$$\Delta q_{it} = \Phi_i + \Gamma_{i1}\Delta q_{it-1} + \Gamma_{i2}\Delta q_{it-2} + \ldots + \Gamma_{ip-1}\Delta q_{it-p+1} + \Pi q_{it-p} + B_i z_{it} + u_{it} \tag{3.8}$$

where $\mathbf{q}_{it}$ is an mx1 vector of I(1) variables, ϕ_i is an mx1 vector of constants, $\mathbf{z}_{it}$ is an sx1 vector of I(0) variables, the $\mathbf{\Gamma}_{ij}$'s and $\mathbf{\Pi}$ are mxm matrices of parameters, $\mathbf{B}_i$ is an mxs matrix and $\mathbf{u}_{it} \sim N(0,\Sigma)$. The Johansen maximum likelihood procedure is based on the estimation of (3.8) subject to the hypothesis that the matrix $\mathbf{\Pi}$ has a reduced rank $r<m$, which may be written formally as H(r): $\mathbf{\Pi}=\alpha\delta'$, where α and δ are mxr matrices. $\delta'\mathbf{x}_t$ represents the cointegrating vectors. In the Johansen cointegration framework, we can also test the hypothesis that the estimated coefficients on the country-specific factors or the estimated coefficients on the global factors are zero. These tests may in fact shed light on the relative

importance of the two sets of factors employed as explanatory variables, and therefore indicate whether pull (country-specific) factors or push (global) factors are the major determinants of longer-run capital flow movements[12].

Since e_{it} may alternatively be interpreted as the deviation from the long-run equilibrium ($e_{it}=f_{it}-\beta' x_{it}-\gamma' w_{it}$), it may be used in an analysis of the short-run dynamics of capital flows through estimation of the error correction model (ECM):

$$\Delta f_{i,t} = \psi_i - \rho_i (f - \beta' x - \gamma' w)_{i,t-1} + \theta_i \Delta f_{i,t-1} + \sum_{j=0}^{p} \lambda'_{ij} \Delta x_{i,t-j} + \sum_{j=0}^{p} \delta'_{ij} \Delta w_{i,t-j} + \omega_{i,t} \tag{3.9}$$

where i (i=1,...,N) is a country index, ψ_i is a constant term, j (=0,...,p) denotes the number of lags and $\omega_{i,t}$ is approximately white noise. Equations (3.9) are a panel data generalisation of the error-correction representation of cointegrated variables established by Engle and Granger (1987) and follow directly from the Granger representation theorem (Granger, 1983; Engle and Granger, 1987). They may be interpreted as a statistical approximation to the theoretical error-correction form (3.6) derived above, on the assumption that desired capital flows $\mathbf{F}^*$ are achieved in the long run and are well approximated by the cointegrating vector.

The adoption of a system of seemingly unrelated error correction models, employed in other dynamic modelling contexts in international finance, allows us to shed some light on the dynamics versus long-run determination of bond and

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[12] A more detailed explanation of the Johansen cointegration procedure is given in Chapter 4.

equity flows and differentiates the present study from the earlier studies on portfolio flows determination.

Equations (3.9) provide the full dynamic interaction of the determinants of capital flows. In particular, the analysis of the parameters ρ_i may shed some light on the degree of persistence in the flow series considered (Claessens, Dooley and Warner, 1995; Dooley, Fernandez-Arias and Kletzer, 1996), conditional on the dynamics of the underlying fundamental determinants.

Overall, therefore, if we assume that actual and desired capital flows coincide in the long-run, then we can think of the cointegrating relationship as determining the desired level of capital flows: $f^*_{it}=\beta' x_{it}+\gamma' w_{it}$. Under this interpretation, (3.7) is the empirical analogue of the theoretical equation (3.2). Hence, as the error term e_{it} can be thought of as the difference between desired and actual flows, $e_{it}=f_{it}-f^*_{it}$, we can then estimate an error correction equation which is the empirical counterpart of the theoretical error correction model (3.6), in which changes in flows are a function of changes in the variables determining the desired level of flows - ie. changes in x_{it} and w_{it} - as well as of the error correction term itself, e_{it}. Estimating the dynamic error correction form then allows us to determine which are the important factors determining short-run movements in capital flows.

Note also that, in principle, the country-specific factors could be influenced by the global factors; for example, global interest rates may affect the pattern of secondary market debt prices and credit ratings (see eg. Dooley, Fernandez-Arias and Kletzer, 1996; Fernandez-Arias, 1996). To the extent that country-specific factors affect portfolio flows only insofar as they are collinear with the global factors, this would show up as the country-specific factors being insignificant when both sets of explanatory variables are included.

3.5 Empirical results

3.5.1 Long-run fundamental determinants of portfolio flows to developing countries

As a preliminary step to testing for cointegration in (3.7), we execute augmented Dickey-Fuller (ADF) unit root test statistics on the series used. The results, reported in Table 3.1, suggest that all of the series are realizations from integrated processes of order one. The null hypothesis of nonstationarity is never rejected for both portfolio flows (equities and bonds) and global factors, while it is rejected once (Chile) for the credit rating and five times out of eighteen (Malaysia, Philippines, Taiwan, Argentina and Uruguay) for the black market exchange rate premium. In general, the ADF test statistics are relatively closer to the rejection region for credit ratings and black market premia, perhaps indicating - as one would expect - less evidence of nonstationarity in these series. In fact, credit ratings are discrete random variables that cannot assume a very large number of outcomes; also, black market premia tend to widen ahead of crises but may become very thin over time as a result, for example, of international capital market integration. Overall, however, the unit root tests executed suggest that a permanent component is statistically significant in credit ratings and *generally* statistically significant also in black market premia, at the five percent nominal significance level, therefore implying that these variables can potentially contribute to the long-run determination of portfolio flows to the countries concerned.

Given the finding that the series considered appear to be realizations of I(1) processes, we are justified in testing for cointegration in equation (3.7). In Table

3.2 (Panel A to D), we report the ADF test statistics on the residuals of the cointegrating relationship (3.7) for all the countries examined. We test for non-stationarity of the residuals in (3.7) considering first equity flows and then bond flows as dependant variables. Apart from 6 of the 36 regressions examined, we are always able to reject the null hypothesis of no cointegration at the five percent nominal level of significance, which suggests that a combination of domestic factors and global factors share a common trend with capital (equity and bond) flows. In order to gauge which of the two sets of variables, $\mathbf{x}_{it}$ and $\mathbf{w}_{it}$, are more important in determining long-run movements of capital flows, we employ the ADF test statistic on a regression including only country-specific factors ($\mathbf{x}_{it}$) and then only global factors ($\mathbf{w}_{it}$). In fact, a test of the null hypothesis H_0: $\beta=0$ or H_0: $\gamma=0$ cannot be executed in such a framework because of the strong bias characterising the estimated standard errors. The results, reported in Table 3.2, show that cointegration is often established in both regressions. Nevertheless, in 3 cases out of 6 for which the ADF fails to reject the null hypothesis in regressions of the type (3.7), the results suggest that if the null of no-cointegration cannot be rejected in a regression of capital flows on $\mathbf{x}_{it}$ ($\mathbf{w}_{it}$), we are not able to reject non-cointegration either in a regression of capital flows on $\mathbf{w}_{it}$ ($\mathbf{x}_{it}$), perhaps supporting the view that it is appropriate to consider both sets of variables as explanatory variables of capital flows.

It must be noted, however, that the results in Table 3.2 are subject to the problem of small sample bias in cointegrating relationships originally highlighted by Banerjee, Dolado, Hendry and Smith (1986) and by Stock (1987). Therefore, in order to strengthen the finding of cointegration in equations (3.7), we employ a relatively more powerful cointegration technique, the Johansen (1988) procedure in a vector autoregressive system including capital flows, country-specific as well as global factors (Kremers, Ericsson and Dolado, 1992). Moreover, in this framework we can also test - using likelihood ratio tests - for zero restrictions on the coefficients on the specific (pull) factors or on the coefficients on the global (push) factors. In order to save space and given that the results for all regressions are qualitatively the same, in Table 3.3 we report the results from the execution of the Johansen procedure for the six regressions on which cointegration could not be established using the ADF test statistics on the residuals[13]. The results enable us, on the basis of both the test statistics suggested by Johansen, to establish cointegration in *all* 36 systems analysed at conventional nominal levels of significance, therefore suggesting that the cases in which the Engle and Granger procedure fails to detect cointegration may simply be due to the low power of the test[14]. The Johansen procedure also suggests that multiple cointegrating vectors exist between the variables examined, as may be the case whenever cointegration is investigated between more than two variables. Further, the zero restrictions on either the coefficients on country-specific or on global factors are *always* strongly rejected, suggesting the long-run importance of both pull and push factors. This result also suggests that pull factors are important determinants of capital flows in their own right, and not only because they may be correlated with the push factors

[13] Clearly, the results for the remaining regressions are available from the author on request.

[14] In particular, the evidence of cointegration is impressive since it is based on the estimation of individual regressions with a relatively small number of observations. This obviates the need to employ more powerful panel cointegration tests (Quah, 1994; Im, Pesaran and Shin, 1997).

(see Section 3.4.2 above).

3.5.2 A dynamic analysis of portfolio flows to developing countries

The finding that at least one cointegrating relationship exists between a set of variables implies that an error-correction model exists, since - as established by the Granger representation theorem - for any set of I(1) variables, error correction and cointegration are equivalent representations. Therefore, the residuals from the equilibrium regressions (3.7) can be used to estimate by generalised least squares (GLS) the error-correction models (3.9) as seemingly unrelated regressions. The speed of adjustment coefficients ρ_i are of particular interest in that they have important implications for the dynamics of the model. In fact, for any given value of the deviations from long-run equilibrium $(f_{it}-\beta' x_{it}-\gamma' w_{it})$, a large value of the speed of adjustment ρ_i is associated with a large value of the change in capital flows, Δf_{it}. If ρ_i is zero, the change in capital flows does not respond at all to the deviation from long-run equilibrium in period (t-1). If ρ_i is zero and all $\lambda_{ij}=0$ $(\delta_{ij}=0)$, then it can be established that $\Delta \mathbf{x}_{it}$ $(\Delta \mathbf{w}_{it})$ does not Granger cause Δf_{it}. We know, however, that one or both of these coefficients should in fact be significantly different from zero since the variables are found to be cointegrated (Engle and Granger, 1987).

We examine the panel data generalization of the error-correction representation of cointegrated variables (3.9) using the cointegrating residuals retrieved from the cointegrating regressions (3.7). Note that the estimation, executed by GLS, of seemingly unrelated ECMs exploits the nature of the panel data set employed and is an efficient technique for analyzing the effect of a common set of global factors across a group of countries in a dynamic framework. In fact, the disturbances in the individual regressions are expected to be very highly correlated since they certainly include some factors that are common to all of the countries as well as some factors that are specific to a particular country (Dwivedi and Srivastava, 1978).[15]

The results from the GLS estimation of the seemingly unrelated error-correction models for both Asian and Latin American portfolio flows are summarised in Tables 3.4 and 3.5. We adopt the conventional general-to-specific procedure to estimate a parsimonious error-correction model, as suggested by Davidson, Hendry, Srba and Yeo (1978) and Hendry (1983)[16]. The resulting models appeared to be quite adequate in terms of high coefficients of determination and approximately white noise residuals. While space considerations preclude us from reporting each of the estimated equations in detail, the following equation, estimated for bond flows to Colombia, is reasonably representative:

$$\Delta bf_t = \underset{(0.097)}{-0.976}\, e_{t-1} \underset{(0.083)}{+0.197}\, \Delta cr_t \underset{(0.083)}{+0.555}\, \Delta cr_{t-1} \underset{(0.687)}{-1.811}\, \Delta i_{t-1}$$

$$R^2 = 0.79, \quad Q(24) = \underset{[0.917]}{15.138}$$

where R^2 denotes the coefficient of determination and Q(24) denotes the Ljung-Box

[15] See Appendix 2.3

[16] See Cuthbertson, Hall and Taylor (1992) for a text book exposition of general-to-specific modelling.

statistic for residual autocorrelation computed for 24 autocorrelations; figures in parentheses are estimated standard errors and the figure in brackets is the marginal significance level (a constant term was also included). bf denotes the bond flow, cr the country credit rating and i the US short-term interest rate; e denotes the cointegrating residual or error correction variable. The equation shows a strongly significant and relatively large error correction coefficient - indicating rapid adjustment - and demonstrates the importance of both push and pull factors[17].

In Table 3.4, we report the number of statistically significant dependent variables in the estimated seemingly unrelated ECMs, namely the error correction term (e), a once-lagged dependent variable (ldv), the credit rating (cr), the black market exchange rate premium (bm), short- and long-term US interest rates (i and r respectively) and US real industrial production (y). As Table 3.4 shows in detail and Table 3.5 summarises, the underlying dynamics of the error-correction models for Asian capital flows are - especially for equities - slightly more complex than the dynamics of the error-correction models for Latin American capital flows, in the sense that more variables are found to be statistically significant on the right hand side of the ECMs. Global and country-specific factors seem to have roughly the same statistical significance in determining the change in equity flows for both Asian and Latin American countries. The change in bond flows, however, appears to be relatively more strongly determined by global factors than by domestic ones. As shown in Table 3.5, the number of times a global factor is significant in the error correction models for bonds is much greater than the number of significant

[17] The full set of results with the point estimates of the parameters of the seemingly unrelated ECMs is available from the author.

country-specific variables, for both Asia (24 versus 25 and 24 versus 16 for equities and bonds respectively) and - to an even larger extent - Latin America (21 versus 17 and 26 versus 13 for equities and bonds respectively). This suggests that equity flows are much more responsive to changes in country-specific factors than are bond flows.

Even if some significant lags of country-specific factors enter the error-correction model, overall bond flows appear to react predominantly to changes in global factors. From Table 3.5, we can see that among global factors, the effect of a change in US interest rates appears to be more strongly significant in explaining the dynamics of bond flows relative to the other global factor considered in this study, the growth of US industrial production[18]. The finding that the dynamics of bond flows, relative to equity flow dynamics, are determined more by global factors than by country-specific factors contrasts with the finding of Chuhan, Claessens and Mamingi (1993), who find that equity flows are more sensitive than bond flows to global factors, while bond flows are more sensitive to country-specific factors. In Chuhan, Claessens and Mamingi (1993), however, the primary interest is the investigation of the long-term determinants of the large capital flows to developing countries using a panel data approach rather than fully modelling the dynamics of capital flows. Hence, their conclusions on the relative responsiveness of bond and equity flows to the two sets of factors is drawn for illustrative purposes using a simpler approach based on the computation of standardised coefficients and elasticities.

[18] It must be pointed out, however, that the two interest rates series considered are not expected to have necessarily the same cyclical properties.

The relative importance of the global factors suggested by our results is consistent with previous work by Calvo, Leiderman and Reinhart (1993, 1996) who first suggested the importance of US interest rates and of the slowdown in US industrial production over the period 1988-1992 in explaining portfolio flows to emerging markets, arguing also that a reversal of the global conditions could induce a fast outflow of capitals from developing countries. In the light of our results, one may perhaps be more specific in pointing out that interest rates are likely to be the most important determinant of the dynamics of portfolio flows (especially bonds) to developing countries (both Asian and Latin American).

Finally, note that the importance of interest rates as a short-term determinant of portfolio flows is higher for Latin American countries than for Asian countries. This can be seen by noting, from Table 3.5, that interest rates are statistically significant 31 times (34 percent of the total number of significant terms) in the Asian ECMs, while they are significant 35 times (45 percent of the total number of significant terms) in the Latin American ECMs. Put another way, Latin American inflows are as sensitive as Asian inflows to interest rates, but the former appear to be less sensitive to all the other (global and country-specific) factors.

3.6 The "hotness" of portfolio flows to developing countries

The large increase in capital flows to developing countries, mainly consisting of portfolio flows and coming from private institutions, also raises the question of whether these flows are sustainable since they mainly involve the purchase of assets with short maturities rather than fixed assets (Claessens, Dooley and Warner, 1995). Reisen (1993) provides evidence suggesting that most flows to Latin

America are very "hot" - ie. have very low persistence - while Nunnenkamp (1993) argues that the degree of "hotness" of capital flows differs widely across developing countries.

The reason the identification of the hotness of capital flows is considered an important issue stems from the argument that different policy measures have in principle to be undertaken by policy makers in response to flows with different persistence properties. Kiguel and Caprio (1993) and Corbo and Hernandez (1993) describe a number of cases of policy responses adopted by developing countries in response to the increase of "hot" money flows; these include exchange rate policies, fiscal policy, sterilization policies or changes in reserve requirements.

Claessens, Dooley and Warner (1995), using balance of payments capital account data for a range of developing countries, show that the accounting labels "short-term" and "long-term" as traditionally applied to capital flows do not provide any reliable indication as to the degree of persistence or "coolness" of the flows. They argue, moreover, that relying on accounting labels rather than on the actual properties of capital flows for policy purposes may generate disastrous results. Dooley (1995), for example, argues that the 1982 debt crisis was in part generated by the imprudence of private investors induced by capital flows to developing countries in the 1970s which, although labeled as private, should really have been considered as official in that a general governmental guarantee underlay them. More generally, it is clear that, whatever accounting label may be appropriate, if capital inflows display high volatility or have a large temporary component, their reversibility can potentially generate high adjustment costs arising from resource reallocation, sunk costs and other hysteresis effects or other market imperfections.

The implication, therefore, is that the country concerned should try to avoid a long and difficult adjustment process on the basis of capital flows which may be reversed later [19].

In the present study, we also contribute to the literature discussed above in that we try to shed light on the "hotness" of capital flows to developing countries by modelling portfolio flows in a fashion which allows us to quantify the degree of persistence of the flows themselves, on the basis that accounting labels alone may not be very informative in this respect.

3.7 Measuring the persistence of capital flows

3.7.1 Unobserved components

In this study, the persistence of capital flows is examined employing the unobserved components model by Harvey (1989). Consider a panel data set of N countries with capital flows to country i $(i=1,\ldots,N)$ at time t generically denoted f_{it}. Dropping the subscript i for clarity, the unobserved components model may be written[20]:

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[19] As noted by Corbo and Hernandez (1993): "Reversing an initial adjustment could be quite costly if there are irreversible costs involved ... This is particularly relevant in the case of hot money."

[20] A possible extension of the model would be to consider one or more cyclical components as well as a seasonal component. We neglect this possibility because in estimating the model for a number of countries the inclusion of a cycle or of a seasonal component in the model was not found to improve the goodness of fit on the basis of the prediction error variance (PEV) and the Akaike Information Criterion (AIC).

$$f_t = \mu_t + \epsilon_t + \nu_t, \qquad t=1,\ldots,T \tag{3.10}$$

where f_t may be either equity flows (ef_t) or bond flows (bf_t), μ_t is a trend component, the irregular component ϵ_t is approximately normally independently distributed with zero mean and constant variance and ν_t represents a first-order autoregressive, AR(1) component:

$$\nu_t = \rho_\nu \nu_{t-1} + \xi_t \tag{3.11}$$

where ξ_t is approximately normally independently distributed with zero mean and constant variance and the autoregressive coefficient is constrained to be less than unity in absolute value in order to ensure stationarity of the autoregressive component[21].

The stochastic trend component is modelled as:

$$\mu_t = \mu_{t-1} + \beta_{t-1} + \eta_t \tag{3.12}$$

and

$$\beta_t = \rho_\beta \beta_{t-1} + \zeta_t \tag{3.13}$$

[21] The need to impose stationarity on the AR(1) process arises because of the risk of it being confounded with the random-walk component in the trend specification.

where β_t represents the slope or gradient of the trend component μ_t and ρ_β (less than or equal to unity in absolute value) represents the damping factor while the disturbances η_t and ζ_t are approximately normally independently distributed with zero mean and constant variance. The irregular component ϵ_t, the level disturbance η_t and the slope disturbance ζ_t are mutually uncorrelated. The slope component may be treated as fixed rather than stochastic and also excluded from the trend specification when this is appropriate.

Intuitively, (3.10) expresses the capital flow as the sum of a permanent component (μ_t), a purely temporary, zero persistence component (ϵ_t) and a more slowly decaying temporary component (ν_t). In addition, the drift in the random walk component (β_t) may itself vary over time. Thus the model separates out the permanent and temporary components of the data in a very general and comprehensive fashion[21].

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[21] Note that the structural time series model outlined above also nests an I(2) process for f_t. In particular, ignoring the AR(1) component for ease of exposition, the first difference of f_t, Δf_t, may be written as:

$$\Delta f_t = \beta_{t-1} + \eta_t + \epsilon_t - \epsilon_{t-1}$$

Since $\beta_{t-1}=\rho_\beta\beta_{t-2}+\zeta_{t-1}$, if $\sigma_\zeta=0$ or $\rho_\beta=0$, then Δf_t is the sum of an MA(1) process and a white noise process, so that $f_t \sim I(1)$: $f_t \sim ARIMA(0,1,1)$ (Granger and Morris, 1976). If $\sigma_\zeta \neq 0$, however: (a) if $\rho_\beta=1$ then $\Delta^2 f_t$ is the sum of a white noise process, an MA(1) and an MA(2) so $f_t \sim I(2)$: $f_t \sim ARIMA(0,2,2)$ (*ibid.*); (b) if $|\rho_\beta|<1$, then Δf_t is the sum of a stationary AR(1), an MA(1) and a white noise so $f_t \sim I(1)$: $f_t \sim ARIMA(1,1,2)$ (*ibid.*).

Nevertheless, as discussed in Section 3.8, when the stochastic slope is found to be statistically significantly different from zero at conventional nominal levels of significance, it is always the case that the estimated $|\rho_\beta|<1$, clearly implying that all the series modelled in this study are first-difference stationary, which also accords with the evidence, reported in Table 3.1, from ADF unit root tests.

This taxonomy also illustrates how the structural time series model may be viewed as a means of interpreting low-order ARIMA processes in terms of permanent and temporary components.

The statistical treatment of the unobserved components model outlined above is conveniently handled by writing it in state space form (SSF) involving a measurement equation relating the unobserved components (the state vector) to an observed series, together with a transition equation governing the evolution of the state vector. The state-space parameters are then estimated by maximum likelihood methods (Harvey, 1989; Cuthbertson, Hall and Taylor, 1992)[22].

The estimated hyperparameters (ie. the variance parameters) indicate the relative contribution of each component in the state vector to explaining the total variation in the time series under consideration. In some sense, therefore, the estimated hyperparameters allow us - by providing information on the size of the nonstationary and the stationary components in the series - to quantify the degree of persistence of the series in question. If a large and statistically significant proportion of flows is attributed to the stochastic level, for example, then one may expect that a large part of the portfolio flows will remain in the country concerned for an indeterminate period of time. By contrast, if a large portion of the variation in the time series is explained by movements in the temporary components, then the portfolio flows under consideration may be regarded as characterised by low persistence[23].

When the maximum likelihood estimate of the variance of an element of the state vector (ie. σ^2 or one of the diagonal elements of Q) is zero, the model can be

[22] A more detailed, technical exposition of the state space form representation of the unobserved components model is given in Appendix 3.1.

[23] In fact, while the irregular component is totally temporary, the AR(1) component displays some degree of positive persistence according to the size of the damping factor, still being mean-reverting.

reestimated making the corresponding component deterministic. Also, standard tests of the significance of the component itself can be carried out: if the component concerned is not found to be significantly different from zero, the model may be simplified by eliminating the component from the SSF altogether.

In choosing the most appropriate model for each country and label flow, we relied not only on standard measures such as the coefficient of determination: the fit of alternative models was also compared on the basis of the Akaike Information Criterion (AIC), ie. log(PEV)+2(m/T), where PEV is the steady state prediction error variance (Harvey, 1989, pp. 263-270), m represents the number of hyperparameters to be estimated plus the number of nonstationary components and T is the number of observations.

3.7.2 Variance ratio tests

The other measure of persistence we employ in this study is a simple nonparametric test, due originally to Cochrane (1988), generally referred to as the variance ratio test, z(k):

$$z(k) = \frac{1}{k} \frac{Var(f_{it} - f_{it-k})}{Var(f_{it} - f_{it-1})} \tag{3.14}$$

where k is a positive integer and Var stands for variance. If the series follows a random walk, then the ratio in (3.14) should be equal to unity. If capital flows exhibit mean reversion, then the ratio z(k) should be in the range between zero and unity.

Cochrane (1988) shows that the Bartlett estimator provides appropriate standard errors for z(k), with $Var(f_{i\,t-k})$ used for the random walk component[24]. In small samples, however, both the variance ratios and the Bartlett standard errors may be biased upwards and the asymptotic standard errors may not be a satisfactory approximation of the actual standard errors. Thus, in our empirical work, since we are dealing with small samples, we adopt the two corrections for small-sample bias suggested by Cochrane (1988, pp. 907-910). First, we use the sample mean of the first differences to estimate the drift term at all k rather than estimate a different drift term at each k from the mean of the k-differences. Second, we adopt a degrees of freedom correction, T/(T-k-1).

3.8 The persistence of portfolio flows: empirical results

In Table 3.6 (Panel A) we specify the various model specifications selected for portfolio flows to developing countries on the basis of the goodness of fit criteria discussed in Section 3.7.1.

Panels B-E of Table 3.6 report the results of estimating the most appropriate structural time series model - chosen by minimising the AIC[25] - in state space form by the Kalman filter maximum likelihood method for equity and bond flows to the Asian and Latin American countries examined.

Note that the convergence achieved by the BFGS numerical optimisation

[24] Thus, the standard error is consistently estimated as $(4k/3T)^{0.5}Var(f_{i\,t-k})$.

[25] Although the prediction error variance is the basic measure of goodness of fit, when a choice has to be made among alternative models with different numbers of hyperparameters, it is more appropriate to compare them on the basis of the AIC (Harvey, 1989).

method is always very strong in the sense that the three convergence criteria considered are always satisfied using the same tolerance level of 1.0E-7 for the likelihood kernel, the gradient and the parameters[26].

In the second and third columns of Table 3.6 (Panels B-E) we report details of the unobserved components included in the estimated structural time series model. In the fourth column we report the estimated standard deviations of the disturbances of the stochastic components included in the state and in parentheses we report the signal-to-noise ratios (ie. the ratios of each estimated standard deviation to the estimated standard deviation of the irregular component)[27]. As Panels B-E of Table 3.6 display, for both sets of countries and for both equity and bond flows the largest variance of the disturbances is always one of the stationary components in the state (either the irregular or the AR(1) component). Also, in 32 out of 36 cases, the largest estimated hyperparameter is the variance of the disturbance of the irregular component, therefore suggesting that the time series has virtually no persistence at all. In the remaining four cases, the AR(1) hyperparameter has the largest variance, which implies some slight degree of persistence. Note, however, that this persistence is also very small, as suggested by the fact that the estimated damping factors in the AR(1), reported in the fifth column of Panels B - E, components are always relatively low. Indeed the half life of shocks affecting portfolio flows ranges from 0.39 (about one and a half weeks)

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[26] See Appendix 3.1 for a technical treatment of the Kalman filter recursions.

[27] In terms of the state-space presented in Appendix 3.1, the Q-ratios are the square roots of the estimated diagonal elements of Q, and so give the ratio of the standard deviation of the relevant element of the state vector to the standard deviation of the irregular component.

for equity flows to Venezuela to 0.84 (about three and a half weeks) for bond flows to Ecuador.

Moreover, the Q-ratios for the stochastic level and especially for the stochastic slope, when it is found to be statistically significantly different from zero, are very low, suggesting that the contribution of the nonstationary, more persistent, components to explaining the variance of equity and bond flows is extremely low while the temporary component is, by contrast, very large. Nevertheless, a significant stochastic level is always found to be statistically significantly different from zero at conventional nominal levels of significance, as displayed by the fifth column in Table 3.6 (Panels B-E), which reports the estimated coefficients of the final state vector and the respective t-value for the nonstationary stochastic components included in the estimated model.

The sixth and seventh columns in Table 3.6 (Panels B-E) report goodness of fit statistics, namely the coefficient of determination - which may be regarded as quite high in all of the estimated models - and the AIC. Finally, in the last column of Panels B-E we report Ljung-Box test statistics for residual serial correlation.

Overall, therefore, the results of the estimation of the unobserved components model for equity and bond flows to developing countries suggest that a statistically significant nonstationary component is present in the data, but that this is generally very small in size, contributing very little to explaining variation in the series. That is to say, portfolio flows to Asian and Latin American countries over the sample period may be regarded as almost entirely temporary.

Our last exercise is to compute some variance ratio tests corrected for small-sample bias in the fashion suggested by Cochrane (1988, pp. 908-910). The estimated variance ratio tests provide qualitatively identical results for all portfolio flows examined in this study. Hence, in order to save space, we do not report the results for all portfolio flows (although they are available from the author). However, a reasonable representative of the results from computing small-sample adjusted variance ratio tests on portfolio flows data may be Indonesia:

Lag	Equity Flows	Bond Flows
1	1.0000 (0.1529)	1.0000 (0.1529)
2	0.6423 (0.1389)	0.8473 (0.1833)
3	0.4675 (0.1238)	0.5092 (0.1349)
6	0.3115 (0.1167)	0.3362 (0.1259)
9	0.1239 (0.0569)	0.2573 (0.1180)
12	0.0990 (0.0524)	0.1229 (0.0651)
18	0.1167 (0.0757)	0.1022 (0.0663)
24	0.0861 (0.0645)	0.0974 (0.0730)
30	0.0014 (0.0011)	0.70E-3 (0.59E-3)
36	0.0011 (0.0010)	0.0015 (0.0014)

where Bartlett standard errors $[=(4k/3T)^{0.5}Var(f_{ik})]$ are reported in parentheses. The results display a clear, fast decline towards zero of the variance ratio, which also lies between zero and unity. Moreover, the variance ratios are, in most cases, statistically significant at conventional nominal levels of significance only for 9-12 lags (ie. up to one year). These results therefore corroborate the Kalman filter results, clearly suggesting that both equity and bond flows to developing countries have *at most* a very small persistent component.

3.9 Conclusions

In this study we have examined the determinants of US capital flows directed to nine Latin American and nine Asian countries over the sample period 1988-92, extending previous work by Chuhan, Claessens and Mamingi (1993). In particular, we have investigated whether the large capital inflows in the form of bonds and equities received by the countries examined were induced by improvements in the economic performances of those countries - country-specific or pull factors - or by factors related to US economic performance - global or push factors, and have explicitly differentiated between short-run and long-run determinants.

We considered in our set of country-specific factors the domestic credit rating and the black market exchange rate premium, while the set of global factors included two US interest rates and the level of US real industrial production (Calvo, Leiderman and Reinhart, 1993, 1996). We examined the *long-run* determinants of portfolio flows by employing two complementary cointegration techniques to examine the long-run relationship between capital flows and both sets of potential explanatory variables. The results provide unequivocal evidence that long-run equity and bond flows are about equally sensitive to global factors and to country-specific factors, and therefore that both sets of variables contribute in explaining US portfolio flows to the developing countries considered.

Moreover, we also estimated seemingly unrelated error-correction models for equity and bond flows to the developing countries considered in order to shed light on the underlying short-run dynamics of US portfolio flows. The estimated error-correction models for portfolio flows to Asian developing countries, especially for equity flows, are more complex, in terms of the short-run dynamics, than the error-correction models for portfolio flows to Latin American countries.

A count of the number of significant pull and push factors appearing in the error correction forms, classified by type of flow and geographic area, revealed the following. As far as equity flows are concerned, global (push) and country-specific (pull) factors seem to be equally important in determining short-run movements for both Asian and Latin American countries. When bond flows are considered, however, global factors seem to be much more important than domestic factors in explaining the short-run dynamics of flows for both sets of countries. In particular, the change in US interest rates is found to be the single most important determinant of short-run movements in bond flows to the emerging markets.

Also, we modelled portfolio flows to a set of Asian and Latin American countries by representing unobserved components time series models in state space form and estimating them using Kalman filtering maximum likelihood techniques in order to quantify the size of the unobserved nonstationary (permanent) and stationary (temporary) components present in these equity and bond flows. In addition, we employed some simple non-parametric variance ratio tests corrected for small sample bias to measure the degree of persistence present in these portfolio flows.

Overall, the empirical results strongly suggest that the permanent - or at least highly persistent - component in equity and bond flows to developing countries is very small in size relative to the temporary component.

From the policy maker's point of view, therefore, given the very large temporary component of capital inflows, developing countries should be wary of implementing painful adjustment processes that may need to be reversed in the future. Especially for countries with a more flexible exchange rate regime, capital inflows can potentially generate excessive volatility in both the nominal and the real exchange rate (Kiguel and Caprio, 1993; Corbo and Hernandez, 1993; Calvo, Leiderman and Reinhart, 1996). The macroeconomic repercussions of the increase of portfolio flows of a highly temporary nature in the 1990s may be very high for several developing countries.

Also, as first argued by Claessens, Dooley and Warner (1995), accounting labels attached to capital flows may be meaningless in providing information about the properties of the flows, since bond flows, which should be relatively more persistent from an accounting point of view, may be modelled in a similar fashion to equity flows, which one might in principle expect to be relatively less persistent, therefore requiring the same type of policy response needed against flows which are identified as "hot money".

Appendix 3.1 The Kalman filter recursions

The unobserved components model outlined in Section 3.7.1 may be represented in state space form (SSF), where a measurement equation relates the unobserved components (the state vector) to an observed series, while a transition equation governes the evolution of the state vector. The SSF corresponding to the model outlined in (3.10)-(3.13) may be written:

$$f_t = (1 \ 0 \ 1) \begin{pmatrix} \mu_t \\ \beta_t \\ \nu_t \end{pmatrix} + \epsilon_t \tag{A3.1.1}$$

$$\begin{pmatrix} \mu_t \\ \beta_t \\ \nu_t \end{pmatrix} = \begin{pmatrix} 1 & 1 & 0 \\ 0 & \rho_\beta & 0 \\ 0 & 0 & \rho_\nu \end{pmatrix} \begin{pmatrix} \mu_{t-1} \\ \beta_{t-1} \\ \nu_{t-1} \end{pmatrix} + \begin{pmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} \eta_t \\ \zeta_t \\ \xi_t \end{pmatrix} \tag{A3.1.2}$$

where (A3.1.1) represents the measurement equation, which shows how an observed series is related to the state vector, whereas (A3.1.2) is the transition equation, describing the dynamic evolution of the state vector. The SSF given by (A3.1.1) and (A3.1.2) may itself be written in a more compact form, using an obvious notation, as:

$$f_t = z' x_t + \epsilon_t, \qquad t=1,\ldots,T \tag{A3.1.3}$$

$$x_t = M\, x_{t-1} + R\, \kappa_t, \qquad t=1,\ldots,T \tag{A3.1.4}$$

where:

$$\epsilon_t \sim IN(0, \sigma^2) \qquad \text{(A3.1.5)}$$

$$\kappa_t \sim IN(0, \sigma^2 Q) \qquad \text{(A3.1.6)}$$

z' is a known (3x1) vector, M and R are fixed matrices of order (3x3) and Q is also a fixed (3x3) matrix. The (3x1) vector x_t is the unobservable state vector.

Given knowledge of the parameters of the SSF, the Kalman filter provides us with optimal estimates of x_t, using either information up to time t-1 (the prediction equations), information up to time t (the updating equations), or the full sample information (the smoothing equations) (Kalman, 1960a,b)[28].

Suppose we have an optimal estimator of x_{t-1} using all information up to time t-1, and denote this $\tilde{x}_{t-1}$. Then the prediction equation providing us with the optimal predictor of x_t using information up to time t-1, denoted $\tilde{x}_{t|t-1}$, is:

$$\tilde{x}_{t|t-1} = M \, \tilde{x}_{t-1} \qquad \text{(A3.1.7)}$$

and the covariance matrix of $\tilde{x}_{t|t-1}$ can be shown to be given by:

$$P_{t|t-1} = M \, P_{t-1} \, M' + R \, Q \, R' \qquad \text{(A3.1.8)}$$

where P_{t-1} is the covariance matrix of $\tilde{x}_{t|t-1}$. Equations (A3.1.7) and (A3.1.8) describe the prediction equations of the Kalman filter.

The updating equations, which update these predictions on the basis of information at time t, are given by (A3.1.9) and (A3.1.10):

[28] See Harvey (1989), or Cuthbertson, Hall and Taylor (1992) for an accessible introduction to the Kalman filter.

$$\tilde{x}_t = \tilde{x}_{t|t-1} + P_{t|t-1}z(f_t - z'\tilde{x}_{t|t-1})/g_t \tag{A3.1.9}$$

$$P_t = P_{t|t-1} - P_{t|t-1}zz'P_{t|t-1}/g_t \tag{A3.1.10}$$

where $g_t = z'P_{t|t-1}z+1$ [29].

Given a finite sequence of observations f_t, the only state vector estimator which uses all the available information is $\tilde{x}_t$. The smoothing equations, given by (A3.1.11) and (A3.1.12), describe optimal, full sample information estimators:

$$\tilde{x}_{t|T} = \tilde{x}_t + P_t^*(\tilde{x}_{t+1|T} - M\,\tilde{x}_t) \tag{A3.1.11}$$

$$P_{t|T} = P_t - P_t^*(P_{t+1|T} - P_{t+1|t})\,P_T^{*\prime} \tag{A3.1.12}$$

where $P_t^* = P_tM/P_{t+1|t}$, $\tilde{x}_{T|T} = \tilde{x}_T$ and $P_{T|T} = P_T$. Equations (A3.1.7)-(A3.1.12) describe the Kalman filter recursions (Kalman, 1960a,b).

The state space parameters can in practice be estimated by maximum likelihood methods (Harvey, 1989; Cuthbertson, Hall and Taylor, 1992). A natural by-product of the Kalman filter recursions is a sequence of one-step-ahead prediction errors, u_t, defined by:

$$u_t = f_t - z'\tilde{x}_{t|t-1} \tag{A3.1.13}$$

Using a result due to Schweppe (1965), the likelihood function for the sample, $\pounds(\cdot)$,

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[29] The term $(f_t - z'\tilde{x}_{t|t-1})$ in (A3.1.7) is the prediction error. This innovation contains all the new information in f_t and is used to update $\tilde{x}_{t|t-1}$ via the Kalman gain, which is the (3x1) vector $(P_{t|t-1}z)/g_t$ which essentially decides what weight to assign to the innovation.

can be derived and expressed in terms of the innovations u_t and their variances, g_t:

$$\mathcal{L}(f_1,\ldots,f_T) = -\frac{T}{2}\log(2\pi) - \frac{T}{2}\log(\sigma^2) - \frac{1}{2}\sum_{t=1}^{T}\log(g_t) - \frac{1}{2\sigma^2}\sum_{t=1}^{T}\frac{u_t^2}{g_t} \tag{A3.1.14}$$

The likelihood function obtained may then be maximised with respect to the hyperparameters[30] using numerical optimisation procedures (see eg. Cuthbertson, Hall and Taylor, 1992, Chapter 2). In our empirical work, we employ the Broyden-Fletcher-Golgfarb-Shanno quasi-Newton algorithm (Harvey, 1981), based on three convergence criteria - the likelihood kernel, the gradient and the parameters.

[30] The scale factor σ^2 may always be concentrated out of (A3.1.14) by substituting:

$$\hat{\sigma}^2 = \frac{1}{T}\sum_{t=1}^{T} u_t^2/g_t$$

so that the maximisation of the likelihood function becomes equivalent to minimising the function:

$$F = T\log(\hat{\sigma}^2) + \sum_{t=1}^{T}\log(g_t)$$

Concentrating σ^2 out of the likelihood function reduces the dimension of the search involved in the numerical optimisation procedure.

Table 3.1 Unit root tests

Panel A: Global variables

y_t	Δy_t	i_t	Δi_t	r_t	Δr_t
-2.6961	-5.4810	-2.7360	-3.4684	-1.7215	-5.2062

Note: The number of lags included is chosen such that the estimated disturbance is approximately white noise. In testing for a unit root on the variables in levels, a constant and a trend are included (and found statistically significant at conventional nominal significance levels) in the regression; the relevant critical values are -3.50 (5%) and -3.18 (10%). The trend is eliminated in testing for a unit root on the variables in first difference and the critical values are -2.93 (5%) and -2.60 (10%).

Panel B: Asian portfolio flows

Country	ef_{it}	Δef_{it}	bf_{it}	Δbf_{it}
China	-0.5248	-4.4164	-1.5941	-3.8151
India	0.2157	-4.4789	-1.2234	-3.9799
Indonesia	-1.3034	-4.6493	-1.0389	-3.2180
Korea	-0.8939	-3.6372	-1.5115	-4.8245
Malaysia	-1.5453	-3.5428	-1.3370	-3.9836
Pakistan	1.3010	-5.7416	-1.4324	-4.1918
Philippines	0.2897	-3.4286	-1.3165	-4.9880
Taiwan	-1.5266	-4.1754	-1.1413	-3.0646
Thailand	-1.1840	-3.5842	-1.3197	-4.4809

See Note to Panel A.

Panel C: Latin American portfolio flows

Country	ef_{it}	Δef_{it}	bf_{it}	Δbf_{it}
Argentina	-1.2721	-6.2347	0.8670	-3.4676
Brazil	-1.4989	-3.6320	-1.3302	-3.9369
Chile	-1.7486	-3.8676	-1.3684	-5.2494
Colombia	-1.5285	-3.9302	-1.8662	-4.2347
Ecuador	-0.8936	-4.6504	-1.9590	-3.9517
Jamaica	-1.4659	-4.1918	-0.6671	-5.8417
Mexico	0.7812	-5.6005	1.1217	-3.4337
Uruguay	-0.6484	-3.9791	1.3078	-4.5792
Venezuela	-0.8500	-5.3234	-0.0637	-5.4417

See Note to Panel A.

Panel D: Asian domestic factors

Country	cr_{it}	Δcr_{it}	bm_{it}	Δbm_{it}
Argentina	-1.6445	-7.7209	-1.8777^{t}	-5.6641
Brazil	-1.3301^{t}	-8.0871	-2.5466	-8.5098
Chile	-0.8781	-8.5323	-2.1616^{t}	-5.7510
Colombia	-2.8472	-7.5528	-2.0466	-8.4026
Ecuador	-1.1980	-8.3368	-3.9660*	-10.9186
Jamaica	-1.5163^{t}	-5.6234	-2.5214	-7.5662
Mexico	-2.0738	-7.3867	-5.1564*	-10.5621
Uruguay	-2.2326	-7.4828	-3.5186*	-10.4215
Venezuela	-1.8316	-7.8401	-2.9789^{t}	-6.4717

Note: The number of lags included is chosen such that the estimated disturbance is approximately white noise. In testing for a unit root on the variables in levels, a constant is always included and a trend is included if found to be statistically significant at the conventional nominal significance levels - ADF statistics drawn from regressions where a trend was included are depicted with the superscript t. The trend is eliminated in testing for a unit root on the variables in first difference. The asterisk denotes that the null hypothesis of nonstationarity can be rejected at the 5% nominal level of significance. The relevant critical values if both constant and trend are present in the regression are -3.50 (5%) and -3.18 (10%), while if only a constant is present in the regression the critical values are -2.93 (5%) and -2.60.

Panel E: Latin American domestic factors

Country	cr_{it}	Δcr_{it}	bm_{it}	Δbm_{it}
Argentina	-0.4110^{t}	-7.2953	-5.5522*	-12.6013
Brazil	-1.8181	-12.1219	-3.0067^{t}	-8.4141
Chile	-5.1971t*	-11.8500	-2.1381	-6.9522
Colombia	-2.9731^{t}	-7.2809	-3.0409^{t}	-9.5964
Ecuador	-2.1883^{t}	-7.5398	-2.6205	-7.6318
Jamaica	-1.4030	-7.8738	-3.4846^{t}	-6.5092
Mexico	-2.3990^{t}	-10.3578	-2.5843^{t}	-7.8405
Uruguay	-2.5163^{t}	-7.7817	-4.3063^{t}*	-11.4835
Venezuela	-1.4463^{t}	-7.0342	-1.9456^{t}	-6.9795

See Note to Panel D.

Table 3.2 Augmented Dickey-Fuller cointegration tests

Panel A: Asian equity flows

	ADF1	ADF2	ADF3
China	-8.4265 (-5.0141)	-8.2182 (-3.8963)	-8.4364 (-4.2988)
India	-6.2328 (-5.0257)	-6.4854 (-3.9053)	-4.4082 (-4.3064)
Indonesia	-7.1897 (-5.0141)	-6.8962 (-3.8963)	-5.6371 (-4.2988)
Korea	-6.7943 (-5.0141)	-4.4005 (-3.8962)	-6.4559 (-4.2988)
Malaysia	-7.0122 (5.0141)	-6.5245 (-3.8962)	-6.0072 (-4.2988)
Pakistan	-1.7677 (-5.0382)*	-2.3677 (-3.9085)*	-0.9119 (-4.3146)*
Philippines	-6.1994 (-5.0141)	-5.2261 (-3.8963)	-6.1861 (-4.2988)
Taiwan	-8.5079 (-5.0141)	-2.7677 (-3.9085)	-2.6200 (-4.3146)
Thailand	-5.9094 (-5.0141)	-4.1257 (-3.8963)	-5.8286 (-4.2988)

Note: ADF1, ADF2 and ADF3 are augmented Dickey-Fuller test statistics computed on the residuals from the regression of portfolio flows to the country considered on both country-specific and global factors, only country specific factors and only global factors respectively. Critical values are reported in parentheses. The number of lags included is such that the error term is approximately white noise. The symbol * is associated with regression where the null hypothesis of no cointegration cannot be rejected at standard nominal levels of significance.

Panel B: Asian bond flows

	ADF1	ADF2	ADF3
China	-6.4286 (-5.0141)	-6.1033 (-3.8963)	-6.0844 (-4.2988)
India	-8.2003 (-5.0141)	-7.0188 (-3.8963)	-7.3904 (-4.2988)
Indonesia	-5.4419 (-5.0141)	-5.3780 (-3.8963)	-5.1129 (-4.2988)
Korea	-5.7214 (-5.0198)	-7.4051 (-3.8963)	-8.0958 (-4.2988)
Malaysia	-7.6344 (5.0141)	-7.4688 (-3.8963)	-7.6126 (-4.2988)
Pakistan	-7.8331 (-5.0141)	-7.7325 (-3.8963)	-7.7011 (-4.2988)
Philippines	-4.2752 (-5.0198)*	-4.1574 (-3.8992)	-3.9047 (-4.3025)*
Taiwan	-5.4774 (-5.0141)	-2.8194 (-3.8992)*	-4.1504 (-4.2988)*
Thailand	-7.8268 (-5.0141)	-7.7779 (-3.8992)	-8.4181 (-4.2988)

See Note to Panel A.

Panel C: Latin American equity flows

	ADF1	ADF2	ADF3
Argentina	-7.3331 (-5.0318)	-3.5075 (-3.8992)*	-5.1805 (-4.3146)
Brazil	-5.5815 (-5.0198)	-5.3729 (-3.8963)	-6.2050 (-4.2988)
Chile	-6.0774 (-5.0141)	-4.7966 (-3.8963)	-4.9482 (-4.2988)
Colombia	-7.5776 (-5.0141)	-6.1239 (-3.8963)	-6.9158 (-4.2988)
Ecuador	-10.6141 (5.0141)	-10.7085 (-3.8963)	-10.1457 (-4.2988)
Jamaica	-7.9104 (-5.0141)	-7.7771 (-3.8963)	-7.5235 (-4.2988)
Mexico	-3.8560 (-5.0318)*	-5.0159 (-3.9053)	-3.6606 (-4.3146)*
Uruguay	-8.1749 (-5.0141)	-7.5328 (-3.9053)	-8.1380 (-4.2988)
Venezuela	-2.7387 (-5.0382)*	-2.5070 (-3.9085)*	-2.2916 (-4.3146)*

See Note to Panel A.

Panel D: Latin American bond flows

	ADF1	ADF2	ADF3
Argentina	-7.6526 (-5.0141)	-5.5257 (-3.8963)	-7.3757 (-4.3146)
Brazil	-8.7683 (-5.0198)	-7.4884 (-3.8963)	-8.7535 (-4.2988)
Chile	-8.8537 (-5.0141)	-8.3542 (-3.8963)	-8.6428 (-4.2988)
Colombia	-7.8772 (-5.0141)	-6.9796 (-3.8963)	-8.0583 (-4.2988)
Ecuador	-8.2530 (5.0141)	-7.4536 (-3.8963)	-8.1222 (-4.2988)
Jamaica	-6.2529 (-5.0318)	-3.7397 (-3.9053)*	-4.3974 (-4.2988)
Mexico	-2.3368 (-5.0382)*	-3.8186 (-3.9022)*	-2.3687 (-4.3146)*
Uruguay	-5.8362 (-5.0198)	0.8025 (-3.9022)*	-0.8631 (-4.2988)*
Venezuela	-3.6914 (-5.0257)*	-5.1953 (-3.8963)	-3.2371 (-4.3146)*

See Note to Panel A.

Table 3.3 Johansen cointegration tests

Panel A: Asian equity flows

	H_0	H_1	λ_{max}	λ_{trace}	95% c.v.	90% c.v.	χ^2 test on β	χ^2 test on γ
Pakistan	r=0	r=1	121.69		39.43	36.35		
	r<=1	r=2	54.25		33.32	30.84		
	r<=2	r=3	43.21		27.14	24.78		
	r<=3	r=4	23.30		21.07	18.90		
	r<=4	r=5	16.50		14.90	12.91		
	r<=5	r=6	3.02		8.18	6.50		
	r=0	r>=1		261.97	95.18	90.39		
	r<=1	r>=2		140.28	70.60	66.49		
	r<=2	r>=3		86.03	48.28	45.23		
	r<=3	r>=4		42.82	31.52	28.71		
	r<=4	r>=5		19.52	17.95	15.66		
	r<=5	r=6		3.02	8.18	6.50		
							157.25 (0.00)	166.34 (0.00)

Note: The critical values for the λ_{max} and λ_{trace} test statistics are provided by Johansen and Juselius (1990). In the last two columns we report the likelihood ratio test for the zero restriction on the coefficients on the country-specific factors (β) and on the global factors (γ) respectively.

Panel B: Asian bond flows

	H_0	H_1	λ_{max}	λ_{trace}	95% c.v.	90% c.v.	χ^2 test on β	χ^2 test on γ
Philippines	r=0	r=1	112.96		39.43	36.35		
	r<=1	r=2	73.13		33.32	30.84		
	r<=2	r=3	53.10		27.14	24.78		
	r<=3	r=4	34.77		21.07	18.90		
	r<=4	r=5	7.67		14.90	12.91		
	r<=5	r=6	4.47		8.18	6.50		
	r=0	r>=1		286.10	95.18	90.39		
	r<=1	r>=2		173.14	70.60	66.49		
	r<=2	r>=3		100.01	48.28	45.23		
	r<=3	r>=4		46.91	31.52	28.71		
	r<=4	r>=5		12.14	17.95	15.66		
	r<=5	r=6		4.47	8.18	6.50		
							150.23 (0.00)	191.71 (0.00)

See Note to Panel A.

Panel C: Latin American equity flows

	H_0	H_1	λ_{max}	λ_{trace}	95% c.v.	90% c.v.	χ^2 test on β	χ^2 test on γ
Mexico	r=0	r=1	115.90		39.43	36.35		
	r<=1	r=2	41.48		33.32	30.84		
	r<=2	r=3	31.67		27.14	24.78		
	r<=3	r=4	20.95		21.07	18.90		
	r<=4	r=5	13.31		14.90	12.91		
	r<=5	r=6	2.40		8.18	6.50		
	r=0	r>=1		225.72	95.18	90.39		
	r<=1	r>=2		109.82	70.60	66.49		
	r<=2	r>=3		68.34	48.28	45.23		
	r<=3	r>=4		36.66	31.52	28.71		
	r<=4	r>=5		17.71	17.95	15.66		
	r<=5	r=6		2.40	8.18	6.50		
							68.69 (0.00)	99.99 (0.00)
Venezuela	r=0	r=1	64.68		39.43	36.35		
	r<=1	r=2	48.05		33.32	30.84		
	r<=2	r=3	26.98		27.14	24.78		
	r<=3	r=4	21.02		21.07	18.90		
	r<=4	r=5	10.06		14.90	12.91		
	r<=5	r=6	0.34		8.18	6.50		
	r=0	r>=1		171.13	95.18	90.39		
	r<=1	r>=2		106.45	70.60	66.49		
	r<=2	r>=3		58.40	48.28	45.23		
	r<=3	r>=4		31.41	31.52	28.71		
	r<=4	r>=5		10.39	17.95	15.66		
	r<=5	r=6		0.34	8.18	6.50		
							79.75 (0.00)	93.08 (0.00)

See Note to Panel A.

Panel D: Latin American bond flows

	H_0	H_1	λ_{max}	λ_{trace}	95% c.v.	90% c.v.	χ^2 test on β	χ^2 test on γ
Mex.	r=0	r=1	112.42		39.43	36.35		
	r<=1	r=2	86.38		33.32	30.84		
	r<=2	r=3	29.77		27.14	24.78		
	r<=3	r=4	24.77		21.07	18.90		
	r<=4	r=5	17.57		14.90	12.91		
	r<=5	r=6	1.368		8.18	6.50		
	r=0	r>=1		284.59	95.18	90.39		
	r<=1	r>=2		172.17	70.60	66.49		
	r<=2	r>=3		85.79	48.28	45.23		
	r<=3	r>=4		56.02	31.52	28.71		
	r<=4	r>=5		31.25	17.95	15.66		
	r<=5	r=6		1.368	8.18	6.50		
							111.22 (0.00)	145.17 (0.00)
Venez.	r=0	r=1	84.89		39.43	36.35		
	r<=1	r=2	60.07		33.32	30.84		
	r<=2	r=3	48.63		27.14	24.78		
	r<=3	r=4	21.50		21.07	18.90		
	r<=4	r=5	11.31		14.90	12.91		
	r<=5	r=6	4.89		8.18	6.50		
	r=0	r>=1		231.29	95.18	90.39		
	r<=1	r>=2		146.40	70.60	66.49		
	r<=2	r>=3		86.33	48.28	45.23		
	r<=3	r>=4		37.70	31.52	28.71		
	r<=4	r>=5		16.20	17.95	15.66		
	r<=5	r=6		4.89	8.18	6.50		
							87.04 (0.00)	114.10 (0.00)

See Note to Panel A.

Table 3.4 Seemingly unrelated ECMs: number of significant lags of right-hand-side variables

Panel A: Asian equity flows

	e	ldv	cr	bm	i	r	y
China	1	1	1	0	0	0	1
India	1	0	2	2	3	1	3
Indonesia	1	0	0	2	1	0	1
Korea	1	1	2	3	2	2	0
Malaysia	1	0	3	2	0	2	1
Pakistan	1	1	0	2	0	0	0
Philippines	1	0	1	2	0	0	0
Taiwan	1	1	1	1	1	1	1
Thailand	1	0	1	0	1	1	1

Note: e, ldv, cr, bm, i, r and y denote the error correction term, the once-lagged dependent variable, the country's credit rating, the black market exchange rate premium, US short-term interest rate, US long-term interest rate and US real industrial production respectively.

Panel B: Asian bond flows

	e	ldv	cr	bm	i	r	y
China	1	1	0	0	2	0	1
India	1	0	1	2	0	2	1
Indonesia	1	1	2	0	1	1	0
Korea	1	1	0	1	0	1	0
Malaysia	1	0	0	0	2	2	0
Pakistan	1	0	1	1	0	0	3
Philippines	1	1	0	3	1	0	3
Taiwan	1	1	1	2	0	1	0
Thailand	1	0	0	2	2	1	0

See Note to Panel A.

Panel C: Latin American equity flows

	e	ldv	cr	bm	i	r	y
Argentina	1	1	2	0	1	0	0
Brazil	1	1	1	2	2	0	1
Chile	1	0	3	0	0	1	1
Colombia	1	0	0	2	0	1	0
Ecuador	1	0	1	1	0	2	0
Jamaica	1	0	0	0	1	2	0
Mexico	1	0	2	0	1	0	1
Uruguay	1	0	0	1	2	1	1
Venezuela	1	0	1	1	2	1	0

See Note to Panel A.

Panel D: Latin American bond flows

	e	ldv	cr	bm	i	r	y
Argentina	1	0	0	1	2	0	0
Brazil	1	1	0	1	0	0	1
Chile	1	0	1	0	2	1	3
Colombia	1	0	2	0	1	0	0
Ecuador	1	0	0	1	1	2	1
Jamaica	1	1	1	2	2	0	1
Mexico	1	0	2	0	1	1	0
Uruguay	1	0	0	1	4	0	1
Venezuela	1	1	1	0	1	0	1

See Note to Panel A.

Table 3.5 Summary table: number of significant country-specific and global factors in the ECMs

Panel A: Results for Asian developing countries

Capital flow type	Credit rating	Black market premium	US interest rates	US real industrial production	Total
Equities	11	14	15	9	49
Bonds	5	11	16	8	40
Total	16	25	31	17	89

Panel B: Results for Latin American developing countries

Capital flow type	Credit rating	Black market premium	US interest rates	US real industrial production	Total
Equities	10	7	17	4	38
Bonds	7	6	18	8	39
Total	17	13	35	12	77

Table 3.6 Results from estimating unobserved components models

Panel A: Structural time-series models adopted in modelling portfolio flows

Model 1:

$$f_t = \mu_t + \epsilon_t$$

$$\mu_t = \mu_{t-1} + \eta_t$$

Model 2:

$$f_t = \mu_t + \epsilon_t$$

$$\mu_t = \mu_{t-1} + \beta + \eta_t$$

Model 3:

$$f_t = \mu_t + \epsilon_t$$

$$\mu_t = \mu_{t-1} + \beta_{t-1} + \eta_t$$

$$\beta_t = \beta_{t-1} + \zeta_t$$

Model 4:

$$f_t = \mu_t + \epsilon_t + \nu_t$$

$$\mu_t = \mu_{t-1} + \eta_t$$

$$\nu_t = \rho_\nu \nu_{t-1} + \xi_t \qquad |\rho_\nu|<1$$

Model 5:

$$f_t = \mu_t + \epsilon_t$$

$$\mu_t = \mu_{t-1} + \beta_{t-1} + \eta_t$$

$$\beta_t = \rho_\beta \beta_{t-1} + \zeta_t \qquad |\rho_\beta|\leq 1$$

Panel B: Kalman filter estimation: Asian equity flows

Country	Model	Component	Est. s.d. of disturb. and Q-ratios	Est. coef. of final state vector [t-values]	R^2	AIC	LB(9,q)
China	1	Stc Lvl, No Slp, Irr	Irr: 55.489 (1.000) Lvl: 10.141 (0.183)	Lvl: 9.314 [2.688]*	0.919	-1.122	5.36 {0.50}
India	2	Stc Lvl, Fxd Slp, Irr	Irr: 12.453 (1.000) Lvl: 3.658 (0.294) Slp: 0.954 (0.077)	Lvl: 5.263 [2.662]* Slp: 0.388 [2.985]**	0.803	0.565	7.13 {0.31}
Indonesia	2	Stc Lvl, Fxd Slp, Irr	Irr: 5.663 (1.000) Lvl: 3.748 (0.662)	Lvl: 9.268 [2.343]* Slp: 0.169 [2.315]*	0.869	4.191	12.11 {0.59}
Korea	3	Stc Lvl, Stc Slp, Irr	Irr: 88.190 (1.000) Lvl: 22.061 (0.250) Slp: 3.213 (0.036)	Lvl: 2.366 [5.389]** Slp: 0.828 [2.124]*	0.866	4.650	3.67 {0.72}
Malaysia	1	Stc Lvl, No Slp, Irr	Irr: 88.257 (1.000) Lvl: 7.752 (0.088)	Lvl: 7.632 [2.601]*	0.925	5.902	9.93 {0.13}
Pakistan	1	Stc Lvl, No Slp, Irr	Irr: 0.177 (1.000) Lvl: 0.073 (0.414)	Lvl: 7.353 [2.968]**	0.867	-0.828	9.01 {0.17}
Philipp.	3	Stc Lvl, Stc Slp, Irr	Irr: 7.179 (1.000) Lvl: 0.995 (0.138) Slp: 0.160 (0.022)	Lvl: 1.026 [6.035]** Slp: 0.492 [3.004]**	0.883	2.913	7.76 {0.26}
Taiwan	1	Stc Lvl, No Slp, Irr	Irr: 45.446 (1.000) Lvl: 7.549 (0.166)	Lvl: 4.637 [8.431]**	0.822	1.183	4.57 {0.60}
Thailand	1	Stc Lvl, No Slp, Irr	Irr: 66.217 (1.000) Lvl: 13.741 (0.416)	Lvl: 1.359 [3.267]**	0.801	4.300	10.19 {0.12}

Note: The Q-ratios (computed as the ratio of the standard deviation of each component to the standard deviation of the irregular component) are reported in parentheses in the fourth column; in the fifth column we report the t-values (ie. the ratio of each estimated coefficient to its root mean square error) in square brackets - the * (**) symbol indicates significance of the component concerned at the 5% (1%) marginal level of significance. The LB(p,q) - where p is the number of lags considered and q is the difference between p and the number of hyperparameters in the state -is the Ljung-Box test statistic for serial correlation; under the null hypothesis of no serial correlation, LB(p,q) is distributed as χ^2 with q degrees of freedom (marginal significance levels are reported in braces). Abbreviations used are: est. for estimated, disturb. for disturbance, s.d. for standard deviation, stc for stochastic, lvl for level, fxd for fixed, dmp for damped, slp for slope, irr for irregular.

Panel C: Kalman Filter estimation: Asian bond flows

Country	Model	Component	Est. s.d. of disturb. and Q-ratios	Est. coef. of final state vector [t-values]	R^2	AIC	LB(9,q)
China	2	Stc Lvl, Fxd Slp, Irr	Irr: 0.641 (1.000) Lvl: 0.096 (0.150)	Lvl: 5.463 [2.998]** Slp: 0.553 [2.343]*	0.863	4.209	9.45 {0.15}
India	1	Stc Lvl, No Slp, Irr	Irr: 1.140 (1.000) Lvl: 0.337 (0.296)	Lvl: 2.911 [3.409]**	0.890	5.213	4.22 {0.65}
Indonesia	1	Stc Lvl, No Slp, Irr	Irr: 5.248 (1.000) Lvl: 2.334 (0.445)	Lvl: 4.836 [2.036]*	0.807	4.334	9.24 {0.16}
Korea	1	Stc Lvl, No Slp, Irr	Irr: 9.570 (1.000) Lvl: 1.164 (0.122)	Lvl: 1.096 [3.061]**	0.831	4.655	10.89 {0.09}
Malaysia	1	Stc Lvl, No Slp, Irr	Irr: 17.250 (1.000) Lvl: 1.741 (0.101)	Lvl: 5.882 [2.457]*	0.863	4.649	0.31 {0.99}
Pakistan	4	Stc Lvl, No Slp, Irr, AR(1)	Irr: 0.141 (1.000) Lvl: 0.074 (0.525) AR(1): 0.544 (3.858)	Lvl: 3.511 [2.515]* AR(1): 0.356	0.902	-3.243	1.37 {0.93}
Philipp.	3	Stc Lvl, Stc Slp, Irr	Irr: 3.265 (1.000) Lvl: 0.402 (0.123) Slp: 0.233 (0.071)	Lvl: 1.176 [3.258]** Slp: 0.815 [3.410]**	0.921	4.262	6.19 {0.40}
Taiwan	1	Stc Lvl, No Slp, Irr	Irr: 7.985 (1.000) Lvl: 3.489 (0.437)	Lvl: 2.153 [8.217]**	0.886	4.047	9.83 {0.13}
Thailand	4	Stc Lvl, No Slp, Irr, AR(1)	Irr: 0.708 (1.000) Lvl: 0.352 (0.497) AR(1): 7.385 (10.431)	Lvl: 2.135 [3.046]** AR(1): 0.563	0.867	8.226	3.74 {0.59}

See Note to Panel B.

Panel D: Kalman Filter estimation: Latin American equity flows

Country	Model	Component	Est. s.d. of disturb. and Q-ratios	Est. coef. of final state vector [t-values]	R^2	AIC	LB(9,q)
Argentina	4	Stc Lvl, No Slp, Irr, AR(1)	Irr: 9.825 (1.000) Lvl: 1.413 (0.144) AR(1): 5.289 (0.538)	Lvl: 1.161 [3.088]** AR(1): 0.296	0.893	4.957	5.33 {0.38}
Brazil	1	Stc Lvl, No Slp, Irr	Irr: 55.225 (1.000) Lvl: 20.977 (0.380)	Lvl: 5.290 [9.497]**	0.858	5.415	10.47 {0.11}
Chile	3	Stc Lvl, Stc Slp, Irr	Irr: 11.624 (1.000) Lvl: 0.456 (0.039) Slp: 0.781 (0.067)	Lvl: 2.366 [3.668]** Slp: 0.431 [2.355]*	0.801	5.378	3.92 {0.69}
Colombia	1	Stc Lvl, No Slp, Irr	Irr: 1.232 (1.000) Lvl: 0.283 (0.229)	Lvl: 1.049 [2.021]*	0.831	0.473	4.66 {0.59}
Ecuador	1	Stc Lvl, No Slp, Irr	Irr: 0.413 (1.000) Lvl: 0.157 (0.381)	Lvl: 3.228 [6.824]*	0.923	-1.675	4.29 {0.64}
Jamaica	1	Stc Lvl, No Slp, Irr	Irr: 3.677 (1.000) Lvl: 0.746 (0.203)	Lvl: 1.821 [2.881]**	0.846	2.696	0.09 {0.99}
Mexico	1	Stc Lvl, No Slp, Irr	Irr: 75.446 (1.000) Lvl: 21.349 (0.283)	Lvl: 1.339 [11.746]**	0.892	1.184	6.98 {0.32}
Uruguay	1	Stc Lvl, No Slp, Irr	Irr: 3.096 (1.000) Lvl: 0.411 (0.133)	Lvl: 2.410 [2.209]*	0.905	2.445	9.20 {0.16}
Venezuela	4	Stc Lvl, No Slp, Irr, AR(1)	Irr: 5.957 (1.000) Lvl: 2.585 (0.434) AR(1): 11.114 (1.866)	Lvl: 4.367 [3.704]** AR(1): 0.167	0.902	5.420	5.07 {0.41}

See Note to Panel B.

Panel E: Kalman filter estimation: Latin American bond flows

Country	Model	Component	Est. s.d. of disturb. and Q-ratios	Est. coef. of final state vector [t-values]	R^2	AIC	LB(9,q)
Argentina	2	Stc Lvl, Fxd Slp, Irr	Irr: 21.044 (1.000) Lvl: 5.362 (0.255)	Lvl: 6.783 [6.579]** Slp: 0.874 [3.812]**	0.808	4.420	7.02 {0.32}
Brazil	3	Stc Lvl, Stc Slp, Irr	Irr: 40.568 (1.000) Lvl: 4.976 (0.122) Slp: 3.501 (0.086)	Lvl: 7.666 [8.069]** Slp: 0.896 [2.621]*	0.804	4.463	6.78 {0.34}
Chile	5	Stc Lvl, Dmp Slp, Irr	Irr: 9.083 (1.000) Lvl: 0.873 (0.096) Slp: 0.345 (0.038)	Lvl: 5.921 [4.922]** Slp: 0.643 [2.541]*	0.863	4.539	8.99 {0.17}
Colombia	1	Stc Lvl, No Slp, Irr	Irr: 2.981 (1.000) Lvl: 0.056 (0.019)	Lvl: 1.975 [4.312]**	0.915	2.256	1.39 {0.97}
Ecuador	4	Stc Lvl, No Slp, Irr, AR(1)	Irr: 0.312 (1.000) Lvl: 0.569 (1.824) AR(1): 2.268 (7.269)	Lvl: 6.306 [4.281]** AR(1): 0.437	0.902	2.043	0.84 {0.97}
Jamaica	2	Stc Lvl, Fxd Slp, Irr	Irr: 0.487 (1.000) Lvl: 0.061 (0.127)	Lvl: 8.487 [4.653]** Slp: 0.773 [6.590]**	0.799	-1.239	11.56 {0.07}
Mexico	3	Stc Lvl, Stc Slp, Irr	Irr: 68.200 (1.000) Lvl: 10.114 (0.148) Slp: 4.182 (0.061)	Lvl: 2.749 [7.700]** Slp: 0.455 [2.593]*	0.853	3.129	4.58 {0.60}
Uruguay	2	Stc Lvl, Fxd Slp, Irr	Irr: 16.343 (1.000) Lvl: 5.921 (0.362)	Lvl: 8.423 [8.639]** Slp: 0.909 [3.636]**	0.800	4.081	11.72 {0.07}
Venezuela	2	Stc Lvl, Fxd Slp, Irr	Irr: 11.959 (1.000) Lvl: 5.412 (0.453)	Lvl: 7.392 [10.112]** Slp: 0.495 [3.235]**	0.846	4.484	9.86 {0.13}

See Note to Panel B.

CHAPTER 4

The Behaviour of the G5 Real Exchange Rates During the Post-Bretton Woods Period

'Among the puzzles which have emerged concerning the exchange rate over the last twenty years are the following. Real exchange rates are hard to distinguish from random walks .... The behavior of real exchange rates under the float may in fact be similar to behavior in other periods and the relative incidence of different types of shocks also largely the same, but researchers may not be able to discriminate, since, with less than twenty annual data points, the recent float may simply contain too few low-frequency components for researchers to detect mean-reverting real exchange rate behavior using conventional statistical tests.'

(Lothian and Taylor, 1995a, pp. 1-2)

4.1 Introduction

This study provides a number of insights into multivariate tests for mean reversion in real exchange rates. In particular, we both extend the development of and illustrate a potential pitfall in the interpretation of multivariate tests of long-run purchasing power parity which have been used by a number of researchers. In addition, we examine the finite-sample performance of an alternative multivariate test which avoids this pitfall and which is widely available to researchers. Further, we apply the suggested test procedures to provide some new evidence on long-run purchasing power parity among the G5 countries during the recent float.

The purchasing power parity (PPP) hypothesis states that national price levels expressed in a common currency should be equal. Equivalently, PPP implies that movements in the nominal exchange rate should be proportional to the ratio of national price levels or that the real exchange rate should be constant. PPP has variously been viewed as a theory of exchange rate determination, as a short-run or long-run equilibrium condition, and as an efficient arbitrage condition in either goods or asset markets (Officer, 1982; Dornbusch, 1987; Taylor, 1995; Froot and Rogoff, 1995; Rogoff, 1996).

The modern articulation of PPP starts in the immediate post World War I, after the collapse of the world financial system and as an attempt to restore a stable international financial environment. In fact, from the end of the nineteenth century until 1914, the international financial system was governed by the gold standard, a monetary arrangement among the major world countries according to which each participating country was committed to guarantee the free convertibility of the domestic currency into gold at a fixed price. Nevertheless, despite the commitment and the seriousness of the intentions of the member countries' monetary authorities, the collapse of the gold standard could not be prevented after the World War I and the fixed parities against gold could not be maintained especially because of the threat of a speculative attack led by investors expecting future devaluations of some members' currencies attempting to gain seignorage revenues. Also, another factor leading to the abandonment of the gold standard is identified by some researchers in the fact that no country had a strong leadership position in the international financial system during the interwar period: the US had not taken the leading role yet, while the UK was losing its leadership (eg. Kindleberger, 1973). At that time, re-establishing the exchange rates prevailing before the war was not a feasible solution, given the very different inflation and government finances positions experienced by the countries participating to the gold standard after the war. Hence, when 'the intensity of the crisis was so great that the conviction arose that a competitive international monetary system would always lead to extreme disorder' (De Grauwe, 1990), Cassel (1921, 1922) suggested the use of PPP as benchmark to reset exchange rates to new relative gold parities[1]. The very influential and debated proposal of Cassel, who may be considered to have written almost

[1] Officer (1982) and Dornbusch (1987) provide fascinating excursus on the origins of PPP, which was first postulated by some Spanish economists at the University of Salamanca in the sixteenth century.

everything one would wish to know about PPP, was simply to compute cumulative inflation rates on the basis of consumer price indices since the beginning of the war and then calculate the changes in the exchange rates of each country consistent with PPP on the basis of the cross-country inflation differentials[2].

The profession has radically switched position several times with regard to the validity of PPP. During the pre-Bretton Woods period, the professional consensus was quite supportive of PPP as a proposition holding over long periods of time (Friedman and Schwartz, 1963; Gaillot, 1970). Quoting Friedman and Schwartz (1963, pp. 678-679): 'One striking example of basic economic relations is the stability of relative prices in the United States and Great Britain adjusted for changes in the exchange rate between the dollar and the pound.... In the 79 years from 1871 to 1949, vast changes occurred in the economic structure and development of the United States, the place of Britain in the world economy, the internal monetary structures of both the Unites States and Great Britain, and the international monetary arrangements linking them. Yet despite these changes, despite two world wars and despite the statistical errors in the price-index numbers, the adjusted ratio on the base that makes 1929=100 was between 84 and 111 in all but one of the 79 years.'

A much stronger position was, however, taken in the early 1970s by some proponents of the monetary approach to exchange rate determination, who posited continuous, rather than long-run, purchasing power parity (eg. McCloskey and

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[2] While Cassel (1921, 1922) was influential in explicitly introducing the concept of PPP into modern discussions of exchange rate equilibrium, however, versions of the PPP hypothesis are implicit in the work of classical economists at least as far back as Hume (1752) and Ricardo (1821).

Zecher, 1976; Frenkel, 1976; Frenkel and Johnson, 1978).

A radical change in the position of the profession on the behaviour of real exchange rates occurred in the mid to late 1970s. In fact, following the very high variability of real exchange rates after the collapse of the Bretton Woods system, the extreme position of continuous PPP was largely abandoned (eg. Krugman, 1978; Frenkel, 1981). In addition, the relevant literature of the 1980s could not reject the hypothesis of random walk behaviour in real exchange rates (eg. Roll, 1979; Adler and Lehmann, 1983; Piggott and Sweeney, 1985), consistently with the failure to find cointegration between nominal exchange rates and relative prices (eg. Taylor, 1988; Corbae and Ouliaris, 1988; Enders, 1988; Mark, 1990), further reducing professional confidence in purchasing power parity and leading to the widespread belief that PPP was of little or no use empirically (eg. Dornbusch, 1988).

At the same time, however, some authors have reported results supporting long-run purchasing power parity for certain historical periods such as the interwar period (Taylor and McMahon, 1988), the International Gold Standard (Diebold, Husted and Rush, 1991) or the 1950s Canadian float (Choudhry, McNown and Wallace, 1991) or under special circumstances such as high inflation episodes (McNown and Wallace, 1989) or among the exchange rates of member countries of the European Monetary System (Cheung and Lai, 1995), thereby creating a puzzle as to why PPP failed to hold for the major exchange rates during the recent float.

A number of authors have argued that the data period for the recent float alone may simply be too short to provide any reasonable degree of test power in the normal statistical tests for non-stationarity of the real exchange rate (Frankel, 1989;

Lothian and Taylor, 1995; Hakkio and Rush, 1991). Accordingly, researchers have sought to remedy this by increasing the sample period under investigation (eg. Frankel, 1986, 1989; Edison, 1987; Abuaf and Jorion, 1990; Kim, 1990; Lothian, 1990; Hakkio and Joines, 1990; Diebold, Husted and Rush, 1991[3]; Lothian and Taylor, 1996). As noted by Frankel and Rose (1996) and others, however, the long samples required to generate a reasonable level of statistical power with standard nonstationarity tests may be unavailable for many currencies and may potentially be inappropriate because of regime changes. While some authors, notably Lothian and Taylor (1996), have argued that reliable inferences can be drawn - at least concerning the stability of the *first* moments of real exchange rate series - by extending the data across exchange rate regimes, others remain sceptical of this view. A number of authors, including Baxter and Stockman (1989), Mussa (1986), Frankel (1989) and Hegwood and Papell (1996), argue that the statistical properties of the real exchange rate appear to vary strongly across nominal exchange rate regimes[4]. To settle the issue of whether the real exchange rate has behaved in a mean-reverting fashion over the post-Bretton Woods period, therefore, would seem to require inference based on data for the recent float alone.

A second approach has therefore been taken by some researchers, involving the use of panel data on exchange rates over relatively shorter periods of time. Flood and Taylor (1996), for example, analyse a panel of annual data on 21

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[3] Diebold, Husted and Rush (1990) apply fractional integration techniques - see also Cheung and Lai (1993a).

[4] Relatedly, Grilli and Kaminksy (1991) argue that real exchange rate behaviour varies more with the historical period *per se* than across exchange rate regimes.

industrialised countries over the floating rate period and find strong support for mean reversion towards long-run purchasing power parity by regressing five-, ten- and twenty-year average exchange rate movements on average inflation differentials against the US. Frankel and Rose (1996) analyse a very large panel of annual data on 150 countries in the post World War II period and also find evidence of mean reversion similar to that evident in studies of long time series. In an influential study, Abuaf and Jorion (1990), develop a multivariate unit root test based on systems estimation of autoregressive processes for a set of real exchange rate series, and use this to reject the joint null hypothesis of non-stationarity of a number of real exchange rates for the recent floating rate period. Panel data methods have also been applied to this issue by, *inter alios*, Wei and Parsley (1995), Wu (1996), Oh (1996), O'Connell (1996) and Papell (1996).

In the present study, we seek to contribute to this literature in a number of ways. Firstly, we provide some further evidence on panel unit root tests of this kind, by calculating the finite-sample empirical distribution of a multivariate augmented Dickey-Fuller (MADF) statistic while allowing for higher-order serial correlation in real exchange rates and relaxing the assumption that the sum of the autoregressive coefficients are identical across the panel under the alternative hypothesis.

Secondly, however, we point out and illustrate through Monte Carlo simulations an important potential pitfall in the interpretation of multivariate unit root tests of this kind. The pitfall is simply this: the null hypothesis in panel unit root tests is usually that *all* of the series under consideration are realizations of unit root processes. Thus, the null hypothesis will be violated even if only *one* of the real exchange rate series in the panel is in fact stationary. Hence, although such multivariate tests may be informative under certain conditions, they may also be relatively uninformative - rejection of the null hypothesis will in general not help the researcher in determining how many of the series under consideration are stationary. We show, *inter alia*, that multivariate unit root tests of this kind may lead to a very high probability of rejection of the joint null hypothesis of non-stationarity when there is a single stationary process among a system of otherwise unit root processes, even when the root of the single stationary process is close to the unit circle.

Thirdly, therefore, we investigate by Monte Carlo methods the finite-sample empirical performance of a multivariate test in which the null hypothesis is that *at least one* of the series in the panel is a realization of a unit root process. This null hypothesis is only violated if *all* of the series are in fact realizations of stationary processes[5]. Moreover, the test procedure we suggest is now widely available to researchers since it simply involves a special application of Johansen's (1988) maximum likelihood procedure for testing for the number of cointegrating vectors in a system. A further attractive property of this test which we demonstrate is that, in the special case we examine - ie. under the null hypothesis that at least one of the series is a realization of a unit root process - it has a known limiting $\chi^2(1)$ distribution. We compute finite-sample critical values for this test but we also show

[5] Given, that is, the maintained hypothesis - common to all test procedures of this kind - that the series are realizations of either I(1) or I(0) processes (where an I(d) process can be thought of as one which must be differenced d times before it becomes stationary). Evidence of explosive behaviour in real exchange rates - for example positive Dickey-Fuller statistics - can easily be checked for prior to applying multivariate test procedures.

that the finite-sample empirical distribution is quite close to the asymptotic distribution in sample sizes exceeding about one hundred, corresponding approximately to the number of quarterly observations currently available for the recent float[6].

The remainder of the chapter is set out as follows. In Section 4.2 we briefly outline the PPP hypothesis and the long-run properties of real exchange rates which it implies. In Section 4.3 we outline two multivariate unit root tests based on a generalisation of the augmented Dickey-Fuller test statistic and of the Johansen maximum likelihood cointegration procedure respectively. In Section 4.4 we discuss some preliminary data analysis and univariate unit root tests on four dollar real exchange rates over the floating rate period. In Section 4.5 we report Monte Carlo evidence on the two multivariate tests described in Section 4.3. In Section 4.6 we employ these tests on quarterly real exchange rate data for the G5 countries. In Section 4.7 we report further empirical and Monte Carlo work based on real exchange rates among the same countries constructed using price indices containing a smaller proportion of non-tradables. A final section summarises and concludes.

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[6] One can distinguish between panel data studies in which the number of time series is relatively large (eg. Frankel and Rose, 1996) and those where the number of series is relatively small (eg. Abuaf and Jorion, 1990; Jorion and Sweeney, 1996). Formally, if N is the number of time series in the panel and T is the sample size, different econometric results follow according to whether T is assumed fixed and N is assumed to be relatively large, or whether N is assumed fixed and T is assumed to be relatively large, or whether both N and T are assumed to be large (which would normally require an additional assumption such as that N/T is small) - see Im, Pesaran and Shin (1997). The conceptual issues raised in this study relate to tests of long-run PPP based on both small and large panels, ie. tests involving any of the standard assumptions regarding N and T. The particular tests we investigate, however, are applicable primarily to relatively small systems of real exchange rates, such as is the case commonly encountered in testing for long-run PPP for a group of industrialised countries' exchange rates over the recent floating rate period.

4.2 Purchasing power parity

The PPP hypothesis states that the nominal exchange rate is proportional to a ratio of foreign and domestic price levels:

$$s_t = p_t - p_t^* \tag{4.1}$$

where s_t denotes the logarithm of the nominal exchange rate (domestic price of foreign currency) observed at time t and p_t and p_t^* are the logarithms of the domestic and foreign price levels respectively.

PPP may be examined through the real exchange rate since the logarithm of the real exchange rate, q_t, can be defined as the deviation from PPP:

$$q_t \equiv s_t + p_t^* - p_t \tag{4.2}$$

While allowing q_t to be non-zero in the short-run, a necessary condition for PPP to hold in the long run is that the real exchange rate q_t be stationary over time, not driven by permanent shocks. If this is not the case, then the nominal exchange rate and the price differential will permanently tend to deviate from one another. This is the rationale for applying non-stationarity tests to real exchange rate data as a means of testing for long-run purchasing power parity.

It should be noted that there are good economic reasons why real exchange rate movements should contain permanent components, particularly where the price indices used in the construction of the real rate contain both tradables and non-tradables.

The well known Harrod-Balassa-Samuelson effect (Harrod, 1933; Balassa, 1964; Samuelson, 1964), for example, implies that relatively fast growing countries may have a tendency to have higher real exchange rates based on relative consumer price indices (see eg. Froot and Rogoff, 1995; Rogoff, 1996; Obstfeld and Rogoff, 1996; Taylor and Sarno, 1997). The underlying argument of the Harrod-Balassa-Samuelson effect is that consumer price indices in fast-growing (rich) countries rise more rapidly than consumer price indices in slow-growing (poor) countries, once prices are expressed into the same currency using the prevailing exchange rate. The crucial underlying assumption is that technological progress is faster in the traded goods sector than in the nontraded goods sector, perhaps because the nontraded goods sector is relatively more service intensive[7]. Then, since rich countries are more productive *especially* in the traded goods sector relatively to poor countries, the price levels tend to be higher in wealthy countries. The mechanism is that if a positive productivity shock occurs in the traded sector, wages tend to rise in the whole economy. The rise in the wages in the traded goods sector can only be met by firms in the nontraded goods sector by increasing prices. Consequently, the overall level of prices rises. As noted by Rogoff (1996), however, while there is reasonably strong evidence supporting the Harrod-Balassa-Samuelson effect for exchange rates between very rich and very poor countries, its empirical relevance for the long-horizon time-series behaviour of real exchange rates among industrialised countries remains a matter of debate, possibly because of the effects

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[7] This argument was first advanced and empirically tested by Baumol and Bowen (1966).

of technology diffusion[8].

Besides the issue of traded-goods productivity bias, however, other arguments may be adduced at a theoretical level to suggest why real exchange rates may have persistent components. For example, differences in aggregate growth rates across countries may induce permanent changes in real exchange rates through preferences if consumers' Engle curves bend towards non-traded goods. Obstfeld and Rogoff (1995b) show that in the presence of sticky goods prices, monetary shocks may have a long-run effect on the real exchange rate because of the residual effects of temporary current account imbalances occasioned by short-run movements in the real exchange rate.

As noted by Lothian and Taylor (1996), it is probably true to say that few economists would rule out the possibility of real long-run effects on real exchange rates altogether, so that in testing for long-run PPP we are, therefore, implicitly testing whether permanent real effects account for only a relatively small part of long-run real exchange rate movements. To that extent, long-run PPP becomes an empirical matter[9].

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[8] Equilibrium models of the exchange rate (Stockman, 1980; Lucas, 1982), in which the real exchange rate is driven primarily by persistent real shocks such as shifts in tastes and technology, have also been used to rationalise persistence in real exchange rate movements. On the other hand, one might expect that *relative* real shocks affecting the real exchange rate between industrialised countries may have a higher mean-reverting component because of technology diffusion and other catch-up effects. Moreover, empirical tests of the implications of such models - for example that the real exchange rate should be invariant to nominal exchange rate regimes - has not by large been favourable. See Taylor (1995) for further discussion of these issues.

[9] The possibility that a very small unit root component may be dominated by a short-run stationary component of the real exchange rate is explicitly considered by Engel (1996). This possibility raises the issue of distinguishing between features of the processes under investigation which are economically

4.3 Multivariate unit root tests

In this section we outline two multivariate unit root tests. The first may be considered as an extension of previous work due to Abuaf and Jorion (1990). The second test we propose is an application of Johansen's maximum likelihood procedure for testing for the number of cointegrating vectors (Johansen, 1988, 1991) in the unusual case where the number of cointegrating vectors tested for is exactly equal to the number of time series in the system.

4.3.1 A multivariate augmented Dickey-Fuller test

Following the work of Fuller (1976) and Dickey and Fuller (1979, 1981), it is possible to test for a unit root in the stochastic process generating a time series q_t by estimating the auxiliary regression:

$$q_t = \mu + \sum_{j=1}^{k} \rho_j q_{t-j} + u_t \tag{4.3}$$

where the number of lags k is chosen such that the residual u_t is approximately white noise. For stationarity we require $\Sigma_{j=1}^{k}\rho_j < 1$, while if q_t is a realization of a unit root process, one should expect to find $\Sigma_{j=1}^{k}\rho_j = 1$. The augmented Dickey-Fuller test statistic is the standard "t-ratio" test statistic for $H_0{:}\Sigma_{j=1}^{k}\rho_j = 1$ and the rejection region consists of large negative values. As is well known, this statistic

important and those which are statistically important. The implicit assumption in this chapter, as in virtually all of the related literature, is that a non-stationary component which is statistically unimportant (in the sense that the unit root hypothesis is rejected) is also economically unimportant. Investigating the validity of this assumption provides an avenue for future research.

does not, however, follow the standard Student's t-distribution under the null hypothesis because of the theoretically infinite variance of q_t and finite-sample critical values have been computed using Monte Carlo methods by Fuller (1976) and MacKinnon (1991). This statistic is normally termed the Dickey-Fuller (DF) statistic for $k=0$ and the augmented Dickey-Fuller statistic (ADF) for $k>0$.

The first multivariate test for unit roots we propose is a multivariate analogue of the standard, single-equation augmented Dickey-Fuller test - the multivariate ADF or MADF test. Consider an (Nx1) dimensional stochastic vector process generated in discrete time according to:

$$q_{it} = \mu_i + \sum_{j=1}^{k} \rho_{ij} q_{it-j} + u_{it} \tag{4.4}$$

for $i=1,\ldots,N$ and $t=1,\ldots,T$, where N denotes the number of series in the panel and T is the number of observations. The disturbances $u_t=(u_{1t} \ldots u_{Nt})'$ are assumed to be independently normally distributed with a possibly non-scalar covariance matrix:

$$u_t \sim IN(0,\Lambda) \tag{4.5}$$

The standard, single-equation ADF unit root test would involve estimating each of the N equations separately and carrying out N individual tests of the null hypothesis:

$$H_{0i}: \sum_{j=1}^{k} \rho_{ij} - 1 = 0 \tag{4.6}$$

For situations where the root of each of the individual autoregressive process is close to but less than unity, it is well known that univariate ADF tests may lack power.

The approach taken in this study is to estimate (4.4) as a system of N equations, taking account of contemporaneous correlations among the disturbances[10], and to test (4.6) jointly on all N equations:

$$H_0: \sum_{j=1}^{k} \rho_{ij} - 1 = 0, \quad \forall i = 1, \ldots, N \tag{4.7}$$

taking the resulting Wald statistic as the MADF statistic.

The obvious way to estimate (4.4) jointly is to employ Zellner's (1962) "seemingly unrelated" (SUR) estimator, which is basically multivariate generalised least squares (GLS) using an estimate of the contemporaneous covariance matrix of the disturbances obtained from individual ordinary least squares estimation. We can write (4.4) in matrix notation as:

[10] O'Connell (1996) demonstrates the importance of accounting for cross-sectional dependence among real exchange rates when testing for long-run PPP. He shows that failure to allow for contemporaneous correlation of the residuals may generate very large size distortion in panel unit root tests.

$$Q = Z\beta + u \tag{4.8}$$

where the NTx1 vector $Q = (q_1' \; q_2' \; \ldots \; q_N')'$, q_i is a Tx1 vector of observations on the i-th real exchange rate, with t-th element q_{it}; Z is an NTxN(k+1) block diagonal matrix with the i-th block a Tx(k+1) matrix with ones in the first column and T observations on k lags of q_{it} in the remainder of the matrix; β is an N(k+1)x1 vector of stacked parameters for each equation; u is an NTx1 vector containing the stacked disturbances, so that

$$u \sim N(O, \Lambda \otimes I_T) \tag{4.9}$$

where I_k denotes - here and throughout this chapter - the kxk identity matrix (k=T,N). The restrictions in the null hypothesis (4.7) may then be written as:

$$\Psi\beta - \iota = O \tag{4.10}$$

where Ψ is an NxN(k+1) block-diagonal matrix with the i-th block a 1x(k+1) row vector with zero as the first element and unity elsewhere, ι is an Nx1 vector of ones and O is an Nx1 vector of zeroes. The MADF test statistic for the unit root hypothesis (4.7) is the standard Wald test statistic which may be written:

$$MADF=\frac{(\iota-\Psi\hat{\beta})'\{\Psi[Z'(\hat{\Lambda}^{-1}\otimes I)Z]^{-1}\Psi'\}(\iota-\Psi\hat{\beta})N(T-k-1)}{(X-Z\hat{\beta})'(\hat{\Lambda}^{-1}\otimes I)(X-Z\hat{\beta})} \quad (4.11)$$

where $\hat{\beta}$ and $\hat{\Lambda}$ are consistent estimates of β and Λ [11]. In general, the Wald statistic for testing N restrictions has a limiting χ^2 distribution with N degrees of freedom under the null hypothesis being tested. In the present case, however, its distribution is unknown because of the theoretically infinite variance of q_t under the null hypothesis (4.7). Its finite-sample empirical distribution can, however, be calculated by Monte Carlo simulation.

Abuaf and Jorion (1990) suggest a similar multivariate test, based on estimation of a first-order autoregressive equation for each individual real exchange rate, with the first-order autocorrelation coefficient constrained to be equal across exchange rates. Their proposed test statistic is then the ratio of the estimated common parameter minus one to its estimated standard error. The multivariate ADF statistic we propose can thus be viewed as a generalisation of the Abuaf-Jorion approach, to allow for higher order serial correlation in real exchange rates and to allow the sum of the autoregressive coefficients to vary across exchange rates under

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[11] Λ was estimated as the covariance matrix of the residuals obtained from OLS applied individually to each equation. Given a non-diagonal contemporaneous residual covariance matrix, the SUR estimator will be a more efficient estimator of β than OLS and so the finite-sample performance of the MADF should be better using SUR rather than individual OLS estimates. Note that the gain in efficiency and hence in test power should not be contingent upon assuming that the speed of adjustment is the same for all real exchange rates (Nelson, 1976).

the alternative hypothesis[12].

4.3.2 The Johansen likelihood ratio test

Johansen (1988, 1991) suggests a maximum likelihood procedure for testing for the number of cointegrating vectors in a multivariate context. Engle and Granger (1987) demonstrate that, among a system of N I(1) series, there can be at most N-1 cointegrating vectors. Hence, the only way there can be N distinct cointegrating vectors among N series is if each of the series is I(0) and so is itself a cointegrating relationship[13]. Thus, a test of the null hypothesis that there are less than N cointegrating vectors (which includes the possibility of zero cointegrating vectors) against the alternative that there are N is a test that each of the series is stationary.

The Johansen likelihood ratio (JLR) test for cointegration is based on the rank of a long-run multiplier matrix in a vector autoregressive system. Consider the data generation process of an Nx1 vector process Q_t - which may be assumed

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[12] Recent papers by Papell (1996) and O'Connell (1996) also allow for higher order serial correlation in tests of this kind, and the Monte Carlo work of Papell (1996) shows that the presence of serial correlation may affect the size of panel unit root tests. These authors do, however, retain the restriction that the autoregressive coefficients are identical across the panel. O'Connell (1996) shows that the generalised least squares estimator of the autoregressive parameters is invariant to the choice of the numeraire so long as they are constrained to be equal across the panel. Engel, Hendrickson and Rogers (1996) point out that allowing the autoregressive coefficients to vary across the panels implies that the form of the ARMA process may alter when the numeraire currency is changed, which is a general property of the addition or subtraction of ARMA processes (Granger and Morris, 1976), although this implies no inconsistency in the context of the present study.

[13] Again, as in Johansen (1988, 1991) and Johansen and Juselius (1990), this implicitly assumes a maintained hypothesis that the series are either I(0) or I(1).

to generate realizations of N real exchange rates at time t - in a vector autoregressive (VAR) form:

$$Q_t = \Pi_1 Q_{t-1} + \ldots + \Pi_k Q_{t-k} + \mu + \omega_t, \quad t=1,2,\ldots,T \quad (4.12)$$

where the terms in front of the Π_is are (NxN) matrices of parameters, μ is an Nx1 vector of constants and ω_t is an Nx1 vector of white noise errors. This VAR system can be reparameterised into the error correction form:

$$\Delta Q_t = \Gamma_1 \Delta Q_{t-1} + \ldots + \Gamma_{k-1} \Delta Q_{t-k+1} + \Gamma_k Q_{t-k} + \mu + \omega_t \quad (4.13)$$

where:

$$\Gamma_i = -I + \Pi_1 + \ldots + \Pi_i, \qquad i=1,\ldots,k \qquad (4.14)$$

and Γ_k represents the long-run solution of the VAR. Indeed, Γ_k is an NxN matrix whose rank defines the number of distinct cointegrating vectors. To see this, note that Γ_k may be written:

$$\Gamma_k = -\alpha\gamma' \qquad (4.15)$$

where α and γ are each Nx[rank(Γ_k)] matrices. γ can be interpreted as the matrix of cointegrating parameters and α as the matrix of error correction coefficients (Johansen, 1988). If, for example, each of the series is individually I(1) and no cointegrating vectors exist, then since all of the other terms in (4.13) are I(0), Γ_k must be the null matrix so that rank(Γ_k)=0. To take the opposite extreme, if Γ_k is of full rank then the space spanned by γ is N-dimensional Euclidean space and contains I_N. Hence, N cointegrating vectors can be formed, each consisting of just one of the series, which implies that each of the series must be I(0) and the vector process X_t is stationary (eg. Johansen and Juselius, 1990).

The rank of a matrix is equal to the number of non-zero latent roots. In the present context, stationarity of all of the processes in the vector autoregressive system is tantamount to Γ_k having full rank and so N non-zero latent roots. Thus, the null hypothesis of one or more non-stationary processes making up an Nx1 vector process can be expressed:

$$H_0: \ rank(\Gamma_k) < N \tag{4.16}$$

and tested against the alternative hypothesis that each of the series is stationary, or equivalently:

$$H_1: \ rank(\Gamma_k) = N \tag{4.17}$$

Since full rank of Γ_k would imply that all of the latent roots are non-zero, a test of (4.16) can be based only on the smallest latent root, since rejection of the hypothesis that the smallest latent root is zero is sufficient to reject the hypothesis that Γ_k has less than full rank.

Following Johansen (1988, 1991), an appropriate test statistic for the null hypothesis that the smallest latent root of Γ_k is zero, equivalent to a likelihood ratio test of (4.16) against (4.17) can be constructed as follows. First, correct for the effects of $\Im_t=\{\iota, \Delta Q_{t-1}, \Delta Q_{t-2}, \ldots, \Delta Q_{t-k+1}\}$ on ΔQ_t and Q_{t-k} (where ι is the unit vector) by projecting each of them onto $\Im_t$ and retrieving the residuals, denoted R_{ot} and R_{kt} respectively. Then form the matrices $S_{ij}=T^{-1}\Sigma_{t=1}^{T}R_{it}R_{jt}'$ (i,j=0,k), and extract the smallest root, λ_N say, of the characteristic equation $|\lambda S_{kk}-S_{ko}S_{oo}^{-1}S_{ok}|=0$. Johansen's likelihood ratio statistic is then[14]:

$$JLR = -T\log(1 - \lambda_N) \tag{4.18}$$

JLR as defined in (4.18) is a likelihood ratio statistic for one restriction ($\lambda_N=0$). Johansen and Juselius (1990) show that JLR converges weakly to a function of Brownian motion:

[14] The Johansen procedure can be related to the estimated Γ_k in the following way. The roots of $|\lambda S_{kk} - S_{ko}S_{oo}^{-1}S_{ok}|=0$ are, of course, the same as the roots of $|\lambda - S_{ko}S_{oo}^{-1}S_{ok}S_{kk}^{-1}|=0$. Nevertheless, by the Frisch-Waugh-Lovell (Frisch and Waugh, 1933; Lovell, 1963) theorem, $S_{ok}S_{kk}^{-1}$ is the multivariate least squares estimator of Γ_k in (4.13). By the same token, $S_{ko}S_{oo}^{-1}$ is the estimated matrix of coefficients of ΔQ_t resulting from applying least squares to (4.13) with ΔQ_t and Q_{t-k} interchanged.

$$JLR \Rightarrow \frac{\left[\int_0^1 (t-1/2)\, dB\right]^2}{\int_0^1 (t-1/2)^2 dt} \tag{4.19}$$

where B is a standard Brownian motion. Although, in general, the likelihood ratio statistics derived by Johansen (1988, 1991) and Johansen and Juselius (1990) have non-standard limiting distributions, the right hand side of (4.19) is in fact distributed as $\chi^2(1)$, so that, in this special case, the Johansen likelihood ratio statistic has a standard χ^2 distribution with one degree of freedom in large samples (Johansen and Juselius, 1990)[15].

4.3.3 *The MADF and the JLR tests compared*

There is a subtle but very important difference between the null hypothesis tested by the MADF statistic - (4.11) - and that tested by the JLR statistic - (4.18). The null hypothesis (4.7) will be violated if one or more of the series in question is a realization of an I(0) process. The null hypothesis (4.16) will, however, only be violated if *all* of the N series are realizations of I(0) processes. This implies that multivariate Dickey-Fuller tests should be interpreted with caution in the context of testing for non-stationarity of real exchange rates. Abuaf and Jorion (1990), for

[15] This result arises for two reasons. First, we are testing for the significance of only one latent root. Second, because we placed no restrictions on the constant intercept terms in the VAR (4.13), we implicitly allowed for the possibility of linear trends (Johansen and Juselius, 1990, p. 171). This is therefore a special case of the result due to West (1988), that if a linear trend is present under the null hypothesis of non-stationarity, then the usual asymptotics hold for the likelihood ratio test.

example, apply a restricted form of the MADF test to a system of ten real exchange rates and reject the null hypothesis of joint non-stationarity at the five percent level: while this may imply ten stationary real exchange rates, it may equally imply only one or two[16]. This suggests that the JLR test statistic may provide a useful alternative or complement to the MADF or similar panel unit root tests.

Below, we investigate the power of the MADF and JLR statistics under a variety of assumptions concerning the number of non-stationary series in the system under consideration.

4.4 Preliminary data analysis and single-equation unit root tests

Quarterly data on bilateral real dollar exchange rates among the G5 countries (ie. sterling-dollar, mark-dollar, franc-dollar and yen-dollar) for the period 1973Q1-1996Q2 were constructed from series obtained from the International Monetary Fund's International Financial Statistics (IFS) data bank[17]. Initially, we examined two real exchange rate series. These correspond to the nominal exchange rate (currency per dollar) deflated by, respectively, relative consumer price indices (CPI) and relative GDP deflators. In Section 4.7 below we analyse real exchange

[16] Papell (1996) finds that rejection of join non-stationarity is more likely with larger panels of real exchange rates. While this may be due to the increased power from exploiting more cross-sectional variability, it may also be due to the increased likelihood of finding one or more stationary processes as the panel size is increased.

[17] In investigating the low-frequency characteristics of time series processes, Shiller and Perron (1985) note that the span of the data set - in terms of years - is far more important than the number of observations *per se*. An intuitive discussion of this point is given in Davidson and MacKinnon (1993, Chapter 20). Our choice of quarterly data should therefore make the present analysis of wider interest to other researchers (since quarterly data are available for a wider range of countries than are monthly data) without a major loss of power of the tests.

rates for these countries constructed using producer price indices. All of the price indices are based on 1990. Each of the real exchange rate series was put into natural logarithms before the econometric analysis.

The first task was to estimate univariate autoregressive equations for each of the series and to construct single-equation unit root tests. In every case, a first-order autoregressive process appeared unsatisfactory in that significant serial correlation remained in the residuals. *A priori*, one might expect a fourth-order autoregression to be more suitable for quarterly data, and, in fact, this turned out to be the case in terms of eliminating serial correlation of the residuals. A fourth-order autoregressive model was also preferred on the basis of the Akaike Information Criterion (AIC) (Akaike, 1973), the Schwartz Information Criterion (SIC) (Schwartz, 1978) or the method proposed by Campbell and Perron (1991, p. 155)[18] [19].

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[18] In their Monte Carlo simulations, Cheung and Lai (1993b) show that for autoregressive processes with no moving average dependencies the Akaike Information Criterion and the Schwartz Information Criterion indicate the right lag order of a vector autoregression used for testing for cointegration in 99.86 percent and 99.96 percent of cases respectively.

[19] The AIC and SIC criteria for a univariate model are calculated as:

$$AIC = T\log(RSS) + 2g$$
$$SIC = T\log(RSS) + g\log(T)$$

where RSS denotes the residual sum of squares, g is the number of parameters estimated and T is the number of observations. The Campbell-Perron method involves starting with a high-order autoregression and sequentially excluding the highest-order lag until the coefficient on the highest-order lag is statistically significant.

In selecting the appropriate lag length for a vector autoregression of four autoregressive real exchange rates, we also employed the multivariate version of the AIC and the SIC:

Table 4.1 (Panel A) lists the estimated coefficients of the fourth-order autoregressions as well as test statistics for serial correlation in the residuals and ADF test statistics for each of the real exchange rates. In all cases, we were unable to reject, at the five percent level, the null hypothesis of non-stationarity on the basis of the single-equation ADF test statistics. Taken alone, this evidence would suggest - in keeping with the literature on univariate tests for mean reversion in the major real exchange rates over the period since 1973 - that real exchange rates for the G5 countries in the post Bretton Woods period are all realizations of stochastic processes integrated of order one. In economic terms, this would mean that real exchange rates show no tendency to settle down in the long run and that long-run purchasing power parity is violated over the recent floating-rate period.

In Table 4.1 (Panel B) we report the standard deviations of the residuals from estimating single-equation AR(4) models for the real exchange rates, together with the contemporaneous cross-exchange rate residual correlation matrix. The likelihood ratio statistics for the null hypothesis that the residual covariance matrices

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$$AIC = T\log|\Sigma| + 2G$$
$$SIC = T\log|\Sigma| + G\log(T)$$

where Σ is the determinant of the variance-covariance matrix and G is the total number of parameters estimated in all equations. In the context of a vector autoregression, the Campbell-Perron method consists of testing the joint exclusion restriction of the last lag included in each autoregression using a likelihood ratio test.

In addition, we also computed the AIC and the SIC for a system of four seemingly unrelated autoregressive processes. Also, in this case, the Campbell and Perron method is employed by estimating seemingly unrelated p-th order autoregressive processes for the four currencies' exchange rates and testing for the joint statistical significance of the p-th lagged dependent variable included in the regressions.

Overall, the results clearly indicate an optimal lag length of four.

are diagonal [*LR(diag)*] massively reject the null hypothesis, suggesting that systems estimation should yield substantial efficiency gains and that panel unit root tests applied to this data without allowing for this cross-sectional dependence would very likely be subject to substantial size distortion (O'Connell, 1996).

4.5 Monte Carlo simulations

The Monte Carlo experiments were based on a data generation process consisting of one to four autoregressive models, each of order four. From the descriptive statistics reported in Table 4.1 (Panel B) we derive the average contemporaneous covariance matrix for the AR(4) residuals, averaged across the CPI and GDP deflator adjusted real exchange rate residual covariance matrices (Table 4.1, Panel C). We employ this average covariance matrix in executing the Monte Carlo simulations. For each autoregressive model we took the average parameters for the estimates of the AR(4) process using the two real exchange rate series for each of the countries, as reported in Table 4.1, adjusting the first two slope parameters equally in order to adjust their sum[20]. Specifically, the systems

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[20] It might be argued that a separate data generating process should be formulated for each data set (ie. CPI and GDP deflator adjusted real exchange rates), rather than averaging across the covariance matrix or across the estimated autoregressive parameters. While the approach adopted in this study has the advantage of making the results of the Monte Carlo simulations more readily comprehensible by reducing the quantity of results to assimilate, it is clearly important to check for the generality of the results. In order to check the robustness of our Monte Carlo results to slight changes in the assumed data generating process, we therefore performed a number of safeguards. First, we carried out Monte Carlo simulations with data generating processes corresponding more specifically to each of the two data sets, for experiments corresponding to at least one cell in every column reported in Tables 4.3 - 4.8. In every case, the results were qualitatively unaffected, affecting at most the second decimal place of the rejection frequencies. In addition, we report in Table 4.2 how the empirical critical values for the MADF are affected by quite wide adjustments to the assumed

were based on the following data generating process:

$$q_{1t} = -0.016 + (1.107+\delta_1)q_{1t-1} + (-0.188+\delta_1)q_{1t-2} + 0.072q_{1t-3} - 0.090q_{1t-4} + u_{1t} \tag{4.20}$$

$$q_{2t} = -0.005 + (1.241+\delta_2)q_{2t-1} + (-0.359+\delta_2)q_{2t-2} + 0.274q_{2t-3} - 0.224q_{2t-4} + u_{2t} \tag{4.21}$$

$$q_{3t} = -0.010 + (1.318+\delta_3)q_{3t-1} + (-0.480+\delta_3)q_{3t-2} + 0.210q_{3t-3} - 0.120q_{3t-4} + u_{3t} \tag{4.22}$$

$$q_{4t} = -0.017 + (1.317+\delta_4)q_{4t-1} + (-0.462+\delta_4)q_{4t-2} + 0.246q_{4t-3} - 0.144q_{4t-4} + u_{4t} \tag{4.23}$$

where $(u_{1t}\ u_{2t}\ u_{3t}\ u_{4t})' \sim N(0,\Lambda)$ with Λ as given in Table 4.1 (Panel C), and the δ_i denote the relevant adjustments to the parameters. In order to generate the critical values for the test statistics, for example, we needed to generate the replications with the coefficients summing to unity in each case. Thus, we set $\delta_1=(1-0.901)/2$, $\delta_2=(1-0.932)/2$, $\delta_3=(1-0.928)/2$ and $\delta_4=(1-0.957)/2$. To take another example, in order to examine the behaviour of the statistics under the alternative hypothesis where all of the autoregressive processes had coefficients summing to 0.95, we set $\delta_1=(0.95-0.901)/2$, $\delta_2=(0.95-0.932)/2$, $\delta_3=(0.95-0.928)/2$ and $\delta_4=(0.95-0.957)/2$.

All of the Monte Carlo experiments discussed below were constructed using 5,000 replications in each experiment, with identical random numbers across

covariance matrix. In Section 4.7, we also report the results of a case study of real exchange rates constructed using producer price indices, the Monte Carlo simulations for which are based on a data generating process calibrated on those particular series.

experiments (Hendry, 1984). The simulations were executed for a number of different sample sizes (T=25, 50, 75, 100, 200, 300, 500). At each replication we started with the first four initial values of each of the artificial series set to zero. We then generated a sample size of 105+T (T=25, 50, 75, 100, 200, 300, 500) and discarded the first 105 observations, leaving a sample of size T for the analysis[21]. For the simulated I(0) process - ie. those with autoregressive coefficients summing to less than unity - this should reduce the dependence of the results on the initialization. By definition, however, the simulated I(1) processes are long-memory in nature and so are unavoidably contingent upon the initialization. An initialization of zero for the log real exchange rate does, however, seem reasonable.

4.5.1 The MADF test: Monte Carlo evidence

Table 4.2 reports the five percent empirical critical values for the MADF test with N=4 and the various sample sizes considered. In order to see how sensitive the results were to the assumed covariance matrix of the innovations, we also generated the critical values assuming two alternative covariance matrices: a half correlations matrix, in which the covariances between residual series were halved while keeping variances unchanged; and a diagonal covariance matrix, in which cross-correlations are all set to zero. The critical values obtained from experiments using the alternative covariance matrices (Table 4.2), however, do not differ widely from the critical values generated from the simulations employing the

[21] A sample size of 25 was not used for analysis of the JLR statistic because of the very low degrees of freedom in a fourth-order VAR with four lags with only 25 observations. Sample sizes of 300 and 500 were only used for analysis of the MADF test because of the high computational expense of Monte Carlo analysis with samples of this size with the Johansen procedure.

full historical matrix.

Table 4.3 reports the percentage of rejections of the null hypothesis of joint non-stationarity when each of the series is a realization of a stationary process, for various true values of the sum of the autoregressive coefficients, under a multivariate system with four time series using the full historical covariance matrix. As Table 4.3 displays, the MADF test is quite powerful even when the sum of the autoregressive parameters is very close to unity. With a sample size of 100, for example, the null hypothesis of non-stationarity is rejected in around 30 percent of the replications when each of the series is generated by a process with a root of 0.99. For roots of 0.975, the rejection rate rises to around 60 percent, to close to 90 percent for roots of 0.95, nearly 99 percent for roots of 0.925 and close to 100 percent for roots of 0.9 or less.

For purposes of comparison, we also calculated the empirical power function for a univariate ADF test applied to the first equation in the system (assuming an AR(4) process), using five percent critical values calculated from the response surface results given in MacKinnon (1991). The results are reported in Table 4.4. This exercise demonstrated, however, that the MADF test is very much more powerful than the univariate ADF test. For a sample size of 100 and a single root of 0.990, for example, the ADF statistic has a rejection frequency of just a little over six percent, compared to nearly 30 percent for the MADF statistic when there are four processes with the same root (Table 4.3). For roots of 0.95 the MADF statistic has a rejection frequency of nearly 90 percent, compared with just over 17 percent for the univariate ADF.

In Table 4.5 we report the power of the MADF test to reject the null hypothesis of joint non-stationarity of the processes when, respectively, one, two or three of them have a root less than unity. The MADF test remains quite powerful under these circumstances. For example, when two of the processes are I(1) and two are stationary (Panel B), the rejection frequency is around 23 percent at the five percent significance level and a sample size of 100 when the roots of the stationary processes are each equal to 0.99, rising quickly to over 45 percent for roots of 0.975 and of over 80 percent for roots of 0.95. [22]

Table 4.5 illustrates, however, the potential pitfall in the use - or rather the interpretation - of the MADF or similar panel unit root tests in testing for long-run PPP: rejection of the null hypothesis does not necessarily indicate that all of the processes in the system under consideration are stationary. For N=4, T=100, with three unit root processes in the system together and *with just one stationary process* with a root of 0.95, for example, the null hypothesis is rejected at the five percent level in 65 percent of the replications (Panel C). For one stationary process with a root of 0.9 together with three unit root processes and a sample size of 100, the rejection frequency is in excess of 95 percent. For larger sample sizes, the rejection frequencies when there is only a single stationary process are even greater.

Following Abuaf and Jorion (1990), we also investigated the possibility that the small-sample behaviour of the MADF test may be affected, in terms of both critical values and power, by the presence of autoregressive conditionally

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[22] We also performed Monte Carlo experiments for the MADF with smaller systems - ie. with N=2 and N=3. For N=3 we omitted (4.23) from the data generating process, and for N=2 we omitted (4.22) and (4.23) (as well as the corresponding rows and columns of Λ in each case). The major noteworthy characteristic of these experiments was a noticeable drop in power as the number of series in the system is reduced.

heteroskedastic (ARCH) effects and generalised ARCH (GARCH) in the disturbances. As part of our preliminary data analysis, we tested for ARCH effects in the residuals of the AR(4) real exchange rate processes. In Table 4.1 (Panel D) we report the estimated ARCH(1) parameters using CPI- and GDP deflator-adjusted real exchange rate data, in addition to the average of the estimated ARCH parameters from the two real exchange rates series. In general the estimated ARCH parameters are found to be very small and not statistically significantly different from zero at conventional nominal levels of significance. Thus, we investigate, using Monte Carlo simulations, the effect on the MADF test statistic of an ARCH(1) process of the form:

$$u_{it+1}|t \sim N(0,h_{it+1}), \ h_{it+1} = \eta_0 + \eta_1 u_{it}^2 \tag{4.24}$$

as well as a GARCH(1,1) process of the form:

$$u_{it+1}|t \sim N(0,h_{it+1}), \ h_{it+1} = \eta_0 + \eta_1 u_{it}^2 + \zeta h_{it} \tag{4.25}$$

In executing the Monte Carlo simulations, we use the averaged ARCH parameters given in Table 4.1 (Panel D). Table 4.6 reports the five percent critical values and the estimated power function obtained from executing 5,000 Monte Carlo simulations. Using an ARCH(1) and a GARCH(1,1) specification based on the estimated paramaters obtained from the preliminary data analysis (Table 4.1, Panel D), we found that the power of the MADF tests was little affected. More precisely, in Table 4.6 (Panel A), we report the critical values for the MADF with N=4 for three different cases: (a) the error terms are generated by ARCH(1) processes set on the basis of the estimated parameters reported in Table 4.1 (Panel D); (b) the error terms are generated by ARCH(1) processes set using the sample estimates given in Table 4.1 (Panel D) for η_0 and with η_1 arbitrarily set to 0.6; (c) the error terms are generated by GARCH(1,1) processes using the sample estimates given in Table 4.1 (Panel D) for η_0 and η_1 and setting ζ to 0.6. While the power of the MADF tests was - in general - little affected, as the parameters were arbitrarily increased, however, both the power and the actual test size increased slightly. Overall, the results reported in Table 4.6, indicating that the MADF test statistic does not appear too sensitive to the presence of conditional heteroskedasticity, are in line with the results reported by Abuaf and Jorion (1990).

4.5.2 The JLR test: Monte Carlo evidence

As we showed above, the JLR statistic does in fact have a limiting $\chi^2(1)$ distribution under the null hypothesis of less than full rank of the long-run matrix. Nevertheless, we generated its small-sample empirical distribution, since Cheung and Lai (1993b) show that there may be substantial finite-sample bias toward rejection of the null hypothesis in Johansen likelihood ratio statistics.

In Table 4.7, therefore, we report the critical values, obtained from executing 5,000 simulations, for the JLR statistic under the null hypothesis that the rank of the long-run matrix is less than full. As noted above, less than full rank of this matrix corresponds to the case of one or more non-stationary processes in the system. Accordingly, the critical values were calculated under all possible cases which satisfy the null hypothesis - ie. for all values of the number of non-stationary processes from four down to one - and the arithmetic average taken as the appropriate entry of Table 4.7. For the smaller samples, the estimated finite-sample critical values are quite large, compared to the corresponding critical values from the $\chi^2(1)$ distribution. For the larger sample sizes of 100 or more, however, the critical values are quite close to those of the $\chi^2(1)$ distribution. For a sample size of T=100, for example, the 5 percent critical value for a system with N=4 is 4.0686. This compares with the five percent critical value form the $\chi^2(1)$ distribution of 3.84 and the large-sample critical value for JLR estimated by Johansen and Juselius (1990, Table A1) of 3.962. This suggests that in sample sizes of one hundred or more - corresponding roughly to the number of quarterly observations available for the floating rate period - researchers could assume that JLR follows a $\chi^2(1)$ distribution under the null hypothesis with only slight size distortion[23].

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[23] Reinsel and Ahn (1988) suggest a finite-sample scaling factor adjustment of T/(T-Nk) to the asymptotic critical values of Johansen test statistics in order to obtain their finite-sample counterparts. Although the Monte Carlo study of Cheung and Lai (1993b) suggests that the Reinsel-Ahn adjustment does not yield unbiased estimates of the finite sample critical values, we thought it worth examining this hypothesis in the present situation where the limiting distribution is a known $\chi^2(1)$ distribution. Accordingly, following Cheung and Lai (1993b), we fitted by ordinary least squares a response surface of the form

$$\frac{\zeta(Tj)}{\zeta[\chi^2(1)]} = \alpha_0 + \alpha_1 \frac{T}{(T-Nk)} + errors$$

where $\zeta(Tj)$ is the finite-sample five percent level critical value for the j-th experiment and $\zeta[\chi^2(1)]$ is the corresponding critical value from the $\chi^2(1)$ distribution. This yielded:

In Table 4.8 we report the estimated power function of the JLR test when all of the processes are integrated of the same order, using the finite-sample five percent critical value. Note that these rejection frequencies are not directly comparable to those discussed above for the MADF test because the null hypotheses for the two tests are quite different. Nevertheless, the JLR test does appear to be moderately powerful. For N=4, T=100 and at a significance level of five percent, for example, the rejection frequency is around 12 percent for roots of 0.99, rising to just over 16 percent for roots of 0.975, to just under 25 percent for roots of 0.95, to about 38 percent for roots of 0.925 and to 54 percent for roots of 0.9.[24]

$$\frac{\zeta(Tj)}{\zeta[\chi^2(1)]} = \underset{(0.335)}{-0.149} + \underset{(0.282)}{1.096}\frac{T}{(T-Nk)}$$

$$R^2=0.95 \quad DW=2.21 \quad LM(1)=\underset{[0.59]}{0.29} \quad RST(1)=\underset{[0.67]}{0.17} \quad JB(2)=\underset{[0.66]}{0.84}$$

$$BP(1)=\underset{[0.70]}{0.14} \quad ARCH(1)=\underset{[0.74]}{0.11} \quad \chi^2(\alpha_0=0,\alpha_1=1)=\underset{[0.54]}{1.23}$$

where figures in parenthesis below estimated coefficients are estimated standard errors and in brackets below test statistics are p-values. R^2 is the coefficient of determination, DW is the Durbin-Watson statistic, LM(1) is a Lagrange multiplier test statistic for first-order serial correlation, RST is Ramsey's (1969) test statistic for functional form misspecification, JB is the Jarque-Bera (1980) test for normality of the residuals, BP is the Breusch-Pagan (1979) test for heteroskedasticity, ARCH is a test for first order autoregressive conditional heteroskedasticity, and the final test statistic relates to a linear Wald test that the intercept is zero and the slope coefficient is unity. Higher-order powers of the regressor were found to be insignificant. These results suggest that, in this particular case, the Reinsel-Ahn adjustment may provide a reasonable approximation to the finite-sample critical values.

[24] This compares to rejection frequencies calculated for a univariate ADF statistic of around six percent, ten percent, 17 percent, 28 percent and 42 percent respectively.

4.6 Empirical results for CPI-adjusted and GDP deflator-adjusted real exchange rates

In Table 4.9 we report the results of applying the MADF and JLR tests to four real dollar exchange rates - dollar-sterling, dollar-mark, dollar-franc and dollar-yen - using quarterly data for the period 1973Q1 through 1996Q2. We used lag lengths of four for each of the autoregressions (in the construction of the MADF test statistic) and in the vector autoregression (in the construction of the JLR test statistic).

The MADF test (Panel A) rejects the null hypothesis of joint non-stationarity at the five percent significance level for both types of real exchange rates (ie. deflated either by relative CPIs or relative GDP deflators), thereby implying that at least one of the series in each of the systems is a realization of a stationary process.

Applying the JLR test (Panel B), we easily reject the null hypothesis that the long-run impact matrix has less than full rank when we consider the real exchange rates deflated by relative CPIs, implying that all the series in question are realizations of stationary processes.

When we consider the real exchange rates constructed using relative GDP deflators, however, we are not able to reject the null hypothesis at the five percent level. In both cases, the JLR test results are qualitatively unaffected whether we use the finite-sample critical values given in Table 4.7, the relevant percentiles of the $\chi^2(1)$ distribution, or the asymptotic critical values calculated by Johansen and Juselius (1990, Table A1).

Taken together, therefore, the MADF and JLR test results imply the following. For the CPI-adjusted real exchange rates, we reject the hypothesis that each of the four series is generated by an I(1) process (Panel A). For the same real exchange rates, we can also reject the null hypothesis that at least one of them is generated by a non-stationary process (Panel B). Hence, the strong implication is that they are each realizations of stationary processes over the floating exchange rate period.

For the GDP deflator-adjusted real exchange rates, however, while we can reject at the five percent level the hypothesis that each of the four series is generated by an I(1) process (Panel A), we are unable to reject, at the same significance level, the null hypothesis that at least one of them is generated by a non-stationary process (Panel B). The JLR test thus indicates the need for caution in interpreting the MADF test result applied to the GDP-adjusted real exchange rates: the most we can say is that *at least one* of them appears to be generated by a stationary process over the floating rate period.

The difference in the test results applied to the two sets of real exchange rates is perhaps not surprising since real exchange rates constructed using relative CPIs may be viewed as more appropriate for testing PPP than those constructed using relative GDP deflators, as GDP deflators will typically be constructed using a much larger proportion of non-tradable goods prices (Froot and Rogoff, 1995; Rogoff, 1996; Taylor and Sarno, 1997, Chapter 3).

4.7 A case study: producer price indeces

As a check on the robustness of the simulation results discussed above to the specific data generating processes assumed, we also investigated the mean-reverting behaviour of real exchange rates among among the G5 constructed using producer price indices (PPIs). Since PPIs cover a higher proportion of tradables goods prices than either CPIs or GDP deflators, one might expect long-run PPP to hold more strongly using these indices to construct the real exchange rate series[25]. Quarterly data on PPIs were obtained from the IFS data bank for the sample period 1980Q1-1996Q2 (66 data points), since the PPI series for France was only available from 1980i onwards.

As for the CPI-adjusted and the GDP deflator-adjusted real exchange rate series, a first-order autoregressive model appeared unsatisfactory for each of the real exchange rate series in that significant serial correlation remained in the residuals. Elimination of the residual serial correlation led to the choice of a fourth-order model in every case, and this lag length was also optimal according to the other selection criteria considered, namely the Akaike Information Criterion, the Schwartz Information Criterion and the Campbell-Perron method[26].

Table 4.10 (Panel A) lists the estimated coefficients as well as ADF test statistics and residual diagnostics. In all cases, we were unable to reject, at the five percent level, the null hypothesis of non-stationarity on the basis of the single-equation ADF test statistics.

The contemporaneous covariance matrix for the AR(4) residuals, given in Table 4.10 (Panel B), demonstrates strong cross-sectional effects in the data, which is confirmed formally on the basis of a likelihood ratio statistic for the diagonality

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[25] Keynes (1932) and McKinnon (1971) argue that the Harrod-Balassa-Samuelson effect is likely to be more pronounced in real exchange rates constructed using PPIs rather than CPIs.

[26] See footnote 19.

of the covariance matrix. Using this estimated covariance matrix and the estimated parameters adjusted to fit the null hypothesis of four unit roots, as described in Section 4.7, we then constructed the MADF and JLR five percent critical values for a sample size of 66 and 5,000 replications[27].

The five percent empirical critical value for the MADF test is 18.7894, which is in the range between the critical values, given in Table 4.2, for samples of T=50 (21.2993) and T=75 (18.5062) as generated from the average covariance matrix of CPI-adjusted and GDP deflator-adjusted real exchange rate residuals. In fact, using the data generation process described in Section 4.5, we estimated the five percent critical value for a sample size of T=66 at 18.9834.

The five percent critical value for the JLR statistic was estimated at 4.6856, which again is in the range between the critical values, given in Table 4.5, for samples of T=50 (5.5065) and T=75 (4.3133) as generated from the average covariance matrix given in Table 4.1 (Panel C). Using the data generation process described in Section 4.5, we estimated the JLR five percent critical value for a sample size of T=66 at 4.6834.

The MADF test statistic calculated on the actual PPI-adjusted real exchange rates for the period 1980Q1-1996Q2 is 19.7432 which, compared to the critical value of 18.7894, enables us to reject the null hypothesis of joint non-stationarity at the five percent level, implying that at least one of the PPI-adjusted real exchange rate series in the system is a realization of a stationary process.

The JLR test computed on the same system of PPI-adjusted real exchange

[27] Following our previous practice, we initialised the first four observations to zero, generated 105+66 observations and discarded the first 105.

rate series is, however, 4.0969 which, compared to the five percent critical value of 4.6632, does not enable us to reject the null hypothesis that the long-run impact matrix has less than full rank. In fact, the empirical marginal significance level of this value of the test statistic[28] is 6.64 percent, so that while we were unable to reject the null hypothesis at the five percent level, it could be rejected at the seven percent level. Given that we were able to reject the null hypothesis using the JLR test applied to the real exchange rate series constructed using consumer price indices, which would cover a higher proportion of non-tradables than the producer price indices, this suggests that the marginal inability to reject the null hypothesis at the five percent level using the PPI-adjusted real exchange rates may be due to a loss of power because of the smaller sample size[29].

4.8 Conclusions

Given that univariate tests for stationarity may lack power with sample periods corresponding to the span of the recent floating rate period, a number of authors have employed multivariate unit root tests to systems of real exchange rates in order to provide a more powerful test of the hypothesis that real exchange rates are realizations of non-stationary processes. In this study we have provided a number of insights into multivariate tests of long-run purchasing power parity.

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[28] ie. the percentage of experiments which generated a larger value of the test statistic under the null hypothesis.

[29] As note earlier, this problem would not be alleviated by increasing the frequency of the data, since it is the total span in terms of years that is available which is important for investigating long-run time-series behaviour (Shiller and Perron, 1985).

First, in an extension of the previous literature, we developed a multivariate augmented Dickey-Fuller (MADF) test for the null hypothesis of joint non-stationarity of the processes in question, to allow for higher-order serial correlation and for the roots of the autoregressive processes to differ across the panel under the alternative hypothesis. We generated the finite-sample empirical distributions of this statistic using Monte Carlo methods, and also investigated its power characteristics under various departures from the null hypothesis. Although the MADF test was shown to be quite powerful, we also demonstrated a potential pitfall in the use - or rather the interpretation - of multivariate ADF tests of this kind. Namely, that rejection of the null hypothesis of joint stationarity may be due to as few as one of the real exchange rate series in the system under investigation being generated by a stationary process. For a sample size of around 100, for example, the presence of a single stationary process in a system together with three unit root processes will lead to rejection, at the five percent level, of the joint null hypothesis of non-stationarity on about 65 percent of occasions when the root of the stationary process is as large as 0.95, and on more than 95 percent of occasions if the single stable root is 0.9 or less.

We therefore suggested an alternative or complementary multivariate test of non-stationarity where the null hypothesis is not that *all* of the series are generated by non-stationary processes, but, rather, that *at least one* of the series is generated by a non-stationary process. This null hypothesis will only be violated if *all* of the series in question are realizations of stationary processes. Moreover, the suggested test procedure is widely available to researchers since it is just a special case of Johansen's (1988) maximum likelihood ratio (JLR) for testing for the number of cointegrating vectors in a system, applied to the unusual case where the number of cointegration vectors tested for is equal to the number of series in the system under consideration. In addition, it turns out - again unusually - that this statistic has a limiting $\chi^2(1)$ distribution under the null hypothesis.

Again using Monte Carlo simulation techniques, we generated the finite sample empirical distributions of this test statistic, and showed that the $\chi^2(1)$ distribution - or indeed the large sample distribution tabulated by Johansen and Juselius (1990) - provides a reasonable approximation in samples larger than 100. We also demonstrated that the JLR test statistic has reasonable power characteristics for various departures from the null hypothesis.

We also applied both the MADF and JLR multivariate tests for non-stationarity to quarterly data on real dollar exchange rates constructed using consumer price indices and GDP deflators for the UK, Germany, France and Japan over the floating rate period 1973Q1-1996Q2, and to quarterly data on real dollar exchange rates for the same countries constructed using producer price indices for the period 1980Q1-1996Q2.

For real exchange rates constructed using relative consumer price indices, both tests strongly reject their respective null hypotheses at the five percent level, indicating mean reversion in these real exchange rates over the floating rate period.

For real exchange rates constructed using relative implicit GDP deflators, however, while the MADF statistic strongly rejects at the five percent level the hypothesis that all of the series are generated by non-stationary processes, the JLR test cannot reject at the five percent level the hypothesis that at least one of the series is generated by a non-stationary process. Taken together, therefore, the

MADF and JLR tests imply that some but probably not all of the GDP deflator-adjusted real exchange rate series are mean reverting. As noted by a number of authors, this is perhaps not surprising given the higher proportion of non-tradables prices covered by the implicit GDP deflator compared with the consumer price index.

For real exchange rates constructed using relative producer price indices, we were able to reject at the five percent level the null hypothesis that all of the series are realizations of I(1) processes using the MADF statistic, while the JLR statistic for the null hypothesis that at least one of the series was a realization of an I(1) process had a marginal significance level of 6.64 percent. Given the finding of mean reversion in the CPI-adjusted real exchange rates, which would be more prone to Harrod-Balassa-Samuelson effects than the PPI-adjusted real exchange rates because of the higher proportion of non-tradables prices covered, one interpretation is that the failure to establish mean-reversion of all the PPI-adjusted real exchange rates may be due to the loss in power resulting from the smaller available span of data.

In any case, the evidence that CPI-adjusted real exchange rates among the G5 are apparently mean reverting over the floating rate period is, by itself, an important finding of our research, corroborating other recently emerging evidence that long-run PPP may hold after all. Indeed, it seems that the profession's confidence in long-run PPP, having been low for a number of years, may itself be mean reverting[30].

[30] The fact that our MADF statistics are significant at the five percent level for all real exchange rate systems considered contrasts with the recent results reported by O'Connell (1996) and Engel, Hendrickson and Rogers (1996), who are

Appendix 4.1 The Engle-Granger procedure

According to the seminal work of Engle and Granger (1987), any two non-stationary series, which are found to be integrated of the same order, are cointegrated if a linear combination of the two exists which is integrated of lower order. If two series are cointegrated, then the non-stationarity of one series exactly offsets the non-stationarity of the other and a long-run relationship is established between the two variables. Generally, if two variables x_t and y_t both have a stationary, invertible, non-deterministic ARMA representation after differencing d times, ie. they are both integrated of order d or I(d), then the linear combination:

$$x_t + \alpha y_t = z_t \tag{A4.1.1}$$

will in general found to be I(d) as well. If, exceptionally, a cointegrating parameter α exists such that the residual in (A4.1.1) is integrated of order I(d-c), $c>0$, then the two variables are said to be cointegrated of order d,c, or CI(d,c). If this is the case, then a strong long-run relationship exists between the two variables considered, since they share a common stochastic trend and cointegration of a pair of variables is at least a necessary condition for them to have a stable long-run (linear) relationship (Stock and Watson, 1988).

A widely used econometric procedure to test for cointegration is the two-step procedure of Engle and Granger (1987), who formalise the idea of variables sharing an equilibrium relationship in terms of cointegration between time series. The preliminary step before proceeding to investigate cointegration is to test for

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unable to reject the unit root hypothesis using panel unit root tests. As noted by Engel, Hendrickson and Rogers (1996, p. 22), however, while these authors are unable to reject unit roots, their point estimates do suggest mean reversion with half lives similar to thos reported in the literature. They therefore suggest that their particular test may lack power. In contrast, O'Connell (1996) is careful to construct tests which have good power characteristics. The prime suspect for the differences in our results must therefore be data differences. While O'Connell uses quarterly IFS data for a comparable sample period to us, his panels cover available data for the world (64 countries), Europe (20 countries), Asia (13 countries), South America (13 countries) and Africa (13 countries). In contrast, we study only the G5 countries.

nonstationarity of the series considered (see Spanos 1976). The work of Fuller (1976) and Dickey and Fuller (1979, 1981) is crucial at this stage. In fact, it is possible to test for a unit root on a time series x_t by estimating the auxiliary regression:

$$\Delta x_t = \beta x_{t-1} + \sum_{i=1}^{n} \gamma_i \Delta x_{t-i} + \gamma_0 + u_t \tag{A4.1.2}$$

where the number of lags n is chosen such that u_t is approximately white noise. For stationarity we require $\beta<0$, while if x_t is a nonstationary series, we would find $\beta=0$. The test statistic is the standard "t-ratio" for the estimate of β and the rejection region consists of large negative values. Unfortunately, the t-ratio statistic for the estimate of β does not follow the standard Student's t-distribution under the null hypothesis, because of the theoretically infinite variance of x_t. Empirical critical values computed using Monte Carlo methods have to be considered. They are given by Fuller (1976) for the Dickey-Fuller statistic (where no lags are included in (A4.1.2)) and for the augmented Dickey-Fuller (ADF, so termed if lags are included in (A4.1.2)). Also Mackinnon (1991) lists the critical values for the DF and ADF statistics for a number of cases and degrees of freedom.

In order to fully understand the rationale behind this test statistic, consider the following (n+1)-th order representation of x_t:

$$x_t = \lambda_0 + \lambda_1 x_{t-1} + \lambda_2 x_{t-2} + \ldots + \lambda_{n+1} x_{t-n-1} + u_t \tag{A4.1.3}$$

Now reparameterise (A4.1.3) as:

$$\begin{aligned}\Delta x_t &= \lambda_0 + \left(\sum_{i=1}^{n+1} \lambda_i - 1\right) x_{t-1} - \left(\sum_{i=2}^{n+1} \lambda_i\right) \Delta x_{t-1} \\ &\quad - \left(\sum_{i=3}^{n+1} \lambda_i\right) \Delta x_{t-2} - \ldots - \lambda_{n+1} \Delta x_{t-n} + u_t\end{aligned} \tag{A4.1.4}$$

which is of the form (A4.1.2). The condition we require for stationarity is that the sum of the autoregressive parameters λ_i, $\Sigma\lambda_i$, is less than one, whilst $\Sigma\lambda_i$ is equal to one if the series x_t has a unit root. Thus, we require the coefficient of x_{t-1} in (A4.1.2) to be significantly negative for a unit root to be precluded.

If the null hypothesis of nonstationarity is rejected, then we cannot go any further. By contrast, if the null is not rejected, then it is correct and advisable to test for a unit root on the first difference of the series in question in order to exactly specify the order of integration[31]. Only at this stage the possibility of cointegration may arise. If a linear combination between two or more series is defined which makes the difference between them stationary, the variables in the estimated regression are said to be cointegrated and the error term in the regression will present well-defined first and second order moments[32]. Ordinary least squares (OLS) is therefore legitimate for estimating the cointegrating regression:

$$x_t = \delta + \alpha y_t + v_t \qquad \text{(A4.1.5)}$$

where v_t is the fitted residual. In fact, if cointegration characterises the estimated regression, the OLS estimator of the cointegrating parameter α converges to the true α as the number of observations increases more rapidly than it usually is for the OLS estimator under standard conditions. This is the so-called property of "super consistency". Formally, estimates of α are highly efficient with variances $0(T^{-2})$ compared to more usual situations where the variances are $0(T^{-1})$, T being the

[31] Dickey and Pantula (1988) suggest an extension of the basic procedure if more than one unit root is suspected. The methodology simply consists of performing (augmented) Dickey-Fuller tests on successive differences of the series in question, starting from higher-order unit roots and testing down until stationarity is found.

[32] This definition of cointegration is not incompatible with the one given previously. From the last definition given, the case is not considered that a set of I(2) variables may be cointegrated of order CI(2,1), so there exists a linear combination that is I(1). This case is, however, of no interest in economics. Most of the cointegration literature focuses on the case in which each variable contains a single unit root.

sample size (Stock, 1987). Offsetting the result of super consistency, however, there is another feature of cointegration analysis: small-sample bias is present in the OLS estimator of the cointegrating parameter and its limiting distribution is non-normal with a non-zero mean. Moreover, Banerejee, Dolado, Hendry and Smyth (1986) show that a high R^2 can provide an index of the value of the estimate. That is to say, $1-R^2$ has the same limiting distribution of the bias[33].

On the other side, if the no-cointegration hypothesis cannot be rejected, then the estimated regression is just a "spurious" one and has no economic meaning. If no bounded combination of the levels exists, then the error term in the regression must be non-stationary under the null hypothesis. Yule (1926) proves that the R^2 of a spurious regression tends to unity and therefore it is meaningless in that case. Granger and Newbold (1974) also point out that the least square estimator is not even consistent and the customary tests of statistical inference break down. They suggest using the Durbin-Watson (DW) statistic, which would tend to zero since the residuals would be nonstationary. A simple general rule suggested is that $R^2 > DW$ constitutes a serious indication that the estimated regression is spurious (Granger and Newbold, 1974)[34].

So, if we cannot reject the hypothesis that both x_t and y_t are I(1), then the second step of the procedure is to test the hypothesis of no-cointegration:

$$H_0 : v_t \sim I(1) \tag{A4.1.6}$$

and in practice we test for a unit root on the residual from the cointegrating

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[33] Note, however, that in multivariate models a high R^2 does not imply that each element in the cointegrating vector is estimated with neglegible bias. Therefore, the use of R^2 as a rough index of the small-sample bias is limited to simple bivariate models (Banerejee, Dolado, Hendry and Smyth, 1986).

[34] In the light of the likely event of incurring in spurious regression, time-series analysts following the traditional Box-Jenkins approach suggest differencing and prewhitening the series prior to proceeding to estimation. The limitation of this approach is, obviously, that it does not give due attention to long-run relationships between the levels of the series in question, to which economic theory is instead usually devoted.

regression (A4.1.3) in the same way as outlined for testing for a unit root on any time series. The critical values tabulated by Fuller (1976) cannot be used, however, when testing for non-stationarity in the residuals. It is clear even intuitively that, since OLS chooses the estimator which minimises the residual variance, we might expect to reject the hypothesis of non-stationarity too often. Appropriate critical values, computed by Engle and Granger (1987) by using Monte Carlo simulations, have to be used to test for a unit root on the cointegrating residuals[35] [36].

Nevertheless, the Engle and Granger two-step procedure is not a very powerful technique for testing for cointegration[37]. A relatively more powerful econometric technique for testing for cointegration is the Johansen procedure.

In addition, the Engle and Granger procedure implicitly assumes that there is a unique cointegrating vector, but this may well not be the case when the number of variables in the cointegrating regression is greater than two. In the latter case there is no guarantee that we have an estimate of a unique cointegrating vector. In general, given N I(1) variables, there may be up to (N-1) distinct cointegrating vectors (Engle and Granger, 1987).

The Johansen (1988) maximum likelihood estimator circumvents this problem and enables us to test for the presence of multiple cointegrating vectors. Moreover, as discussed in Section 4.3.2, Johansen also shows how to test for linear restrictions on the parameters of the cointegrating vectors, therefore making it

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[35] Another way of testing for non-stationarity on the OLS residuals from a cointegrating regression, suggested by Sargan and Barghava (1983), consists of testing the Durbin-Watson statistic against a value of zero (CRDW). Intuitively, since $CRDW \simeq 2(1-\rho)$, where ρ is the first-order autocorrelation coefficient, $CRDW \simeq 0$ when $\rho = 1$, ie. a unit root is present.

Also a more general alternative exists to the Dickey-Fuller tests for testing for a unit root on a time series: the Phillips-Perron test. Its main feature is that this test does not require the error term to be serially uncorrelated or homogeneous. Rather, the Phillips-Perron test allows the error terms to be weakly dependent and heterogeneously distributed (Perron, 1988).

[36] A useful reference is, once again, MacKinnon (1988), who reports the critical values not only for the DF and ADF statistics, but also for the CRDW statistics for a number of cases and degrees of freedom.

[37] By "powerful" we mean having a relatively high probability of rejecting the null hypothesis when it is false.

possible to test the theory by drawing statistical inferences on the relative magnitudes of the estimated coefficients.

Table 4.1 Preliminary data analysis

Panel A: Single-equation AR(4) estimates and ADF tests on bilateral dollar exchange rate

	Lag 1	Lag 2	Lag 3	Lag 4	Intercept	Adj. R^2	Q(27)	DW	Sum of AR Coeffs.	ADF
UK										
CPI	1.104	-0.207	0.096	-0.097	-0.014	0.86	31.21 (0.26)	2.016	0.896	-2.407
GDPD	1.110	-0.169	0.048	-0.083	-0.018	0.88	36.12 (0.11)	2.012	0.906	-2.392
Mean	1.107	-0.188	0.072	-0.090	-0.016				0.901	
Germ.										
CPI	1.247	-0.391	0.307	-0.232	-0.003	0.92	16.58 (0.94)	2.051	0.931	-2.198
GDPD	1.235	-0.327	0.241	-0.216	-0.007	0.93	17.15 (0.93)	2.052	0.933	-2.163
Mean	1.241	-0.359	0.274	-0.224	-0.005				0.932	
France										
CPI	1.328	-0.504	0.234	-0.129	-0.012	0.91	12.06 (0.99)	2.016	0.929	-2.096
GDPD	1.308	-0.456	0.186	-0.111	-0.008	0.91	11.84 (0.99)	2.012	0.927	-2.128
Mean	1.318	-0.480	0.210	-0.120	-0.010				0.928	
Japan										
CPI	1.330	-0.493	0.287	-0.163	-0.018	0.95	27.04 (0.46)	2.011	0.961	-1.678
GDPD	1.304	-0.431	0.205	-0.125	-0.016	0.94	22.96 (0.69)	2.011	0.953	-1.703
Mean	1.317	-0.462	0.246	-0.144	-0.017				0.957	

Notes: CPI and GDPD denote the real exchange rate deflated by relative consumer price indices and implicit GDP deflators respectively. Adj. R^2 denotes the degrees-of-freedom adjusted coefficient of determination; Q(27) is the Ljung-Box test statistic for residual serial correlation up to 27 lags (p-values in parentheses); DW denotes the Durbin-Watson test for first-order serial correlation; and ADF is the augmented Dickey-Fuller statistic (ie. the "t-ratio" for the slope coefficients to sum to unity).

Panel B: residual standard deviations and contemporaneous correlations

	UK	Germany	France	Japan
<u>CPI</u>				
Standard Deviations	0.052	0.046	0.044	0.049
Correlations:				
UK	1.000			
Germany	0.631	1.000		
France	0.639	0.907	1.000	
Japan	0.457	0.565	0.554	1.000
LR(diag) = 249.88				
<u>GDPD</u>				
Standard Deviation	0.054	0.049	0.047	0.051
Correlations:				
UK	1.000			
Germany	0.633	1.000		
France	0.644	0.900	1.000	
Japan	0.460	0.563	0.553	1.000
LR(diag) = 245.47				

Notes: Covariance and correlation matrices were constructed using the residuals from the AR(4) regressions reported in Panel A *LR(diag)* is a likelihood ratio statistic for the null hypothesis that the off-diagonal elements of the relevant covariance matrix are zero, and has a limiting χ^2 distribution with six degrees of freedom if the true covariance matrix is diagonal.

Panel C: Average covariance matrix

	UK	Germany	France	Japan
UK	0.283 E-2			
Germany	0.159 E-2	0.224 E-2		
France	0.111 E-2	0.120 E-2	0.208 E-2	
Japan	0.172 E-2	0.215 E-2	0.127 E-2	0.254 E-2

Notes: Entries correspond to the arithmetic average of the corresponding entries in the two covariance matrices for the residuals from the AR(4) regressions given in Panel B.

Panel D: Estimated ARCH parameters

	est. η_0	est. η_1
UK		
CPI	0.279E-2	0.177E-1
GDPD	0.307E-2	0.000
Mean	0.293E-2	0.885E-2
Germany		
CPI	0.215E-2	0.485E-2
GDPD	0.236E-2	0.187E-1
Mean	0.226E-2	0.118E-1
France		
CPI	0.197E-2	0.000
GDPD	0.216E-2	0.000
Mean	0.207E-2	0.000
Japan		
CPI	0.217E-2	0.000
GDPD	0.237E-2	0.000
Mean	0.227E-2	0.000

Notes: est. η_0 and est. η_1 are the estimated ARCH parameters. In fact, the error term in each autoregressive process was modelled as an ARCH(1) process of the form:

$$u_{it+1}|t \sim N(0, h_{it+1}), \quad h_{it+1} = \eta_0 + \eta_1 u_{it}^2$$

Table 4.2 MADF: empirical critical values at the 5 percent level

	T=25	T=50	T=75	T=100	T=200	T=300	T=500
FHM	30.7115	21.2993	18.5062	16.8701	14.5119	13.5531	12.2813
HCM	28.8331	19.3964	16.4575	15.2197	13.2503	12.4483	11.3285
DM	33.4695	22.1978	18.7771	17.2629	14.1732	12.9284	11.7198

Notes: The critical values correspond to a fourth-order system. FHM, HCM and DM stand for full historical covariance matrix, half-historical covariance matrix and diagonal covariance matrix respectively. T is the sample size.

Table 4.3 MADF: estimated power function

$\Sigma\rho$	T=25	T=50	T=75	T=100	T=200	T=300	T=500
0.990	6.53	11.45	18.71	29.57	70.96	92.18	99.94
	(0.35)	(0.45)	(0.55)	(0.64)	(0.64)	(0.38)	(0.03)
0.975	8.17	20.67	38.84	59.88	97.70	99.58	100.00
	(0.40)	(0.57)	(0.69)	(0.69)	(0.21)	(0.09)	(0.00)
0.950	10.46	33.25	66.38	89.39	100.00	100.00	100.00
	(0.43)	(0.67)	(0.67)	(0.43)	(0.00)	(0.00)	(0.00)
0.925	12.25	47.79	86.37	98.51	100.00	100.00	100.00
	(0.46)	(0.71)	(0.48)	(0.17)	(0.00)	(0.00)	(0.00)
0.900	13.57	62.17	96.37	99.90	100.00	100.00	100.00
	(0.48)	(0.68)	(0.26)	(0.04)	(0.00)	(0.00)	(0.00)

Notes: Entries give the percentage of rejections at the five percent level of the null hypothesis that the autoregressive coefficients sum to unity in each of the four equations, using the critical values (full historical covariance matrix) given in Table 4.2, computed from 5,000 replications. Figures in parentheses are the estimated standard errors. $\Sigma\rho$ is the sum of the autoregressive coefficients in each of the four equations in the data generating process, and T is the sample size.

Table 4.4 ADF: estimated power function

$\Sigma\rho$	T=25	T=50	T=75	T=100	T=200	T=300	T=500
0.990	3.86	4.52	5.16	6.28	12.84	19.46	36.88
	(0.27)	(0.29)	(0.31)	(0.34)	(0.47)	(0.56)	(0.68)
0.975	4.50	6.54	8.32	10.08	23.46	40.34	77.72
	(0.29)	(0.35)	(0.39)	(0.42)	(0.60)	(0.69)	(0.59)
0.950	5.60	8.34	12.32	17.34	51.26	84.22	99.68
	(0.32)	(0.39)	(0.46)	(0.53)	(0.71)	(0.51)	(0.08)
0.925	7.34	10.74	18.00	28.02	78.76	98.36	100.00
	(0.37)	(0.44)	(0.54)	(0.64)	(0.58)	(0.18)	(0.00)
0.900	7.78	13.80	25.89	42.24	93.34	99.92	100.00
	(0.38)	(0.49)	(0.62)	(0.77)	(0.34)	(0.04)	(0.00)

Notes: Entries give the percentage of rejections of the null hypothesis that the autoregressive coefficients sum to unity in a single AR(4) equation, using five percent level critical values calculated from MacKinnon (1991), computed from 5,000 replications. Figures in parentheses are the estimated standard errors. $\Sigma\rho$ is the sum of the autoregressive coefficients in the autoregressive equation and T is the sample size.

Table 4.5 MADF estimated power function with less than four unit root processes

Panel A: three stationary processes, one unit root process

$\Sigma\rho$	T=25	T=50	T=75	T=100	T=200	T=300	T=500
0.990	5.98 (0.33)	9.60 (0.42)	16.36 (0.52)	26.46 (0.62)	70.74 (0.64)	90.24 (0.42)	99.72 (0.07)
0.975	7.25 (0.37)	15.86 (0.52)	33.06 (0.66)	53.83 (0.70)	96.80 (0.25)	98.97 (0.14)	100.00 (0.00)
0.950	8.66 (0.40)	26.43 (0.62)	60.98 (0.69)	86.52 (0.48)	99.94 (0.03)	100.00 (0.00)	100.00 (0.00)
0.925	9.41 (0.41)	40.82 (0.69)	83.52 (0.52)	97.51 (0.22)	100.00 (0.00)	100.00 (0.00)	100.00 (0.00)
0.900	11.14 (0.44)	55.44 (0.70)	94.36 (0.33)	99.70 (0.08)	100.00 (0.00)	100.00 (0.00)	100.00 (0.00)

Panel B: two stationary processes, two unit root processes

$\Sigma\rho$	T=25	T=50	T=75	T=100	T=200	T=300	T=500
0.990	5.90 (0.33)	9.20 (0.41)	14.48 (0.50)	22.74 (0.59)	58.96 (0.69)	84.78 (0.51)	99.34 (0.11)
0.975	7.03 (0.36)	15.28 (0.51)	28.48 (0.64)	45.49 (0.70)	94.00 (0.33)	99.88 (0.05)	100.00 (0.00)
0.950	8.20 (0.39)	25.21 (0.61)	52.98 (0.70)	81.06 (0.55)	99.06 (0.14)	100.00 (0.00)	100.00 (0.00)
0.925	8.91 (0.40)	37.79 (0.68)	77.52 (0.59)	95.89 (0.28)	100.00 (0.00)	100.00 (0.00)	100.00 (0.00)
0.900	9.72 (0.42)	51.86 (0.71)	90.96 (0.40)	99.36 (0.11)	100.00 (0.00)	100.00 (0.00)	100.00 (0.00)

Panel C: one stationary process, three unit root processes

$\Sigma\rho$	T=25	T=50	T=75	T=100	T=200	T=300	T=500
0.990	5.44	7.01	10.02	15.28	46.06	73.33	95.90
	(0.32)	(0.36)	(0.42)	(0.51)	(0.70)	(0.62)	(0.28)
0.975	6.11	9.62	17.66	31.58	83.76	98.46	100.00
	(0.34)	(0.42)	(0.54)	(0.66)	(0.52)	(0.17)	(0.00)
0.950	6.32	15.51	36.80	64.89	98.46	100.00	100.00
	(0.34)	(0.51)	(0.68)	(0.67)	(0.17)	(0.00)	(0.00)
0.925	6.82	23.78	58.24	86.95	100.00	100.00	100.00
	(0.36)	(0.60)	(0.70)	(0.48)	(0.00)	(0.00)	(0.00)
0.900	7.65	33.00	76.42	95.78	100.00	100.00	100.00
	(0.37)	(0.66)	(0.60)	(0.28)	(0.00)	(0.00)	(0.00)

Notes: Entries give the percentage of rejections of the null hypothesis that the autoregressive coefficients sum to unity in each of the four equations, using the critical values (full historical covariance matrix) given in Table 4.2, computed from 5,000 replications. Figures in parentheses are the estimated standard errors. $\Sigma\rho$ is the sum of the autoregressive coefficients in each of i equations in the data generating process (the sum of the coefficients in the first 4-i equations is set to unity), with i=3 in Panel A, i=2 in Panel B and i=1 in Panel C. T is the sample size.

Table 4.6 MADF with ARCH and GARCH disturbances

Panel A: Empirical critical values

	T=25	T=50	T=75	T=100	T=200	T=300	T=500
ARCH1	32.9844	22.1175	19.5125	17.2653	14.8646	13.9033	12.5775
ARCH2	36.9487	25.4134	21.8299	19.6811	16.5323	15.2188	14.8223
GARCH	34.9293	24.5427	21.6089	20.1777	17.1882	15.3876	14.0479

Notes: The critical values correspond to a fourth-order system where the error term in each autoregressive process was modelled as an ARCH or a GARCH process. More precisely, ARCH1, ARCH2 and GARCH correspond to the three different cases where: (a) the error terms are generated by ARCH(1) processes set on the basis of the estimated parameters reported in Table 4.1 (Panel D); (b) the error terms are generated by ARCH(1) processes set using the sample estimates given in Table 4.1 (Panel D) for η_0 and with η_1 arbitrarily set to 0.6; (c) the error terms are generated by GARCH(1,1) processes using the sample estimates given in Table 4.1 (Panel D) for η_0 and η_1 and setting ζ to 0.6.

Panel B: Estimated power function (ARCH1)

$\Sigma\rho$	T=25	T=50	T=75	T=100	T=200	T=300	T=500
0.990	6.70	11.92	19.86	31.90	75.54	94.68	100.00
0.975	8.42	21.24	39.58	60.08	98.52	100.00	100.00
0.950	10.82	33.90	66.08	90.60	100.00	100.00	100.00
0.925	12.76	48.58	87.02	98.96	100.00	100.00	100.00
0.900	13.36	63.18	97.52	100.00	100.00	100.00	100.00

Notes: Entries give the percentage of rejections at the five percent level of the null hypothesis that the autoregressive coefficients sum to unity in each of the four equations, using the critical values (full historical covariance matrix) given in Panel A, computed from 5,000 replications. $\Sigma\rho$ is the sum of the autoregressive coefficients in each of the four equations in the data generating process, and T is the sample size. ARCH1 corresponds to the case where the error terms are generated by ARCH(1) processes set on the basis of the estimated parameters reported in Table 4.1 (Panel D).

Panel C: Estimated power function (ARCH2)

$\Sigma\rho$	T=25	T=50	T=75	T=100	T=200	T=300	T=500
0.990	7.10	13.72	22.34	33.62	79.12	97.66	100.00
0.975	9.20	22.50	42.54	63.16	99.98	100.00	100.00
0.950	11.86	36.62	69.16	94.72	100.00	100.00	100.00
0.925	14.22	50.20	90.12	100.00	100.00	100.00	100.00
0.900	15.14	68.70	99.60	100.00	100.00	100.00	100.00

Notes: Entries give the percentage of rejections at the five percent level of the null hypothesis that the autoregressive coefficients sum to unity in each of the four equations, using the critical values (full historical covariance matrix) given in Panel A, computed from 5,000 replications. $\Sigma\rho$ is the sum of the autoregressive coefficients in each of the four equations in the data generating process, and T is the sample size. ARCH2 corresponds to the case where the error terms are generated by ARCH(1) processes set using the sample estimates given in Table 4.1 (Panel D) for η_0 and with η_1 arbitrarily set to 0.6.

Panel D: Estimated power function (GARCH)

$\Sigma\rho$	T=25	T=50	T=75	T=100	T=200	T=300	T=500
0.990	7.00	11.86	20.78	32.60	77.42	95.12	100.00
0.975	8.92	21.02	40.84	63.06	98.60	100.00	100.00
0.950	11.22	34.12	67.98	92.82	100.00	100.00	100.00
0.925	13.64	49.62	89.14	98.68	100.00	100.00	100.00
0.900	14.12	68.70	98.40	100.00	100.00	100.00	100.00

Notes: Entries give the percentage of rejections at the five percent level of the null hypothesis that the autoregressive coefficients sum to unity in each of the four equations, using the critical values (full historical covariance matrix) given in Panel A, computed from 5,000 replications. $\Sigma\rho$ is the sum of the autoregressive coefficients in each of the four equations in the data generating process, and T is the sample size. GARCH corresponds to the case where the error terms are generated by GARCH(1,1) processes using the sample estimates given in Table 4.1 (Panel D) for η_0 and η_1 and setting the GARCH parameter ζ to 0.6.

Table 4.7 JLR: average empirical critical values

N	T=50	T=75	T=100	T=200
4	5.5065	4.3133	4.0686	3.9712

Notes: Using a fourth-order system, for each sample size, four experiments were performed in which, respectively all four, the first three, the first two, and the first of the four autoregressive equations in the data generating process had coefficients summing to unity while the remainder were set equal to the estimated values given in Table 4.1 (Panel A). Each experiment involved 5,000 replications and the critical value was taken as the 95th percentile. The average empirical critical value was then taken and is given in the table. T is the sample size.

Table 4.8 JLR: estimated power function with four processes integrated of the same order

$\Sigma\rho$	T=50	T=75	T=100	T=200
0.990	6.56 (0.35)	8.86 (0.40)	11.77 (0.45)	26.04 (0.62)
0.975	9.03 (0.40)	13.26 (0.48)	16.10 (0.52)	46.04 (0.70)
0.950	9.45 (0.41)	16.44 (0.52)	24.28 (0.61)	83.61(0.52)
0.925	9.55 (0.41)	21.46 (0.58)	38.39 (0.69)	97.18 (0.23)
0.900	9.84 (0.42)	28.24 (0.64)	53.96 (0.70)	99.49(0.10)

Notes: Entries give the percentage of rejections, at the five percent level, of the null hypothesis that the autoregressive coefficients sum to unity in each of the four equations, using the critical values given in Table 4.7, computed from 5,000 replications. Figures in parentheses are the estimated standard errors. $\Sigma\rho$ is the sum of the autoregressive coefficients in each of the four equations in the data generating process, and T is the sample size.

Table 4.9 Empirical results

Panel A: MADF test statistics

Countries	CPI	GDPD
UK, GE, FR, JA	26.5497	26.5774

Panel B: JLR test statistics

Countries	CPI	GDPD
UK, GE, FR, JA	5.8851	3.7712

Notes: In Panel A, the null hypothesis is that all four real exchange rate series are realizations of unit root processes, the alternative hypothesis is that at least one of them is a realization from a stationary process. The five percent critical value, taken from Table 4.2, is 16.8701.

In Panel B the null hypothesis is that at least one of the four real exchange rate series is a realization of a unit root process, the alternative hypothesis is that all of them are realizations of stationary processes. The five percent critical value, taken from Table 4.5, is 4.0686.

Table 4.10 Estimates using PPI-adjusted real exchange rates

Panel A: Single-equation AR(4) estimates and ADF tests on bilateral dollar exchange rate

	Lag 1	Lag 2	Lag 3	Lag 4	Intercept	Adj. R^2	Q(27)	DW	Sum of AR Coeffs.	ADF
UK	1.148	-0.282	0.132	-0.070	0.005	0.88	21.95 (0.58)	1.992	0.928	-1.619
Germ.	1.243	-0.279	0.095	-0.114	0.005	0.93	13.04 (0.96)	2.056	0.945	-1.568
France	1.291	-0.421	0.175	-0.115	0.014	0.91	12.57 (0.97)	2.040	0.930	-1.763
Japan	1.263	-0.422	0.297	-0.181	-0.008	0.94	7.99 (0.99)	1.986	0.957	-1.334

Notes: Adj. R^2 denotes the degrees-of-freedom adjusted coefficient of determination; Q(27) is the Ljung-Box test statistic for residual serial correlation up to 27 lags (p-values in parentheses); DW denotes the Durbin-Watson test for first-order serial correlation; and ADF is the augmented Dickey-Fuller statistic (ie. the "t-ratio" for the slope coefficients to sum to unity).

Panel B: Covariance Matrix

	UK	Germany	France	Japan
UK	0.264 E-2			
Germany	0.178 E-2	0.235 E-2		
France	0.167 E-2	0.216 E-2	0.226 E-2	
Japan	0.116 E-2	0.151 E-2	0.139 E-2	0.199 E-2
LR(*diag*)= 328.48				

Notes: Covariance and correlation matrices were constructed using the residuals from the AR(4) regressions reported in Panel A. *LR(diag)* is a likelihood ratio statistic for the null hypothesis that the off-diagonal elements of the covariance matrix are zero, and has a limiting χ^2 distribution with six degrees of freedom if the true covariance matrix is diagonal.

CONCLUSIONS

In this thesis we have investigated four different topics in macroeconomics and international finance which are currently under debate, receiving widespread attention by researchers. In this final section, we briefly summarise our key findings and suggest potential avenues for future research.

The first task accomplished in this thesis was an investigation of the effect of movements in the real interest rate on changes in real consumption over a sample period which is characterised by very significant regime changes in financial policies in the two countries examined, the UK and France. After estimating a number of alternative models for consumption - which are found to be quite unsatisfactory, especially using UK data - we proposed and estimated a nonlinear model for consumption which extends and significantly outperforms the linear consumption function due originally to Campbell and Mankiw (1989). Modelling the proportion of income going to liquidity constrained consumers, λ_t, as a non-linear function of a proxy for financial deregulation, results in an estimated model which cannot be rejected against an alternative error correction specification for both countries examined, yielding sensible and well-determined estimates of the parameters. Most tellingly, the implied time-varying and non-monotonic behaviour of λ_t accords well with the observed pace and pattern of financial deregulation in the two countries considered over the sample period, providing a feasible explanation of previous findings of researchers who detected evidence of a linear upward trend in λ_t for the UK and detected no significant change in the path of λ_t for France.

The empirical evidence provided here suggests, therefore, that imperfections in capital markets lead to variations in loan supply having a substantial impact on private consumption expenditure and may be responsible for the commonly reported rejection of the pure life-cycle permanent income hypothesis. Nevertheless, while the results from estimating our nonlinear model for consumption expenditure of nondurables and services should be taken with caution because of the use of aggregate data, they provide a clear and sensible explanation of the large errors in forecasting consumption - and overall economic developments - obtained in the late 1980s and early 1990s using macroeconometric models where the consumption functions do not adequately capture the effects of financial deregulation. Overall, in fact, the main interest of the findings provided here is for practical purposes, such as macro-modelling and forecasting. Also, for example, modelling explicitly the precautionary motive for saving in the estimated model as well as modelling the transition probability of consumers changing their status from liquidity-constrained to unconstrained or *vice versa* may further improve the performance of the model proposed in this thesis, hence representing immediate avenues for future research in this area.

As the removal of imperfections in capital markets and financial deregulation make - in the domestic economy - the consumption-smoothing paradigm implied by the life-cycle permanent income hypothesis more applicable relative to models allowing for liquidity-constrained consumers, the widespread reduction or removal of barriers to international capital movements and financial liberalization are expected to make - in the world economy - the intertemporal approach to the current account more fully applicable, ie. leaving current account imbalances to the role of residual, thanks to the enhanced increase in the degree of international capital mobility. In particular, the abolition of remaining capital controls by France and Italy in 1991 and the creation of a European Financial Area also raise important questions concerning the nature and the extent of capital mobility in Europe. The second issue investigated in this thesis involves, in fact, a range of issues concerning capital flows in Europe. In particular, we reexamined the validity of savings-investment correlations as a means to shed light on the degree of international capital mobility distinguishing between the short-run and the long-run saving-investment correlation coefficient. Using data for the UK, where in October 1979 exchange control was abolished ending a long period of restrictions on capital flows between the UK and the international economy, we provide strong evidence that the short-run savings-investment correlation is significantly higher than the long-run correlation, in the sense that the temporary components are relatively more highly correlated. This suggests that transitory movements in savings, given the frictions existing in international transactions, are more likely to remain in the UK, since it may not be worthwhile facing the cost of analysing foreign investment opportunities or evading or avoiding exchange controls. Conversely, when an increase in saving occurs which is perceived to be permanent, saving is more likely to flow abroad. Overall, the results suggest that the Feldstein-Horioka definition of international capital mobility, correctly interpreted, may provide a valid way to shed light on the degree of international capital mobility. In fact, the correlation, both between temporary (short-run) and permanent (long-run) components of the change in savings and investment shares are significantly lower in the period after the abolition of UK exchange controls. Moreover, the correlation between the permanent components in the post-1979 period is not significantly different from zero, suggesting a very high degree of capital mobility when the change in savings is perceived to be permanent.

Capital flows in a financially integrated area also provide a means of insuring individuals' incomes against changes in local productivity, since private investors can create a portfolio which is diversified in the returns to both human and financial capital. We also examined this issue and provide further evidence on the extent and effect of European capital mobility through an examination of GNP/GDP ratios across the largest European countries and a comparison of these with similar calculations executed for regions within the UK. The empirical evidence provided indicates that, as European financial integration continues, one may expect an increasingly diverse pattern of capital ownership across the EU to emerge. In turn, this may be expected to enhance higher intra-EU current account imbalances, generating a greater possibility of European output and income measures diverging and greater opportunities for individuals to insure themselves against region-specific risk. Further research on those issues is certainly warranted.

As an effect of the ongoing process of abolition of various sorts of impediments and capital controls and the broader liberalization of financial markets experienced also in developing countries from the late 1980s, a sharp expansion of capital flows together with a marked increase in the participation of foreign investors and foreign financial institutions in the financial markets of developing countries represents one of the main features of the recent development of international capital flows. An interesting feature of the recent trend of capital flows to developing countries is that private (bond and equity) as opposed to official capital flows are increasingly a crucial source of financing of large current account imbalances.

We investigated the determinants of US capital flows directed to some Latin American and Asian countries over the sample period 1988-92, in order to shed light on whether the large capital inflows in the form of bonds and equities received by the countries examined were induced by improvements in the economic performances of those countries - country-specific or pull factors - or by factors related to US economic performance - global or push factors. Using an estimation strategy which explicitly distinguishes between long-run and short-run determinants, we provide unequivocal evidence that long-run equity and bond flows are about equally sensitive to global factors and to country-specific factors, and therefore that both sets of variables contribute to explaining US portfolio flows to the developing countries considered. However, the short-run dynamics of portfolio flows to Asian developing countries, especially for equity flows, is found to be more complex relative to the short-run dynamics of portfolio flows to Latin American countries. As far as equity flows are concerned, global (push) and country-specific (pull) factors seem to be equally important in determining short-run movements for both Asian and Latin American countries. When bond flows are considered, however, global factors seem to be much more important than domestic factors in explaining the short-run dynamics of flows for both sets of countries. In particular, the change in US interest rates is found to be the single most important determinant of short-run movements in bond flows to the emerging markets.

Also, given that much of the recent capital inflow consists of flows labelled short term according to the World Bank's classification, a crucial issue for policy makers is to quantify the degree of persistence characterising capital flows received by developing countries. We addressed this question by measuring the relative size and statistical significance of the permanent and temporary components of capital flows to Latin American and Asian developing countries, on the basis that accounting labels may not in themselves be a reliable guide as to the "coolness" or "hotness" - persistence or temporariness - of capital flows. Using Kalman filtering maximum likelihood techniques in order to quantify the size of the unobserved nonstationary (permanent) and stationary (temporary) components present in equity and bond flows, supplemented by some simple non-parametric tests to measure their degree of persistence, we find that the permanent - or at least highly persistent - component in equity and bond flows to developing countries is very small in size relative to the temporary component.

The reason why the identification of the relative importance of push and pull factors and of the degree of persistence of capital flows are important for the effective design of policy is that large capital flows may, under various circumstances, have adverse effects on developing countries unless proper policies designed to neutralise such effects are adopted. More precisely, if the causes of capital flows are deemed largely external and exogenous to the developing country in question, then compensatory policies are appropriate. If, on the other hand, the causes of the large capital flows are deemed largely domestic, then a more direct policy design may be more appropriate and effective. From the policy maker's point of view, therefore, given the very large temporary component of capital inflows, developing countries should be wary of implementing painful adjustment processes that may need to be reversed in the future. Especially for countries with a more flexible exchange rate regime, capital inflows can potentially generate excessive volatility in both the nominal and the real exchange rate. The macroeconomic repercussions of the increase of portfolio flows of a highly temporary nature in the 1990s may be very high for several developing countries. Also, accounting labels attached to capital flows may be meaningless in providing information about the properties of the flows, since bond flows, which should be relatively more persistent from an accounting point of view, may be modelled in a similar fashion to equity flows, which one might in principle expect to be relatively less persistent, therefore requiring the same type of policy response needed against flows which are identified as "hot money".

An immediate, straightforward avenue for research in this area is to extend the sample period employed in this study. In particular, some of the global factors considered here have changed over the last two years or so and, while structural breaks may have been caused by these changes generating additional estimation difficulties, it will be interesting to investigate whether and how portfolio flows to developing countries have been affected by these changes.

In the last part of the thesis, we devoted our attention to the examination of nonfinancial - as opposed to financial - integration in the world economy, where nonfinancial refers to the mobility of goods and services and also the mobility of the production factors, labor and physical capital. The concept of nonfinancial integration in the world economy is, in practice, captured by the purchasing power parity proposition.

In particular, we provide a number of insights into multivariate tests for mean reversion in real exchange rates, both extending the development of and illustrating a potential pitfall in the interpretation of multivariate tests of long-run purchasing power parity which have been used by a number of researchers. Also, we examined the finite-sample performance of an alternative multivariate test which avoids this pitfall. Under the multivariate unit root test we suggest, in fact, we do not test the null hypothesis - common to most multivariate unit root tests - that *all* of the series are generated by non-stationary processes, but, rather, that *at least one* of the series is generated by a non-stationary process. This null hypothesis will only be violated if *all* of the series in question are realizations of stationary processes. Moreover, the suggested test procedure is widely available to researchers since it is just a special case of Johansen's maximum likelihood ratio for testing for the number of cointegrating vectors in a system, applied to the unusual case where the number of cointegration vectors tested for is equal to the number of series in the system under consideration. In addition, it turns out - again unusually - that this statistic has a limiting $\chi^2(1)$ distribution under the null hypothesis. Using Monte Carlo simulations, we could generate the finite sample empirical distributions of this test statistic and showed that the $\chi^2(1)$ distribution provides a reasonable approximation in samples larger than one hundred. We also demonstrated that the JLR test statistic has reasonable power characteristics for various departures from the null hypothesis.

Further, we apply the suggested test procedures to provide some new evidence on long-run purchasing power parity among the G5 countries during the recent float and, for real exchange rates constructed using relative consumer price indices, both tests strongly reject their respective null hypotheses at the five percent level of significance, indicating mean reversion in all the G5 real exchange rates over the floating rate period. For real exchange rates constructed using relative implicit GDP deflators, however, while we could strongly reject the hypothesis that all of the series are generated by non-stationary processes, we could not reject at the five percent level of significance the hypothesis that at least one of the series is generated by a non-stationary process. Taken together, therefore, the two test statistics imply that some but probably not all of the GDP deflator-adjusted real exchange rate series are mean reverting. This finding may be easily explained, however, on the basis of the higher proportion of non-tradables prices covered by the implicit GDP deflator compared with the consumer price index. Also, the same result was obtained for real exchange rates constructed using relative producer price indices. Given the finding of mean reversion in the real exchange rates constructed using consumer price indices, which would be more prone to Harrod-Balassa-Samuelson effects than the real exchange rates constructed using producer price indices because of the higher proportion of non-tradables prices covered, one interpretation is that the failure to establish mean-reversion of all the real exchange rates adjusted by producer price indices may be due to the loss in power resulting from the smaller available span of data.

In any case, the evidence that real exchange rates adjusted by consumer price indices among the G5 are apparently mean reverting over the floating rate period is, by itself, an important finding of our research, corroborating other recently emerging evidence that long-run purchasing power parity may hold after all. The strong implication of this finding is that shocks impinging upon real exchange rates may be highly persistent, but are temporary as there are forces that tend to restore purchasing power parity in the world economy in the long-run when the latter is disturbed.

It is a task for future research in this area to reconcile the finding that deviations from purchasing power parity slowly die out with the observed very high volatility of real exchange rates, which involves the modelling of financial and monetary shocks which, despite medium-run and long-run neutral, may be the major source of the large temporary deviations from purchasing power parity. In fact, clearly, real shocks - productivity or preferences shocks - are not volatile enough to explain deviations from purchasing power parity. On the other side, however, the speed of adjustment of real exchange rates towards long-run equilibrium seems too slow to be explained only by monetary and financial shocks, which suggests that a role should also be given to other factors such as - for example - trading frictions across countries due to transport costs, tariff and nontariff barriers, information costs, low factors mobility. In fact, we believe that the reconciliation of a slow mean-reversion in real exchange rates with large deviations from long-run equilibrium requires appropriate modelling of those factors.

In sum, we have that this thesis adds to the relevant literature in macroeconomics and international finance on the four different issues examined by providing insights and evidence to researchers, and indicating potential avenues for future research.

References

Abuaf, N. and Jorion, P. (1990), "Purchasing Power Parity in the Long Run", *Journal of Finance*, 45, pp. 157-174.

Adler, M. and Lehmann, B. (1983), "Deviations from Purchasing Power Parity in the Long Run", *Journal of Finance*, 38, pp. 1471-1487.

Agenor, P.-R. (1997), "The Surge in Capital Flows: Analysis of "Pull" and "Push" Factors", *International Journal of Finance and Economics*, forthcoming.

Akaike, H. (1973), "Information Theory and an Extension of the Maximum Likelihood Principle", in Petrov, B.N. and Csaki, F. (eds.), *Second International Symposium on Information Theory*, Budapest: Akademini Kiado, pp. 267-81.

Altonji, J.G. and Siow, A. (1987), "Testing the Response of Consumption to Income Changes with (Noisy) Panel Data", *Quarterly Journal of Economics*, 102, pp. 293-328.

Amemiya, T. (1974): "The Non-Linear Two Stage Least Squares Estimator", *Journal of Econometrics*, 2, pp. 105-110.

Amemiya, T. (1985), *Advanced Econometrics*, Cambridge, Mass.: Harvard University Press.

Anderson, T.W. and Rubin, H. (1949), "Estimation of the Parameters of a Single Equation in a Complete System of Stochastic Equations", *Annals of Mathematical Statistics*, 21, pp. 570-582.

Argimon, I. and Roldan, J.M. (1994), "Saving, Investment and International Capital Mobility in EC Countries", *European Economic Review*, 38, pp. 59-67.

Artis, M.J. and Bayoumi, T. (1991), "Global Capital Market Integration and the Current Account", in Taylor, M.P. (ed.), *Money and Financial Markets*, Cambridge, Mass. and Oxford: Blackwell, pp. 297-307.

Artis, M.J. and Taylor, M.P (1990), "Abolishing Exchange Control: The UK Experience", in Couriakis, A.S. and Taylor, M.P. (eds.), *Private Behaviour and Government Policy in Interdependent Economies*, Oxford: Clarendon Press.

Atkeson, A. and Bayoumi, T. (1993), "Do Private Capital Markets Insure Regional Risk? Evidence from the United States and Europe", *Open Economies Review*, 4, pp. 303-324.

Attanasio, O.P. (1995), "The Intertemporal Allocation of Consumption: Theory and Evidence", *Carnegie-Rochester Conference Series on Public Policy*, 42, pp. 39-89.

Attanasio, O.P. and Browning, M. (1995), "Consumption over the Life-cycle and over the Business Cycle", *American Economic Review*, 85, pp. 1187-1237.

Attanasio, O.P. and Weber, G. (1993), "Consumption Growth, the Interest Rate and Aggregation", *Review of Economic Studies*, 60, pp. 631-649.

Attanasio, O.P. and Weber, G. (1995), "Is Consumption Growth Consistent with Intertemporal Optimization? Evidence from the Consumer Expenditure Survey", *Journal of Political Economy*, 103, pp. 631-649.

Baffes, J. (1997), "Explaining stationary variables with nonstationary regressors", *Applied Economics Letters*, 4, pp. 69-75.

Balassa, B. (1964), "The Purchasing Power Parity Doctrine: A Reappraisal", *Journal of Political Economy*, 72, pp. 584-96.

Ballabriga, F.C., Dolado, J.J. and Vinals, J. (1991), "Investigating Public and Private Saving-Investment Gaps", *CEPR Working Paper* No. 607.

Banerjee, A., Dolado, J.J., Galbraith, J.W. and Hendry, D.F. (1993), *Co-Integration, Error Correction, and the Econometric Analysis of Non-Stationary Data*, Oxford University Press, New York.

Banerjee, A., Dolado, J.J., Hendry, D.F. and G.W. Smith (1986), "Exploring Equilibrium Relationships in Econometrics Through State Models: Some Monte Carlo Evidence", *Oxford Bulletin of Economics and Statistics*, 42(3).

Banerjee, A., Lumsdaine, R.L. and Stock, J.H. (1992), "Recursive and Sequential Tests of the Unit Root and Trend Break Hypothesis: Theory and International Evidence", *Journal of Business and economic Statistics*, 10.

Barro, R. (1974), "Are Government Bonds Net Wealth?", *Journal of Political Economy*, 81, pp. 1095-1117.

Barro, R.J. (1990), "Government Spending in a Simple Model of Endogenous Growth", *Journal of Political Economy*, October.

Barro, R.J. and Sala-i-Martin, X. (1990), "Public Finance in Models of Economic growth", *NBER Working Paper* No. 3362, May.

Barro, R.J., Mankiw, N.G. and Sala-i-Martin, X. (1995), "Capital Mobility in Neoclassical Models of Growth", *American Economic Review*, 85, pp. 103-115.

Basmann, R.L. (1957), "A Generalized Classic Method of Linear Estimation of Coefficients in a Structural Equation", *Econometrica*, 25, pp. 77-83.

Baumol, W. and Bowen, W. (1966), *Performing Arts: The Economic Dilemma*, New York: The Twentieth Century Fund.

Baxter, M. and Crucini, M.J. (1993), "Explaining Saving/Investment Correlations", *American Economic Review*, 83, pp. 416-436.

Baxter, M., and Stockman, A.C. (1989), "Business Cycles and the Exchange Rate System", *Journal of Monetary Economics* 23, pp. 377-400.

Bayoumi, T. (1990), "Financial Innovation and Consumption in the United Kingdom", *International Monetary Fund Working Paper* WP/90/95, Washington DC.

Bayoumi, T. (1990), "Saving-Investment Correlations", *International Monetary Fund Staff Papers*, 37, pp. 360-387.

Bayoumi, T. (1993), "Financial Deregulation and Consumption in the United Kingdom", *Review of Economics and Statistics*, 75, pp. 536-539.

Bayoumi, T. and Koujianou, P. (1990), "Consumption, Liquidity Constraints and Financial Deregulation", *Greek Economic Review*, 12, pp. 195-210.

Bayoumi, T. and Rose, A.K. (1993), "Domestic Savings and Intra-National Capital Flows", *European Economic Review*, 37, pp. 1197-1202.

Bayoumi, T. and Sterne, G. (1993), "Regional Trading Blocks, Mobile Capital, and Exchange Rate Coordination", London, Bank of England, mimeo.

Bean, C.R. (1986), "The Estimation of 'Surprise' Models and the 'Surprise' Consumption Function", *Review of Economic Studies*, 53, pp. 497-516.

Bekaert, G. (1995), "Market Integration and Investment Barriers in Emerging Equity Markets", *The World Bank Economic Review*, 9, pp. 75-107.

Bhandari, J. and Mayer, T. (1990), "A Note on Saving-Investment Correlations in the EMS", *International Monetary Fund Working Paper* No. 90/97.

Blanchard, O. and Quah, D. (1989), "The Dynamic Effects of Demand and Supply Disturbances", *American Economic Review*, 79, pp. 655-673.

Blundell, R., Browning, M. and Meghir, C. (1994), "Consumer Demand and the Lifetime Allocation of Consumption", *Review of Economic Studies*, 53, pp. 497-516.

Blundell-Wignall, A. and Browne, F. (1991), "Increasing Financial Market Integration, Real Exchange Rates, and Macroeconomic Adjustment", *OECD Working Paper* No. 96.

Bohn, H. and Tesar, L. (1996), "U.S. Equity Investment in Foreign Markets: Portfolio Rebalancing or Return Chasing?", *American Economic Review*, 86, pp. 77-81.

Breusch, T.S. and Pagan, A.R. (1979), "A Simple Test for Heteroskedasticity and Random Coefficient Variation", *Econometrica* 47, pp. 1287-94.

Browning, M. and Lusardi, A. (1996), "Household Saving: Micro Theories and Micro Facts", *Journal of Economic Literature*, 34, pp. 1797-1855.

Browning, M. and Meghir, C. (1991), "The Effect of Male and Female Labor Supply on Commodity Demands", *Econometrica*, 59, pp. 925-951.

Browning, M., Deaton, A.S. and Irish, M. (1985), "A Profitable Approach to Labor Supply and Commodity Demand over the Life Cycle", *Econometrica*, 53, pp. 503-544.

Bruno, M. (1993), *Foreword* in Claessens, S. and S. Gooptu (ed.) (1993), "Portfolio Investment in Developing Countries", *World Bank Discussion Papers*, The World Bank, Washington DC.

Buiter, W.H. (1993), "Saving and Endogenous Growth: A Survey of Theory and Policy", in Heertje, A. (ed.), *World Savings: An International Survey*, Cambridge: Blackwell.

Calvo, G., Leiderman, L. and Reinhardt, C. (1993), "Capital Inflows and the Real Exchange Rate Appreciation in Latin America: The Role of External Factors", *International Monetary Fund Staff Papers*, 40(1).

Calvo, G., Leiderman, L. and Reinhardt, C. (1996), "Inflows of Capital to Developing Countries in the 1990s", *Journal of Economic Perspectives*, 10, pp.123-139.

Campbell, J.Y. and Deaton, A.S. (1989), "Why is Consumption so Smooth?", *Review of Economic Studies*, 56, pp. 357-374.

Campbell, J.Y. and Mankiw, N.G. (1989), "Consumption, Income, and Interest Rates: Reinterpreting the Time Series Evidence", *NBER Working Paper* No. 2924, also published with discussants' comments in Blanchard, O.J. and Fischer, S. (eds.) *NBER Macroeconomics Annual 1989*, Cambridge: MIT Press.

Campbell, J.Y. and Mankiw, N.G. (1990), "Permanent Income, Current Income and Consumption", *Journal of Business and Economic Statistics*, 8, pp. 265-279.

Campbell, J.Y. and Mankiw, N.G. (1991), "The Response of Consumption to Income: A Cross-Country Investigation", *European Economic Review*, 35, pp. 715-721.

Campbell, J.Y. and Perron, P. (1991), "Pitfalls and Opportunities: What Macroeconomists Should Know About Unit Roots", in O.J. Blanchard and Fisher, S. (eds.), *NBER Macroeconomic Annuals* 6 (Cambridge, Mass., MIT Press).

Carroll, C.D. (1992), "Buffer Stock Saving and the Permanent Income Hypothesis", *Brookings Papers of Economic Activity*, 2, pp. 61-156.

Carroll, C.D. and Summers, L.H. (1991), "Consumption Growth Parallels Income Growth: Some New Evidence", in Bernheim, B.D. and Shoven, J.B. (eds.), *National Saving and Economic Performance*, Chicago: Chicago University Press for NBER, pp. 305-343.

Cassel, G. (1921), *The world's money problems*, New York: E.P Dutton and Co..

Cassel, G. (1922), *Money and foreign exchange after 1914*, New York: Macmillan.

Cheung, Y.W. and Lai, K.S. (1993a), "A Fractional Cointegration Analysis of Purchasing Power Parity", *Journal of Business Economics and Statistics*, 11, pp. 103-112.

Cheung, Y.W. and Lai, K.S. (1993b), Finite-Sample Sizes of Johansen's Likelihood Ratio Tests for Cointegration, *Oxford Bulletin of Economics and Statistics*, 55, pp. 313-328.

Cheung, Y.W. and Lai, K.S. (1995), "Purchasing Power Parity Under the European Monetary System", *Journal of International Money and Finance*, 14, pp. 179-189.

Choudhry, T., McNown, R. and Wallace, M. 1991, Purchasing Power Parity and the Canadian Float in the 1950s, *Review of Economics and Statistics*, 73, pp. 558-563.

Chuhan, P., Claessens, S. and Mamingi, N. (1993), "Equity and Bond Flows to Asia and Latin America", *Policy Research Working Papers* No. 1160, The World Bank, Washington DC.

Church, K.B., Smith, P.N. and Wallis, K.F. (1994), "Econometric Evaluation of Consumers' Expenditure Equations", *Oxford Review of Economic Policy*, 10, pp. 71-85.

Claessens, S., Dooley, M. and Warner, A. (1995), "Portfolio Capital Flows: Hot or Cold?", *The World Bank Economic Review*, 9, pp. 153-174.

Clarida, R.H. (1987), "Consumption, Liquidity Constraints and Asset Accumulation in the Presence of Random Income Fluctuations", *International Economic Review*, 28, pp. 339-351.

Clarida, R.H. (1991), "Aggregate Stochastic Implications of the Life-cycle Hypothesis", *Quarterly Journal of Economics*, 106, pp. 851-867.

Clarke, R. (1996), "Equilibrium Interest Rates and Financial Liberalisation in Developing Countries", *Journal of Development Studies*, 32, pp. 391-413.

Coakley, J., Kulasi, F. and Smith, R. (1994), "The Saving-Investment Association", *Birkbeck College Working Paper.*

Coakley, J., Kulasi, F. and Smith, R. (1995a), "The Feldstein-Horioka Puzzle and Capital Mobility", *Birkbeck College Discussion Paper.*

Coakley, J., Kulasi, F. and Smith, R. (1995b), "Current Account Solvency and the Feldstein-Horioka Puzzle", *Birkbeck College Discussion Paper.*

Cochrane, J.H. (1988), "How Big Is the Random Walk in GNP?", *Journal of Political Economy*, 96, pp. 893-920.

Corbae, D. and Ouliaris, S. (1988), "Cointegration and Tests of Purchasing Power Parity", *Review of Economics and Statistics*, 70, pp. 508-511.

Corbo, V. and Hernandez, L. (1993), "Macroeconomic Adjustment to Portfolio Capital Inflows: Rationale and Some Recent Experiences", in Claessens and Gooptu, op. cit..

Corrigan, E.G. (1989), "Reflections on the 1980s", *Federal Reserve Bank of New York Seventy-Fifth Annual Report for the Year Ended December 31, 1989*, pp. 5-22.

Cuthbertson, K., Hall, S.G. and Taylor, M.P. (1992), *Applied Econometric Techniques*, Harvester Wheatsheaf, Exeter.

Daly, V. and Hadjimatheou, G. (1981), "Stochastic Implications for the Life Cycle-Permanent Income Hypothesis: Evidence for the UK Economy", *Journal of Political Economy*, 89, pp. 596-599.

Daveri, F. (1995), "Costs of Entry and Exit from Financial Markets and Capital Flows to Developing Countries", *World Development*, 23, pp. 1375-1385.

Davidson, J.E.H. and Hendry, D.F. (1981), "Interpreting Econometric Evidence: Consumers' Expenditures", *European Economic Review*, 16, pp. 177-192.

Davidson, J.E.H., Hendry, D.F., Srba, F. and Yeo, S., (1978), "Econometric Modelling of the Aggregate Time-Series Relationship Between Consumers' and Income in the United Kingdom", *Economic Journal*, 88, pp. 661-692.

Davidson, R. and MacKinnon, J.G. (1993), *Estimation and Inference in Econometrics*, New York, USA and Oxford, UK: Oxford University Press.

De Grauwe, P. (1990), *International Money: Post-War Trends and Theories*, Oxford: Oxford University Press.

De Grauwe, P. (1994), *The Economics of Monetary Integration*, 2nd Revised Edition, New York: Oxford University Press.

Deaton, A.S. (1985), "Panel Data for Time Series of Cross Sections", *Journal of Econometrics*, 30, pp. 109-126.

Deaton, A.S. (1987), "Life-cycle Models of Consumption: Is the Evidence Consistent with the Theory?", in Bewley, T.F. (ed.) *Advances in Econometrics, Fifth World Congress*, 2, Cambridge and New York: Cambridge University Press, 121-148.

Deaton, A.S. (1991), "Saving and *Liquidity Constraints*", *Econometrica*, 1959, pp. 1221-1248.

Deaton, A.S. (1992), *Understanding Consumption*, Oxford, Clarendon Press.

Dickey, D. and Pantula, S. G. (1988), "Determining the Order of Differencing in Autoregressive Processes", *Journal of Business and Economic Statistics*, 5, pp. 455-461.

Dickey, D. and Fuller, W.A. (1979), "Distribution of the Estimators for Autoregressive Time Series with a Unit Root", *Journal of the American Statistical Association*, 74, pp. 427-431.

Dickey, D. and Fuller, W.A. (1981), "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root", *Econometrica*, 49, pp. 1057-1072.

Diebold, F., Husted S. and Rush, M. (1991), Real Exchange Rates under the Gold Standard, *Journal of Political Economy*, 99, pp. 1252-1271.

Dooley, M., Fernandez-Arias, E. and Kletzer, K. (1996), "Is the Debt Crisis History? Recent Private Capital Inflows to Developing Countries", *The World Bank Economic Review*, 10, pp. 27-50.

Dooley, M.P. (1995), "A Retrospective on the Debt Crisis", in Kenen, P. (ed.), *Understanding Interdependence*, Princeton University Press, Princeton, New Jersey.

Dooley, M.P., Frankel, J.A. and Mathieson, D.J. (1987), "International Capital Mobility: What Do Saving-Investment Correlations Tell Us?", *International Monetary Fund Staff Papers*, 34, pp. 503-530.

Dornbusch, R. (1987), "Purchasing Power Parity", in Eatwell, J., Milgate, M. and Newman, P. (eds.), *The New Palgrave: A Dictionary of Economics* (London: Macmillan, New York: Stockton press), pp. 1075-85.

Dornbusch, R. (1988), "Real Exchange Rates and Macroeconomics: A Selective Survey", *National Bureau of Economic Research Working Paper* No. 2775.

Dornbusch, R. (1988), "Real Exchange Rates and Macroeconomics: A Selective Survey", *National Bureau of Economic Research Working Paper* No. 2775.

Dornbush, R. (1991), Comment to Feldstein and Bacchetta in Bernheim, B.D. and Shoven, J.B. (eds.), *National Saving and Economic Performance*, Chicago, University of Chicago Press, pp. 201-220.

Dwivedi, T. and Srivastava, K. (1978), "Optimality of Least Squares in the Seemingly Unrelated Regressions Model", *Journal of Econometrics*, 7, pp. 391-395.

Edison, H.J. (1987), "Purchasing Power Parity in the Long Run", *Journal of Money, Credit and Banking* 19, pp. 376-387.

Enders, W. (1988), "ARIMA and Cointegration Tests of PPP under Fixed and Flexible Exchange Rate Regimes", *Review of Economics and Statistics* 70, pp. 504-508.

Engel, C. (1996), "Long-Run PPP May Not Hold After All", *National Bureau of Economic Research Working Paper* No. 5646.

Engel, C., Hendrickson, M.K. and Rogers, J.H. (1996), "Intra-National, Intra-Continental and Intra-Planetary PPP", University of Washington, mimeo.

Engle, R. and Granger, C.W.J. (1987), "Cointegration and Error Correction: Representation, Estimation, and Testing", *Econometrica*, 55, pp. 251-276.

Feldstein, M. (1983), "Domestic Saving and International Capital Movements in the Long Run and the Short Run", *European Economic Review*, 21, pp. 129-151.

Feldstein, M. and Bacchetta, P. (1991), "National Saving and International Investment", in Bernheim, B.D. and Shoven, J.D. (eds.), *National Saving and Economic Performance*, Chicago, University of Chicago Press, pp. 201-220.

Feldstein, M. and Horioka, C. (1980), "Domestic Saving and International Capital Flows", *Economic Journal*, 90, pp. 314-329.

Fernandez-Arias, E. (1996), "The New Wave of Capital Inflows: Push or Pull?", *Journal of Development Economics*, 48, pp. 389-418.

Fernandez-Arias, E. and Montiel, P.J. (1996), "The Surge in Capital Inflows to Developing Countries: An Analytical Overview", *The World Bank Economic Review*, 10, pp. 51-77.

Fischer, S. and Merton, R.C. (1984), "Macroeconomics and Finance: The Role of the Stock Market", in *Carnegie-Rochester Conference Series on Public Policy*, 21, by Brunner, K. and Meltzer, A.H. (eds.), Amsterdam: North-Holland, pp. 57-108.

Fitoussi, J. and Le Cacheux, J. (1993), "Saving in Western Europe", in A. Heertje (ed.), op. cit..

Flavin, M. (1981), "The Adjustment of Consumption to Changing Expectations About Future Income", *Journal of Political Economy*, 89, pp. 974-1009.

Flemming, J.S. (1973), "The Consumption Function when Capital Markets are Imperefect, The Permanent Income Hypothesis Reconsidered", *Oxford Economic Papers*, 25, pp. 160-172.

Flood, R.P. and Taylor, M.P. (1996), "Exchange Rate Economics: What's Wrong with the Conventional Macro Approach?", in J.A. Frankel, Galli, G. and Giovannini, A. (eds.), *The Microstructure of Foreign Exchange Markets*, Chicago: Chicago University Press for NBER.

Frankel, J.A. (1993), "Quantifying International Capital Mobility in the 1980s", in Frankel, J.A. (ed.), *On Exchange Rates*, Cambridge, Mass.: MIT Press, pp. 41-69.

Frankel, J.A. and Rose, A.K. (1996), "A Panel Project on Purchasing Power Parity: Mean Reversion Within and Between Countries", *Journal of International Economics*, 40, pp. 209-224.

Frankel, J.A. and MacArthur, A.T. (1988), "Political vs. Currency Premia in International Real Interest Rate Differentials: A Study of Forward Rates for 24 Countries", *European Economic Review*, 32, pp. 1083-1121.

Frankel, J.A. (1986), "International Capital Mobility and Crowding Out in the US Economy: Imperfect Integration of Financial Economics or of Goods Markets?", in Hafer, R. (ed.), *How Open is the US Economy?*, Lexington MA: Lexington Books.

Frankel, J.A. (1989), "Zen and the Art of Modern Macroeconomics: The Search for Perfect Nothingness", in Haraf, W. and Willett, T. (eds.), *Monetary Policy for a Volatile Global Economy*, Washington DC: American Enterprise Institute.

Frenkel, J. and Razin, A. (1987), *Fiscal Policies and World Economy*, Cambridge, Mass.: MIT Press.

Frenkel, J.A. (1976), "A Monetary Approach to the Exchange Rate: Doctrinal Aspects and Empirical Evidence", *Scandinavian Journal of Economics* 78, pp. 200-224.

Frenkel, J.A. (1981), "The Collapse of Purchasing Power Parities During the 1970s", *European Economic Review* 16, pp. 145-65.

Frenkel, J.A., and Johnson, H.G. (1978), *The Economics of Exchange Rates*, Reading MA: Addison-Wesley.

Friedman, M. and Schwartz, A.J. (1963), *A Monetary History of the United States: 1867-1960*, Princeton, NJ: Princeton University Press for the National Bureau of Economic Research.

Frisch, R. and Waugh, F.V. (1933), "Partial Time Regressions As Compared With Individual Trends", *Econometrica*, 1, pp. 387-401.

Froot, K.A. and Rogoff, K. (1995), "Perspectives on PPP and Long-run Real Exchange Rates", in Rogoff, K. and Grossman, G. (eds.), *Handbook of International Economics*, Amsterdam: North Holland.

Fuller, W.A. (1976), *Introduction to Statistical Time Series*, New York: Wiley and Sons.

Gaillot, H.J. (1970), "Purchasing Power Parity as an Explanation of Long-Term Changes in Exchange Rates", *Journal of Money Credit and Banking*, 2, pp. 348-357.

Gali, J. (1990), "Finite Horizons, Life-cycle Savings and Time-series Evidence on Consumption", *Journal of Monetary Economics*, 26, pp. 433-452.

Gali, J. (1991), "Budget Constraints and Time-series Evidence on Consumption", *American Economic Review*, 81, pp. 1238-1253.

Gallant, A. R. (1987), *Non-linear Statistical Models*, New York: John Wiley.

Ghosh, A.R. (1995), "International Capital Mobility among the Major Industrialised Countries: Too Little or Too Much", *Economic Journal*, 105, pp. 107-128.

Glick, R. and Rogoff, K. (1995), "Global versus Country-Specific Productivity Shocks and the Current Account", *Journal of International Monetary Economics*, 35, pp. 159-192.

Goldfeld, S.M. and Quandt, R.E. (1972), *Nonlinear Methods in Econometrics*, Amsterdam: North-Holland.

Goldstein, M., Mathieson, D.J. and Lane, T. "Determinants and Systemic Consequences of International Capital Flows", in *Determinants and Systemic Consequences of International Capital Flows* (March 1991), Occasional Paper No. 77, International Monetary Fund, Washington DC.

Gooptu, S. (1993), "Portfolio Investment Flows to Emerging Markets", in Claessens and Gooptu, op. cit..

Grabel, I. (1995), "Assessing the Impact of Financial Liberalisation on Stock Market Volatility in Selected Developing Countries", *Journal of Development Studies*, 31, pp. 903-917.

Granger, C.W.J. (1983), "Co-integrated Variables and Error Correcting Models", *University of California, San Diego Discussion Paper*.

Granger, C.W.J. and Morris, M.J. (1976), "Time Series Modelling and Interpretation", *Journal of the Rotal Statistical Society*, Series A, 139, pp. 246-257.

Granger, C.W.J. and Newbold, P. (1974): "Spurious Regressions in Econometrics", *Journal of Econometrics*, 26, pp. 1045-1066.

Green, C.J. and Llewellyn, D.T. (eds.) (1991), *Surveys in Monetary Economics*, Vol. 2, Oxford: Blackwell.

Grilli, V. and Kaminky, G. (1991), Nominal Exchange Rate Regimes and the Real Exchange Rate: Evidence for the United States and Great Britain, 1885-1986, *Journal of Monetary Economics*, 27, pp. 191-212.

Grossman, G.M. and Helpman, E. (1991), *Innovation and Growth in the Global Economy*, Cambridge, MA: MIT Press.

Guiso, L., Jappelli, T. and Terlizzese, D. (1994), "Why is Italy's Saving Rate so High?", in Ando, A., Guiso, L. and Visco, I. (eds.) *Saving and the Accumulation of Wealth*, Cambrdige: Cambridge University Press, pp. 23-69.

Hakkio, C.S. and Joines, D.H. (1990), "Real and Nominal Exchange Rates Since 1919", *Federal Reserve Bank of Kansas City Working Paper*.

Hakkio, C.S. and Rush, M. 1991, Cointegration: How Short is the Long-Run?, *Journal of International Money and Finance*, 10, pp. 571-581.

Hall, A.R., Rudebusch, G.D. and Wilcox, D.W. (1994), "Judging Instrument Relevance in Instrumental Variables Estimation", *North Carolina State University Discussion Paper*, September.

Hall, M.J.B. (1991), "Financial Regulation in the UK: Deregulation or Reregulation?", in Green and Llewellyn (1991).

Hall, R.E. (1978), "Stochastic Implications of the Life Cycle-Permanent Income Hypothesis: Theory and Evidence", *Journal of Political Economy*, 96, pp. 971-987.

Hall, R.E. (1978), "Stochastic Implications of the Life Cycle-Permanent Income Hypothesis: Theory and Evidence", *Journal of Political Economy*, 96, pp. 971-987.

Hall, R.E. (1988), "Intertemporal Substitution in Consumption", *Journal of Political Economy*, 96, pp. 339-357.

Hall, R.E. and Mishkin, F.S. (1982), "The Sensitivity of Consumption to Transitory Income: Estimates from Panel data on Households", *Econometrica*, 50, pp. 461-481.

Hansen, L.P. (1982), "Large Sample Properties of Generalized Method of Moments Estimators", *Econometrica*, 50, pp. 1029-1054.

Harrod, R. (1933), *International Economics*, London: James Nisbet and Cambridge University Press.

Harvey, A.C. (1981), *Time Series Models*, Hemel Hempstead, Philip Allan.

Harvey, A.C. (1989), *Forecasting, Structural Time Series Models and the Kalman Filter*, Cambridge University Press, Cambridge Mass..

Heckman, *J.J. (1974)*, "Life-cycle Consumption and Labor Supply: An Exploration of the Relationship Between Income and *Consumption Over the Life Cycle*", *American Economic Review*, 64, pp. 188-194.

Hegwood, N.D. and Papell, D.H. (1996), "Quasi Purchasing Power Parity, mimeo, Department of Economics", *University of Houston Discussion Paper.*

Helpman, E. (1991), "Endogenous Growth Macroeconomic Theory", Tel Aviv University, mimeo, August.

Hendry, D.F. (1984), "Monte Carlo Experimentation in Econometrics", in Griliches, Z. and Intrilligator, M.D. (eds.), *Handbook of Econometrics* (Amsterdam, New York and Oxford: North-Holland), pp. 939-76.

Hendry, D.F. (1983), "Econometric Modelling: The Consumption Function in Retrospect", *Scottish Journal of Political Economy*, 30, pp. 193-220.

Hendry, D.F. and von Ungern-Sternberg, T. (1981), "Liquidity and Inflation Effects on Consumers' Expenditure", in A.S. Deaton (ed.), *Essays in Theory and Measurement of Consumers' Behavior*, Cambridge: Cambridge University Press, pp. 237-260.

HMSO (Her Majesty's Stationery Office) (1980), *Monetary Control* (consultation paper issued by HM Treasury and the Bank of England). London: HMSO, Cmnd 7858.

HMSO (Her Majesty's Stationery Office) (1984), *Building Societies: A New Framework*, London: HMSO, Cmnd 9316.

Horioka, C.Y. (1993), "Saving in Japan", in A. Heertje (ed.), op. cit..

Hume, D. (1752), *On the Balance of Trade*, reprinted in R.N. Cooper, ed., 1969, *International Finance: Selected Readings*, London: Penguin.

Ickes, B.W. (1993), "Saving in Estern Europe and the former Soviet Europe", in A. Heertje (ed.), op. cit..

Im, K.S., Pesaran, M.H. and Shin, Y. (1997), "Testing for Unit Roots in Heterogeneous Panels", mimeo, Department of Applied Economics, University of Cambridge, *Econometrica*, forthcoming.

Imrohoroglu, A. (1989), "Cost of Business Cycle with Indivibilities and Liquidity Constraints", *Journal of Political Economy*, 100, pp. 118-142.

International Monetary Fund (November 1995), *Private Market Financing for Developing Countries*, International Monetary Fund, Washington DC.

International Monetary Fund (October 1994a), *World Economic Outlook, World Economic and Financial Surveys*, International Monetary Fund, Washington DC.

International Monetary Fund (September 1994b), *International Capital Markets: Developments, Prospects and Policy Issues*, International Monetary Fund, Washington DC.

Jappelli, T. (1990), "Who is Credit Constrained in the US Economy?", *Quarterly Journal of Economics*, 105, pp. 219-234.

Jappelli, T. and Pagano, M. (1989), "Consumption and Capital Market Imperfections: An International Comparison", *American Economic Review*, 79, pp. 1088-1105.

Jappelli, T. and Pagano, M. (1994), "Saving, Growth and Liquidity Constraints", *Quarterly Journal of Economics*, 109, pp. 83-110.

Jarque, C.M. and Bera, A.K. (1980), "Efficient Tests for Normality, Heteroskedasticity and Serial Independence of Regression Residuals", *Economics Letters*, 6, pp. 255-59.

Johansen, S. (1988), "Statistical Analysis of Cointegration Vectors", *Journal of Economic Dynamics and Control*, 12, pp. 231-254.

Johansen, S. (1991), "Estimation and Hypothesis Testing of Cointegrating Vectors in Gaussian Vector Autoregressive Models", *Econometrica*, 59, pp. 1551-1580.

Johansen, S. and Juselius, K. (1990), "Maximum Likelihood Estimation and Inference on Cointegration with Application to the Demand for Money", *Oxford Bulletin of Economics and Statistics*, 52, pp. 169-209.

Jorion, P. and Sweeney, R. (1996), "Mean Reversion in Real Exchange Rates: Evidence and Implications for Forecasting", *Journal of International Money and Finance* 15, pp. 535-550.

Kalman, R.E. (1960a), "A New Approach to Linear Filtering and Prediction Theory", *Journal of Basic Engineering*, March, pp. 1-35.

Kalman, R.E. (1960b), "A New Approach to Linear Filtering and Prediction Problems", *Journal of Basic Engineering, Transactions of the ASME*, Series D, 82, pp. 35-45.

Keane, M. and Runkle, D. (1992), "On the Estimation of Panel-Data Models with Serial Correlation when Instruments are not Strictly Exogenous", *Journal of Business and Economic Statistics*, 10, pp. 1-9.

Kelejian, H.H. (1971), "Two-stage Least Squares and Econometric Systems Linear in Parameters but Nonlinear in Endogenous Variables", *Journal of the American Statistical Association*, 66, pp. 373-374.

Kendall, M.G. and Stuart, A. (1967), *The Advanced Theory of Statistics*, 2, London: Charles Griffin.

Keynes, J.M. (1932), *Essays in Persuasion*, New York: Harcourt Brace.

Kiguel, M. and Caprio, G. (1993), "Capital Inflows - Hot or Cool?", *Outreach* 8.

Kim, Y. (1990), "Purchasing Power Parity in the Long Run: A Cointegration Approach", *Journal of Money, Credit and Banking* 22, pp. 348-357.

Kimball, M.S. (1990), "Precautionary Saving in the Small and in the Large", *Econometrica*, 58, pp. 53-73.

Kindleberger, C.P. (1973), *The World in Depression, 1929-1939*, University of California Press.

Kremers, J.J.M., Ericsson, N.R. and Dolado, J.J. (1992), "The Power of Cointegration Tests", *Oxford Bullettin of Economics and Statistics*, August, pp. 325-348.

Krugman, P. (1978), "Purchasing Power Parity and exchange Rates: Another Look at the Evidence", *Journal of International Economics*, 8, pp. 397-407.

Lau, L.J. (1993), "US Saving Behavior in the Post-War Period", in A. Heertje (ed.), op. cit..

Lemmen, J.J.G. and Eijffinger, S.C.W. (1995a), "The Quantity Approach to Financial Integration: The Feldstein-Horioka Approach Criterion Revisited", *Open Economies Review*, 6, pp. 145-165.

Lemmen, J.J.G. and Eijffinger, S.C.W. (1995b), "Financial Integration in Europe: Evidence from Euler Equation Tests", *Centre for Economic Research Discussion Paper* No. 9532.

Litterman, R.B. and Weiss, L. (1985), "Money, Real Interest Rates and Output: A Reinterpretation of Postwar U.S. Data", *Econometrica*, 53, pp. 129-156.

Llewellyn, D.T. (1991), "Structural Change in the British Financial System", in Green and Llewellyn (1991), op. cit...

Lothian, J.R. and Taylor, M.P (1995a), "Real Exchange Rate Behaviour: The Recent Float from the Perspective of the Past Two Centuries", *Liverpool Research Papers in Economics, Finance and Accounting* No. 9503.

Lothian, J.R. and Taylor, M.P (1995b), "Real Exchange Rate Behavior: The Problem of Power and Sample Size", *Journal of International Money and Finance*, forthcoming.

Lothian, J.R. and Taylor, M.P (1996), "Real Exchange Rate Behavior: The Recent Float from the Perspective of the Past Two Centuries", *Journal of Political Economy*, 104, pp. 488-510.

Lothian, J.R. (1990), A Century Plus of Yen Exchange Rate Behavior, *Japan and the World Economy*, 2, pp. 47-90.

Lovell, M.C. (1963), "Seasonal Adjustment of Economic Time Series", *Journal of the American Statistical Association*, 58, pp. 993-1010.

Lucas, R.E. Jr (1976), "Econometric Policy Evaluation: A Critique", in Brunner, K. and Meltzer, A.H. (eds.) *The Phillips Curve and Labor Markets, Carnegie Rochester Series on Public Policy*, 1, pp. 19-46, Supplement to the *Journal Monetary Economics*.

Lucas, R.E., Jr. (1982), Interest Rates and Currency Prices in a Two-Country World, *Journal of Monetary Economics*, 10, pp. 335-359.

Lucas, R.E.J. (1988), "On The Mechanics of Economic Development", *Journal of Monetary Economics*, 22, pp. 3-42.

Lusardi, A. (1996), "Permanent Income, Current Income, and Consumption: Evidence from Two Panel Data Sets", *Journal of Business Economics and Statistics*, 14, pp. 81-90.

MacKinnon, J.G. (1991), "Critical Values for Cointegration Tests", in Engle, E.F. and Granger, C.W.J. (eds.), *Long-Run Economic Relationships: Readings in Cointegration*, Oxford: Oxford University Press.

Maddala, G.S. and Jeong, J. (1992), "On the Exact Small-Sample Distribution of the Instrumental Variables Estimator", *Econometrica*, 60, pp. 181-3.

Mankiw, N.G. (1982), "Hall's Consumption Hypothesis and Durable Goods", *Journal of Monetary Economics*, 10, pp. 417-426.

Mankiw, N.G. and Shapiro, M. (1985), "Trends, Random Walks and Tests of the Permanent Income Hypothesis", *Journal of Monetary Economics*, 16, pp. 165-174.

Mankiw, N.G. and Shapiro, M. (1986), "Do We Reject Too Often? Small Sample Properties of Tests of Rational Expectations Models", *Economics Letters*, 20, pp. 139-145.

Mankiw, N.G., Romer, D. and Weil, D.N. (1992), "A Contribution in the Empirics of Economic Growth", *Quarterly Journal of Economics*, 107, pp. 4-7-437.

Mark, N. (1990), "Real and Nominal Exchange Rates in the Long Run: An Empirical Investigation", *Journal of International Economics*, 28, pp. 115-136.

Masson, P.R. and Taylor, M.P. (1993), "Currency Unions: A Survey of the Issues", in Masson, P.R. and Taylor, M.P. (eds.) *Policy Issues in the Operation of Currency Unions* (1993), Cambridge: Cambridge University Press.

Mathieson, D. and Rojas-Suarez, L. (1990), "Financial Market Integration and Exchange Rate Policy", *International Monetary Fund Working Paper* No. 90/2.

Mathieson, D. and Rojas-Suarez, L. (1992), "Liberalization of the Capital Account", *International Monetary Fund Working Paper*, WP/92/46.

Mathieson, D.J. and Rojas-Suarez, L. (1990), "Financial Market Integration and Exchange Rate Policy," *International Monetary Fund Working Paper* WP/90/2, Washington DC.

McCloskey, D. and Zecher, J.R. (1976), "How the Gold Standard Worked, 1880-1913", in Frenkel, J.A. and Johnson, H.G. (eds.), *The Monetary Approach to the Balance of Payments*, Toronto: University of Toronto Press.

McKinnon, R.I. (1971), *Monetary Theory and Controlled Flexibility in the Foreign Exchanges*, Princeton, NJ: International Finance Section, Princeton University.

McNown, R., and Wallace, M. (1989), "National Price Levels, Purchasing Power Parity and Cointegration: A Test of Four High-Inflation Economies", *Journal of International Money and Finance*, 8, pp. 533-545.

Miles, D. (1992), "Housing Markets', Consumption, and Financial Liberalization in the Major Economies", *European Economic Review*, 36, pp. 1093-1127.

Modigliani, F. (1993), Introduction in Heertje (ed.), *World Savings: An International Survey*, Cambridge: Blackwell.

Montiel, P. (1993), "Capital Mobility in Developing Countries", *PRE Paper* No.1103, The World Bank, Washington DC.

Muellbauer, J. (1994), "The Assessment: Consumer Expenditure", *Oxford Review of Economic Policy*, 10, pp. 1-41.

Muellbauer, J. (1994), "The Assessment: Consumer Expenditure", *Oxford Review of Economic Policy*, 10, pp. 1-41.

Muellbauer, J. and Murphy, A. (1990), "Is the US Balance of Payments Sustainable?", *Economic Policy*, pp. 347-382.

Muellbauer, J. and Murphy, A. (1993), "Income Expectations, Wealth and Demography in the Aggregate US Consumption Function", unpublished paper presented to H.M. Treasury Academic Panel, Nuffield College, University of Oxford.

Mullebauer, J. (1988): "Habits, Rationality and Myopia in the Life Cycle Consumption Function", *Annales d'Economie et de Statistique*, 9, pp. 47-70.

Mussa, M. (1986), "Nominal Exchange Rate Regimes and the Behavior of Real Exchange Rates: Evidence and Implications", *Carnegie-Rochester Conference Series on Public Policy* 25, pp. 177-213.

Nelson, C.R. and Startz, R. (1990a), "The Distribution of the Instrumental Variables Estimator and its t-Ratio When the Instrument Is a Poor One", *Journal of Business*, 63, pp. 125-140.

Nelson, C.R. and Startz, R. (1990b), "Some Further Results on the Exact Small-Sample Properties of the Instrumental Variables Estimator", *Econometrica*, 58, pp. 967-76.

Nelson, C.R. (1976), "Gains in Efficiency from Joint Estimation of Systems of Autoregressive Moving Average Processes", *Journal of Econometrics*, 4, 331-348.

Nicoletti-Altimari, S. and Thomson, M.D. (1995), "The Effect of Liquidity Constraints on Consumption and Labor Supply: Evidence from Italian Households", *Temi di Discussione Banca d'Italia*, No. 252.

Nunnenkamp, P. (1993), "The Return of Foreign Capital to Latin America", *Kiel Working Paper* 574.

O'Connell, P.G.J. (1996), The Overvaluation of Purchasing Power Parity, *Journal of International Economics*, forthcoming.

Obstfeld, M. (1986), "Capital Mobility in the World Economy: Theory and Measurement", *Carnegie-Rochester Conference Series on Public Policy*, 24, pp. 55-103.

Obstfeld, M. (1995), "International Capital Mobility in the 1990s", in Kenen, P.B. (ed.), *Understanding Interdependence: The Macroeconomics of the Open Economy*, Princeton University Press.

Obstfeld, M. and Rogoff, K. (1995a), "The Intertemporal Approach to the Current Account", in Grossman, G.M. and Rogoff, K. (eds.), *Handbook of International Economics*, 3, pp. 1731-1799.

Obstfeld, M. and Rogoff, K. (1995b), Exchange Rate Dynamics Redux, *Journal of Political Economy*, 103, pp. 624-660.

Obstfeld, M. and Rogoff, K. (1996), *Foundations of International Macroeconomics*, Cambridge, Mass. and London, UK: MIT Press.

Officer, L.H. (1982), *Purchasing Power Parity and Exchange Rates: Theory, Evidence and Relevance*, Greenwich CT: JAI Press.

Oh, K.Y. (1996), "Purchasing Power Parity and Unit Root Tests Using Panel Data", *Journal of International Money and Finance*, 15, pp. 405-418.

Organization for Economic Cooperation and Development (OECD) (1987), *Structural Adjustment and Economic Performance*, Paris, OECD.

Pagan, A.R. and Wickens, M.R. (1989), "A Survey of Some Recent Econometric Methods", *Economic Journal*, 99, pp. 962-1025

Papaioannu, M.G. and Duke, L.K. (1993), "The Internationalization of Emerging Equity Markets", *Finance & Development*, September, pp. 36-39.

Papell, D.H. (1996), "Searching for Stationarity: Purchasing Power Parity Under the Current Float", University of Houston Working Paper, *Journal of International Economics*, forthcoming.

Patat, J.P. (1993), *Monnaie, Institutions Financieres et Politiques monetaires*, Economica (5eme edition).

Patterson, K.D. (1984a), "Net Liquid Assets and Net Illiquid Assets in the Consumption Function: Some Evidence for the United Kingdom", *Economics Letters*, 14, pp. 389-395.

Patterson, K.D. (1984b), "Some Properties of Consumption Functions with Integral Correction Mechanisms", *The Manchester School*, 70, pp. 347-362.

Patterson, K.D. (1987), "The Stability of Some Annual Consumption Functions", *Oxford Economic Papers*, 38, pp. 1-30.

Patterson, K.D. and Pesaran, B. (1992), "The Intertemporal Elasticity of Substitution in Consumption in the United States and the United Kingdom", *Review of Economics and Statistics*, 74, pp. 573-584.

Pecha, P. and Sicsic, P. (1988), "Developpement du Credit a la Concommation et Economie Reele", *Revue d'Economie Financiere*, 5/6, pp. 212-223.

Perron, P. (1988), "Trends and Random Walks in Macro-economic Time Series: Further Evidence from a New Approach", *Journal of Economic Dynamics and Control*, 12, pp. 297-332.

Perron, P. (1989), "The Great Crash, the Oil Price Shock and the Unit Root Hypothesis", *Econometrica*, 57, pp. 1361-1401.

Pesaran, B. (1988), "Exact Maximum Likelihood Estimation of Regression Models with Invertible General Order Moving Average Disturbances", *National Institute of Ecocomic and Social Research Discussion Paper* No. 136.

Pesaran, B. (1990), "Instrumental Variables Estimation of Regression Models with Invertible General Order Moving-Average Disturbances", mimeo Bank of England.

Pesaran, M.H. and Smith, R.P. (1995), "The Role of Theory in Econometrics", *Journal of Econometrics*, 67, pp. 61-79.

Piggott, C. and Sweeney, R.J. (1985), "Purchasing Power Parity and Exchange Rate Dynamics", in Arndt, S.W., Sweeney, R.J. and Willett, D. (eds.), *Exchange Rates, Trade and the US Economy*, Cambridge MA: Ballinger and American Enterprise Institute.

Powell, M.J.D. (1964), "An Efficient Method of Finding the Minimum of a Function of Several Variables without Calculating Derivatives", *The Computer Journal*, 7, pp. 155-162.

Quah, D. (1994), "Exploiting Cross Section Variation for Unit Root Inference in Dynamic Data", *Economics Letters*, 44, pp. 9-19.

Ramsey, J.B. (1969), "Tests for Specification Errors in Classical Linear Least Squares Regression Analysis", *Journal of the Royal Statistical Society, series B*, 31, pp. 350-71.

Razin, A. (1995), "The Dynamic-Optimizing Approach to the Current Account: Theory and Evidence", in Kenen, P.B. (ed.), *Understanding Interdependence: The Macroeconomics of the Open Economy*, Princeton: Princeton University Press.

Reinsel, G.C. and Ahn, S.K. (1988), "Asymptotic Properties of the Likelihood Ratio Test for Cointegration in the Non-Stationary Vector AR Model", Technical Report, Department of Statistics, University of Wisconsin-Madison.

Reisen, H. (1993), "The Case of Sterlized Intervention in Latin America", Paper presented at the 6th Annual Inter-American Seminar in Economics, May 28-29, Caracas, Venezuela.

Reisen, H. and Fischer, B. (1993), *Financial Opening: Policy Issues and Experiences in Developing Countries*, OECD Development Centre, Paris.

Ricardo, D. (1821), *The Principles of Political Economy and Taxation*, London: J.M. Dent and Sons Ltd.

Rogoff, R. (1996), "The Purchasing Power Parity Puzzle", *Journal of Economic Literature*, 34, pp. 647-668.

Rogoff, R. (1996), "The Purchasing Power Parity Puzzle", *Journal of Economic Literature*, 34, pp. 647-668.

Roll, R. (1979), "Violations of Purchasing Power Parity and their Implications for Efficient International Commodity Markets", in Sarnat, M. and Szego, G.P. (eds.), *International Finance and Trade*, Vol 1 (Cambridge MA: Ballinger).

Romer, P. (1986), "Increasing Returns and Long-Run Growth", *Journal of Political Economy*, 94, pp. 1002-1037.

Romer, P. (1990), "Endogenous Technological Change", *Journal of Political Economy*, 98, pp. 71-102.

Runkle, D. (1991), "Liquidity Constraints and the Permanent Income Hypothesis", *Journal of Monetary Economics*, 27, pp. 73-98.

Sachs, J.D. (1981), "The Current Account and Macroeconomic Adjustment in the 1970s", *Brookings Papers in Economic Activity*, 12, pp. 1089-1092.

Sachs, J.D. (1982), "The Current Account in the Macroeconomic Adjustment Process", *Scandinavian Journal of Economics*, 84, pp. 147-159.

Sachs, J.D. and Sala-i-Martin, X. (1992), "Fiscal Federalism and Optimum Currency Areas: Evidence from Europe and from the United States", in Canzoneri, M., Grilli, V. and Masson, P.R. (eds.), *Establishing a Central Bank: Issues in Europe and Lessons from the US*, Cambridge: Cambridge University Press, pp. 1950-1219.

Sala-i-Martin, X. (1990a), "Lecture Notes on Economic Growth: Introduction to the Literature and Neoclassical Models", 1, *Economic Growth Centre Discussion Paper* No. 621, December.

Sala-i-Martin, X. (1990b), "Lecture Notes on Economic Growth: Introduction to the Literature and Neoclassical Models", 2, *Economic Growth Centre Discussion Paper* No. 622, December.

Samuelson, P. (1939), "Interactions between the Multiplier Analysis and the Principle of Acceleration", *Review of Economics and Statistics*, 21, pp. 75-78.

Samuelson, P. (1964), "Theoretical Notes on Trade Problems", *Review of Economics and Statistics*, 46, pp. 145-54.

Sargan, J.D. (1964), "Wages and Price in the United Kingdom: A Study in Econometric Methodology", in Hart, P.E., Mills, G. and Whittaker, J.N. (eds.), *Econometric Analysis for National Economic Planning*, London, Butterworths.

Sargan, J.D. and Barghava, A. (1983), "Testing Residuals from Least Square Regression for Being Generated by the Gaussian Random Walk", *Econometrica*, 51, pp. 153-174.

Sargan, J.D. (1958), "The Estimation of Economic Relationships Using Instrumental Variables", *Econometrica*, 26, pp. 393-415.

Schwartz, G. (1978), "Estimating the Dimension of a Model", *The Annals of Statistics*, 6, pp. 461-464.

Schweppe, F.C. (1965), "Evaluation of Likelihood Functions for Gaussian Signals", *IEEE Transactions on Information Theory*, 11, pp. 61-70.

Shafer, J.R. (1995), "Experience with Controls on International Capital Movements in OECD Countries: Solution or Problem for Monetary Policy?", in Edwards, S. (ed.) (1995), *Capital Controls, Exchange Rates and monetary Policy in the World Economy*, New York: Cambridge University Press.

Shiller, R.J. and Perron, P. (1985), "Testing the Random Walk Hypothesis: Power Versus Frequency of Observation", *Economics Letters*, 18, pp. 381-6.

Sims, C.A. (1980), "Comparison of Interwar and Postwar Cycles: Monetarism Reconsidered", *American Economic Review*, 70, pp. 250-277.

Sinn, S. (1992), "Saving-Investment Correlations and Capital Mobility: on the Evidence from Annual Data", *Economic Journal*, 102, pp. 1162-1170.

Skinner, J. (1988), "Risky Income, Life-Cycle Consumption, and Precautionary Saving", *Journal of Monetary Economics*, 22, pp. 237-255.

Spanos, A. (1986), *Statistical Foundations of Econometric Modelling*, Cambridge: Cambridge University Press.

Staiger, D. and Stock, J.H. (1994), "Instrumental Variables Regression with Weak Instruments", *NBER Technical Working Paper* No. 151.

Stefani, M.L. (1994), "Saving-investment Correlations and Governments Policy under Credit Constraints", *IRES Discussion Paper* No. 9404.

Stiglitz, J.E. and Weiss, A. (1981), "Credit Rationing in Markets with Imperfect Information", *American Economic Review*, 71, pp. 393-410.

Stock, J.H. (1987), "Asymptotic Properties of Least Squares Estimators of Cointegrating Vectors", *Econometrica*, 55, pp. 1035-1056.

Stock, J.H. (1990), "Unit Roots in Real GNP: Do We Know and Do We Care? A Comment", *Carnegie-Rochester Series on Public Policy*, 32, pp. 63-82.

Stock, J.H. (1993), "Unit Root and Trend Breaks in Econometrics", in Pesaran, M.H. and Wickens, M. (eds.), *The Handbook of Applied Econometrics*, forthcoming.

Stock, J. and Watson, M. (1988), "Testing for Common Trends", *Journal of the American Statistical Association*, 83, pp. 1093-1107.

Stock, J. and West, K.D. (1988), "Integrated Regressors and Tests of the Permanent Income Hypothesis", *Journal of Monetary Economics*, 21, pp. 85-95.

Stocker, T. (1993), "Empirical Approaches to the Problem of Aggregation over Individuals", *Journal of Economic Literature*, 31, pp. 1827-1874.

Stockman, A.C. (1980), "A Theory of Exchange Rate Determination", *Journal of Political Economy*, 88, pp. 673-698.

Summers, L.H. (1988), "Tax Policy and International Competitiveness", in Frenkel, J.A. (ed.), *International Aspects of Fiscal Policies*, Chicago, University of Chicago Press, pp. 349-375.

Taylor, M.P. (1988), "An Empirical Examination of Long-Run Purchasing Power Parity Using Cointegration Techniques", *Applied Economics*, 20, pp. 1369-1381.

Taylor, M.P. (1991), *Money and Financial Markets*, Cambridge, Mass. and Oxford: Blackwell.

Taylor, M.P. (1995), "The Economics of Exchange Rates", *Journal of Economic Literature*, 33, pp. 13-47.

Taylor, M.P. and McMahon, P.C. (1988), Long-Run Purchasing Power Parity in the 1920s, *European Economic Review*, 32, pp. 179-197.

Taylor, M.P. and Sarno, L. (1997), *The Economics of Exchange Rates*, Cambridge and New York: Cambridge University Press, forthcoming.

Taylor, M.P. and Tonks, I. (1989), "The Internationalisation of Stock Markets and the Abolition of UK Exchange Control", *Review of Economics and Statistics*, 71, pp. 332-336.

Terasvirta, T. (1994), "Specification, Estimation and Evaluation of Smooth Transition Autoregressive Models", *Journal of the American Statistical Association*, 89, pp. 208-218.

Terasvirta, T. (1995), "Smooth Transition Models", paper presented at the 1995 ERNSI Econometric Workshop on Dynamic Models, Oegstgeest.

Terasvirta, T. and Anderson, H. (1992), "Modelling nonlinearities in Business Cycles Using Smooth Transition Autoregressive Models", *Journal of Applied Econometrics*, 7, pp. 119-136.

Tesar, L. and Werner, I. (1995), "US Equity Investment in Emerging Stock Markets", *The World Bank Economic Review*, 9, pp. 109-129.

Tesar, L.L. (1991), "Savings, Investment and International Capital Flows", *Journal of International Economics*, 31, pp. 55-78.

Tesar, L.L. (1993), "International Risk-Sharing and Nontraded Goods", *Journal of International Economics*, 35, pp. 69-89.

Theil, H. (1953), "Repeated Least Squares Applied to Complete Equation Systems", The Hague, Central Planning Bureau, mimeo.

Tobin, J. (1980), *Asset Accumulation and Economic Activity*, Oxford: Basil Blackwell.

Tobin, J. and Dolde, W.C. (1971), "Wealth, Liquidity and Consumption", Reserve Bank of Boston, Monetary Conference Series, 5.

Tong, H. (1990), *Non-linear Time-series: A Dynamic System Approach*, Oxford: Clarendon Press.

United Nations Economic Commission for Latin America and the Caribbean (1994), *Economic Panorama of Latin America, 1994: Argentina, Brazil, Colombia, Chile, Ecuador, Mexico, Peru, Uruguay, Venezuela* United Nations, Santiago.

van Wincoop, E. and Marriman, J. (1993), "Public and Private Saving and Investment", *Institute for International Economic Studies Seminar Paper*, No. 546.

Wei, S., and Parsley, D. (1995), "Purchasing Power Disparity During the Floating Rate Period: Exchange Rate Volatility, Trade Barriers and Other Culrpits", *National Bureau of Economic Research Working paper* No. 5032.

West, K.D., (1988), "Asymptotic Normality when Regressors Have a Unit Root", *Econometrica*, 56, pp. 1397-1417.

White, H. (1980), "A Hetereskedasticity-Consistent Covariance Matrix Estimator and A Direct Test for Heteroskedasticity", *Econometrica*, 48, pp. 817-838.

Williamson, J. (1993), "Issued Posed by Portfolio Investment in Developing Countries", in Claessens, S. and S. Gooptu (1993), op. cit..

Working, H. (1960), "A Note on the Correlation of First Differences of Averages in a Random Chain", *Econometrica*, 28, pp. 916-928.

World Bank (1995), *Trends in Developing Economies*, Washington DC.

World Bank (1997), *The Road to Financial Integration: Private Capital Flows to Developing Countries*, Washington DC.

World Bank (1997), *The Road to Financial Integration: Private Capital Flows to Developing Countries*, Washington DC.

Wu, Y. (1996), "Are Real Exchange Rates Non-Stationary? Evidence from a Panel-Data Test", *Journal of Money, Credit and Banking*, 28, pp. 54-63.

Yule, G.U. (1926), "Why Do We Sometimes Get Nonsense Correlations Between Time Series? A Study in Sampling and the Nature of Time Series", *Journal of the Royal Statistical Society*, 84, pp. 1-64.

Zeldes, S.P. (1989a), "Optimal Consumption with Stochastic Income: Deviations from Certainty Equivalence", *Quarterly Journal of Economics*, 104, pp. 275-297.

Zeldes, S.P. (1989b), "Consumption and Liquidity Constraints: An Empirical Investigation", *Journal of Political Economy*, 97, pp. 305-346.

Zellner, A. (1962), "An Efficient Method of Estimating Seemingly Unrelated Regressions and Tests for Aggregation Bias", *Journal of the American Statistical Association* 57, pp. 348-368.

Zellner, A., Huang, D.S. and Chau, L.C. (1965), "Further Analysis of the Short Run Consumption Function with Emphasis on the Role of Liquid Assets", *Econometrica*, 33, pp. 571-581.