University of Macedonia - Social & Economic Sciences
Department of Economics
Post – Graduate Program in Economic Science

Thesis

"The role of the term structure of interest rates in the transmission mechanism of monetary policy in Greece"

Sarigiannidou Maria

Supervisor Professor: K. Karfakis

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Contents

Introduction

Monetary policy is now at center stage in discussions about how to promote sustainable growth and low inflation in the economy. Fiscal policy has lost its luster as a tool to stabilize the aggregate economy because of doubts about the ability to time fiscal policy actions to obtain desirable stabilization outcomes as well as concerns about budget deficits.

Most economists would agree that, at least in the short-run, monetary policy can significantly influence the course of the real economy. Indeed, a spate of recent empirical research has confirmed the early findings of Friedman and Schwartz (1963) that monetary policy actions are followed by movements in real output that may last for two years or more. There is far less agreement, however, about exactly how monetary policy exerts its influence: the same research that has established that changes in monetary policy are eventually followed by changes in output is largely silent about what happens in the interim. To a great extent, empirical analysis of the effects of monetary policy has treated the monetary transmission mechanism as a "black box".

In the first part of this essay the conventional views of how monetary policy works are readily available. The presentation begins with the *traditional* view, which supports that monetary policymakers use their leverage over short-term interest rates to influence the cost of capital and, consequently, spending on investment. In turn, changes in aggregate demand affect the level of production. Although some researchers, such as Taylor (1995) see traditional interest rate channels operating through the cost of capital as extremely important in the monetary transmission mechanism, many others find it incomplete in several important ways.

The analysis is continued with a reference to *other asset price channels* besides interest rates. The view that asset prices such as stock prices, foreign exchange rates, housing and land prices are an important part of the monetary transmission mechanism is held by monetarists, Keynesians and neo-Keynesians.

Furthermore, important gaps in the conventional story have led a number of economists to explore the so-called *Credit Channel*. They argue that imperfect information and other frictions in credit markets are crucial to understand how monetary policy affects the economy. This credit channel consists of the *Balance Sheet* Channel and the *Bank Lending* Channel. The overview of the main types of

monetary transmission mechanisms ends with a brief reference to a *cost- or supply-side channel* of monetary transmission.

In Part II the analysis turns to a more specific subject of the monetary transmission mechanism: the *term structure of interest rates.* An understanding of the relationship between short and long-term interest rates, the term spread of interest rates is of central importance for the conduct of monetary policy.

A great deal of research has explored the information in the term structure of interest rates. This research examines whether the relationship between interest rates at different maturities helps to explain future movements in interest rates. An attractive theory of the term structure of interest rates is the *expectations hypothesis* that relates the yield on long-term bonds to expected future short rates. A theoretical presentation, as well as the main empirical evidence of the expectations theory are reviewed in Part II. This part ends with the report of the conclusions of some empirical research on the long-run relationship between short - and long-term interest rates. The exploration of such a linkage is constitutes the object of our econometric work.

The third, and final, part of this essay is the econometric part. My aim in this section is to explore whether there exists a stable long-run relationship between a short-term and a medium-term interest rate and a medium-term and a long-term rate. It is examined whether changes induced to a short-term interest rate, which is used as an instrument in the conduct of monetary policy, are transferred to other interest rates with longer duration. The revealment of such a linkage is considered to be an important item in the transmission process and it can be exploitable by monetary authorities in achieving their goals.

More specifically, I estimate empirically, for the case of Greece, (a) the relationship between the one-day rate (R_1) and the Athibor 3-month interest rate (R_2) both of the inter-bank market, and (b) the linkage between the 3-month Athibor interest rate (R_2) and the one-year government bond rate (R_3). The period under study is March 1994 to November 2000.

The empirical findings support the existence of a stable long - run relationship between the daily and 3-month inter-bank rates and the latter and the one-year government bond rate.

Part I. Channels of Monetary Transmission

Section A. Demand Channels of Monetary Transmission

◻ Traditional Interest Rate Channels

We start the overview of the monetary transmission mechanism with interest-rate channels because these have been a standard feature in the literature for over fifty years and are the key monetary transmission mechanism in the basic Keynesian IS-LM model.

The traditional Keynesian view of how a monetary tightening is transmitted to the real economy can be characterized by a schematic diagram:

$$M\downarrow \Rightarrow i\uparrow \Rightarrow I\downarrow \Rightarrow Y\downarrow$$

where $M\downarrow$ indicates a contractionary monetary policy leading to a rise in real interest rates ($i\uparrow$), which in turn raises the cost of capital, thereby causing a decline in investment spending ($I\downarrow$), thereby leading to a decline in aggregate demand and a fall in output ($Y\downarrow$).

Although Keynes originally emphasized this channel as operating through businesses' decisions about investment spending, later research recognized that consumers' decisions about housing and consumer durable expenditure also are investment decisions. Thus, the interest rate channel of monetary transmission outlined in the schematic above applies equally to consumer spending in which I represents residential housing and consumer durable expenditure.

An important feature of the interest rate transmission mechanism is its emphasis on the *real* rather than the nominal interest rate as that which affects consumer and business decisions. Moreover, it is often the real *long-term* interest rate and not the short-term interest rate that has the major impact on spending. A crucial question arising here is how changes in the short-term nominal interest rate result in a corresponding change in the real interest rate on both short and long-term bonds. The answer is sticky prices. A contractionary monetary policy, which increases the short-term nominal interest rate, also increases the short-term real rate and, this result is

consistent even with rational expectations. The expectations hypothesis of the term structure, which states that the long interest rate is an average of expected future short-term interest rates, suggests that the higher real short-term interest rate leads to a rise in the real long-term interest rate. It is these higher real interest rates that affect business fixed investment, residential housing investment, consumer durable expenditure and inventory investment, all of which produce the fall in aggregate output.

The fact that it is the real interest rate that impacts on spending rather than the nominal rate provides an important alternative mechanism for how monetary policy can stimulate the economy, even if nominal interest rates hit a floor of zero during a deflationary episode. With nominal interest rates at a floor of zero, an expansion in the money supply ($M \uparrow$) can raise the expected price level ($P^e \uparrow$) and therefore expected inflation ($\pi^e \uparrow$), thereby lowering the real interest rate ($i_r \downarrow$) and stimulating spending through the interest rate channel. The schematic for this transmission mechanism is thus:

$$M \uparrow \Rightarrow P^e \uparrow \Rightarrow \pi^e \uparrow \Rightarrow i_r \downarrow \Rightarrow I \uparrow \Rightarrow Y \uparrow$$

This mechanism indicates that monetary policy can still be effective even when nominal interest rates have already been driven down to zero by the monetary authorities. Indeed, this mechanism explains why the U.S. economy was not stuck in a liquidity trap during the Great Depression and why expansionary monetary policy could have prevented the sharp decline in output during this period.

Taylor (1995) takes the position that there are strong interest rate effects on consumer and investment spending and, hence, a strong interest rate channel of monetary transmission. His position, however, is a controversial one because many researchers, as for example Bernanke and Gertler (1995), have an alternative view. They state that empirical studies have had great difficulty in identifying quantitatively significant effects of interest rates through the cost capital. This empirical failure of interest rate channels has provided the stimulus for the search for other transmission mechanisms of monetary policy.

◻ Other Asset Price Channels

As emphasized by Allan Meltzer, a key monetarist objection to the Keynesian paradigm for analyzing monetary policy effects on the economy is that is focuses on only one asset price, the interest rate, rather than on many asset prices. Instead, monetarists hold that it is vital to look at how monetary policy affects the universe of relative asset prices and real wealth. Recognition of these other channels is a feature of macro models built by Keynesians such as Franco Modigliani who also see these other asset price effects as being critical to the monetary transmission mechanism.

There are two key assets besides bonds that receive substantial attention in the literature on the transmission mechanism: foreign exchange and equities.

Exchange Rate Channel

With the growing internationalization of economies throughout the world and the advent of flexible exchange rates more attention has been paid to monetary policy transmission operating through exchange rate effects on net exports. This transmission mechanism is now a standard feature in macroeconomics.

This channel also involves interest rate effects because when domestic real interest rates rise, domestic dollar deposits become more attractive relative to deposits denominated in foreign currencies. This leads to a rise in the value of dollar deposits relative to other currency deposits, that is an appreciation of the dollar, which is denoted by $E\uparrow$. The higher value of the domestic currency makes domestic goods more expensive than foreign goods and thereby causes a fall in net exports (NX$\downarrow$) and hence in aggregate output. The schematic for the monetary transmission mechanism operating through the exchange rate is thus:

$$M\downarrow \Rightarrow i\uparrow \Rightarrow E\uparrow \Rightarrow NX\downarrow \Rightarrow Y\downarrow$$

The importance of the exchange rate channel in the way monetary policy affects the domestic economy is emphasized by Taylor, Maurice Obstfeld, Kenneth Rogoff and others. As they state, a framework for the conduct of monetary policy must be inherently international in scope.

Equity Price Channels

Monetarists are often loath to commit themselves to specific transmission mechanisms, because they see these mechanisms as changing during different business cycles. However, two channels are often emphasized in monetarist stories about the monetary transmission mechanism: these involve Tobin's q theory of investment and wealth effects on consumption, as well as housing and land price channels.

i) Tobin's q Theory.

Tobin's q Theory provides a mechanism through which monetary policy affects the economy through its effects on the valuation of equities. Tobin defines q as the market value of firms divided by the replacement cost of capital. If q is high, the market price of firms is high relative to the replacement cost of capital and new plant and equipment capital is cheap relative to the market value of business firms. Companies can then issue equity and get a high price for it relative to the cost of the plant and equipment they are buying. Thus investment spending will rise because firms can buy a lot of new investment goods with only a small issue of equity. On the other hand, the opposite result arrives when q is low. Firms will not purchase new investment goods, as the market value of firms is low relative to the cost of capital. If they want to acquire capital, they can buy another firm cheaply and acquire old capital instead. As a result, investment spending will be low.

The crux of this discussion is that a link exists between Tobin's q and investment spending. The crucial question arising here is how can monetary policy affect equity prices. In a monetarist story, when the money supply falls, the public finds it has less money than it wants and so tries to acquire it by decreasing its spending. One place the place can spend less is in the stock market, decreasing the demand for equities and consequently lowering their prices. A more Keynesian story comes to a similar conclusion because it sees the rise in interest rates coming from contractionary monetary policy making bonds more attractive relative to equities, thereby causing the price of equities to fall. Combining these views with the fact that

lower equity prices ($P_e \downarrow$) will lead to a lower q ($q \downarrow$), and thus to lower investment spending ($I \downarrow$), leads to the following transmission mechanism of monetary policy:

$$M\downarrow \Rightarrow P_e\downarrow \Rightarrow q\downarrow \Rightarrow I\downarrow \Rightarrow Y\downarrow$$

ii) Wealth Effects

An alternative channel for monetary transmission through equity prices occurs through wealth effects on consumption. This channel has been strongly advocated by Franco Modigliani. In his life-cycle model, the lifetime resources of consumers determine consumption spending. The former is made up of human capital, real capital and financial wealth. A major component of financial wealth is common stocks. When stock prices fall, the value of financial wealth decreases, thus decreasing the lifetime resources of consumers, and consumption should fall. Since contractionary monetary policy can lead to a decline in stock prices ($P_e \downarrow$), we have the following transmission mechanism:

$$M\downarrow \Rightarrow P_e\downarrow \Rightarrow \text{wealth}\downarrow \Rightarrow \text{consumption}\downarrow \Rightarrow Y\downarrow$$

iii) Housing and Land Price Channels.

Meltzer emphasizes that asset price effects extend beyond those operating through interest rates, exchange rates and equity prices. He supports that monetary policy can have an important impact on the economy through its effect on land and property values. The two channels of wealth effects and Tobin's q Theory provide support for this view.

Housing and land prices are an extremely important component of wealth. Therefore, a monetary contraction which leads to a decline in land and property values, causes households' wealth to decline, thereby causing a fall in consumption and aggregate output. This can be described by the mechanism above where P_e also represents land and property values.

Tobin's q Theory also applies equally to structures and residential housing so that the schematic outlining the q-theory mechanism applies. A monetary expansion that leads to a rise in house prices raises their prices relative to replacement cost. This

results in a rise in Tobin's q for housing, thereby stimulating its production and hence, aggregate output.

Differences between Keynesians and Monetarists

The Keynesian model within the IS-LM framework gives a very explicit view of the transmission mechanism. Monetary policy through a series of effects on the cost of capital finally affects real output. The key element here is that interest rates are the linkage between the real and monetary sectors.

The monetarists have a somewhat broader view with which, interestingly enough, Keynesians do not quarrel. Monetary policy leads to monetary-sector disequilibrium. If the public has, as a result of an expansionary monetary policy, excess money balances, it will reduce them by increasing its expenditures. These expenditures can include the purchase of financial assets and of real goods or goods. A direct expansion in output can thus result. This monetarist transmission mechanism involves substitution between money and a broad range of assets which includes new goods. The Keynesian mechanism, on the other hand, concentrates on substitution from money into other financial assets. For Keynesians, the resultant changes in rates of return can induce new production, and so the end results are indistinguishable.

The different views of the transmission mechanism are illustrated by the Keynesian and monetarist ideas about the price of money. Since the Keynesian thinks of money as a financial asset, its price is the interest rate. To the monetarist, money is used for expenditure, to buy goods, and thus its price is the inverse of the price level. As it was mentioned above, according to the monetarist version, a monetary expansion leads people to have excess cash balances and therefore increase their demand for assets, securities or real estate and goods. This drives up the prices of assets and causes interest rates to fall. Over time this drives up prices in general and increases expectations of inflation, which can cause interest rates to rise. The spillover into goods demand also leads to a real-sector expansion. It is apparent that the monetarist transmission mechanism emphasizes price effects and also deemphasizes the role of interest rates, which may over time be affected in either direction.

As the monetarists developed a fuller specification of the transmission mechanism, it began to appear more Keynesian. When the Keynesians described the

transmission mechanism in detail it appeared more general and very much like that of the monetarists. The transmission mechanism as a point of theoretical disagreement between monetarists and Keynesians is thus a nonissue. Although there are differences in emphasis in regard to the more important elements of the linkages between the real and financial sectors, the theoretical models are the same. The Keynesian emphasis on financial asset substitutability leads logically to an emphasis on interest rates. It is no surprise that the Keynesians often view monetary policy as the appropriate tool for the determination of interest rates. The monetarists' emphasis on expenditure leads logically to their policy emphasis on the size of the money supply.

▫ Credit Channels

It is difficult to explain the magnitude, timing and composition of the economy's response to monetary policy shocks solely in terms of interest rate (neoclassical cost-of-capital) effects. This dissatisfaction with the conventional stories has led a number of economists to explore whether imperfect information and other "frictions" in credit markets might help explain the potency of monetary policy. The mechanisms collectively known as the credit channel help fill in the gaps in the traditional story.

Ben S. Bernanke and Mark Gertler don't think of the credit channel as a distinct, freestanding alternative to the traditional monetary transmission mechanism, but rather as a set of factors that amplify and propagate conventional interest rate effects. For this reason, in their view, the term "credit channel" is something of a misnomer; the credit channel is an enhancement mechanism, not a truly independent or parallel channel.

Underlying the concept of the credit channel is the following premise: whenever frictions, such as imperfect information or costly enforcement of contracts, interfere with the smooth functioning of financial markets, we expect to observe a wedge between the cost of funds raised externally (for example, through the issuance of imperfectly collateralized debt) and the opportunity cost of internal funds. This wedge, which is called the *external finance premium*, reflects the deadweight costs associated with the principal-agent problem that typically exists between lenders and borrowers. Among the factors reflected in the external finance premium are the

lender's expected costs of evaluation, monitoring and collection; the "lemons" premium that results from the fact that the borrower inevitably has better information about its prospects than does the lender; and the costs of distortions in the borrower's behavior that stem from moral hazard or from restrictions in the contract intended to contain moral hazard (for example, restrictive covenants or collateral requirements).

According to advocates of the credit channel, monetary policy affects not only the general level of interest rates, but also the size of the external finance premium. This complementary movement in the external finance premium may help explain the strength, timing and composition of monetary policy effects better than is possible by reference to interest rates alone.

Two mechanisms have been suggested to explain the link between monetary policy actions and the external finance premium: The *balance sheet channel* (sometimes called the net worth channel) and the *bank lending channel*.

The Balance Sheet Channel

The balance sheet channel is based on the theoretical prediction that the external finance premium facing a borrower should depend on borrower's financial position. In particular, the greater is the borrower's net worth – defined operationally as the sum of her liquid assets and marketable collateral – the lower the external finance premium should be. Intuitively, a stronger financial position (greater net worth) enables a borrower to reduce her potential conflict of interest with the lender, either by self-financing a greater share of her investment project or by offering more collateral to guarantee the liabilities she does issue. Equivalently, a stronger financial position implies that adverse selection and moral hazard problems are lower. This basic insight underlies many real-world arrangements, such as the requirement that borrowers meet certain financial ratios, or that they post collateral.

Since borrowers' financial positions affect the external finance premium, and thus the overall terms of credit that they face, fluctuations in the quality of borrowers' balance sheets similarly should affect their investment and spending decisions. An extensive theoretical literature has exploited this idea to argue that endogenous pro-cyclical movements in borrower balance sheets can amplify and propagate business cycles, a phenomenon that has been referred to as the "financial accelerator". This

approach has been supported by a wide range of empirical work linking balance sheet and cash flow variables to firms' decisions concerning fixed investment, inventories and other factor demands, and to household purchases of durables and housing.

The balance sheet channel of monetary policy arises because shifts in Fed policy affect not only market interest rates but also the financial positions of borrowers, both *directly* and *indirectly*.

For example, a tight monetary policy *directly* weakens borrowers' balance sheets in at least two ways. First, to the extent that borrowers have outstanding short-term or floating-rate debt, rising interest rates directly increase interest expenses, reducing net cash flows and weakening the borrower's financial position. As a result adverse selection and moral hazard problems increase leading to decreased lending to finance investment spending. Because many firms rely heavily on short-term debt to finance inventories and other working capital, this direct effect of monetary policy on business net cash flows is quite important quantitatively. The following schematic can show this balance sheet channel:

$$M\downarrow \Rightarrow i\uparrow \Rightarrow \text{cash flow}\downarrow \Rightarrow \text{adverse selection}\uparrow \text{ and moral hazard}\uparrow$$
$$\Rightarrow \text{lending}\downarrow \Rightarrow I\downarrow \Rightarrow Y\downarrow$$

An important feature of this transmission mechanism is that it is *nominal* interest rates that affect firms' cash flow. Thus this interest rate mechanism differs from the traditional interest rate channel in which it is the real rather than the nominal interest rate that affects investment. Furthermore, the short-term interest rate plays a special role in this transmission mechanism because it is interest payments on short-term rather than long-term debt that typically have the greatest impact on firm cash flow.

Second, rising interest rates are also typically associated with declining asset prices, which among other things shrink the value of the borrower's collateral. The fall in equity prices following a tight monetary policy leads to a decline in firms' net worth and, thereby, to lower levels of investment spending and aggregate demand because of the increase in adverse selection and moral hazard problems. The following schematic can depict this channel:

$$M\downarrow \Rightarrow P_e\downarrow \Rightarrow \text{adverse selection}\uparrow \text{ and moral hazard}\uparrow$$
$$\Rightarrow \text{lending}\downarrow \Rightarrow I\downarrow \Rightarrow Y\downarrow$$

Borio, Kennedy and Prowse (1994) argue that an asset-price boom and bust leading to real fluctuations occurred during the 1980s in a number of major industrialized countries.

Tight monetary policy may also reduce net cash flows and collateral values *indirectly*. Considering the case of a certain firm, a monetary tightening reduces spending by its customers either for cost-of-capital or balance sheet reasons. As a result, the firm's revenues decline while its various costs, including interest and wage payments, do not adjust in the short run. The resulting increase in the "financing gap", which is the difference between the firm's uses and sources of funds, erodes the firm's net worth and creditworthiness over time. This mechanism may explain why the impact of the credit channel on spending, particularly on inventory and investment spending, may persist well beyond the period of the initial monetary tightening.

Another mechanism involving adverse selection through which expansionary monetary policy that lowers interest rates can stimulate aggregate output involves the *credit-rationing* phenomenon. As demonstrated by Stiglitz and Weiss (1981), credit rationing occurs in cases where borrowers are denied loans even when they are willing to pay a higher interest rate. This is because individuals and firms with the riskiest investment projects are exactly the ones who are willing to pay the highest interest rates since, if the high-risk investment succeeds, they will be the primary beneficiaries. Thus, higher interest rates increase the adverse selection problem and lower interest rates reduce it. When expansionary monetary policy lowers interest rates, less risk-prone borrowers are a higher fraction of those demanding loans and thus lenders are more willing to lend, raising both investment and output.

A additional balance-sheet channel operates through monetary policy effects on the general price level. Because debt payments are contractually fixed in nominal terms, an unanticipated rise in the price level lowers the value of firms' liabilities in real terms (decreases the burden of the debt) but should not lower the real value of the firms' assets. Monetary expansion that leads to an unanticipated rise in the price level ($P\uparrow$) therefore raises real net worth, which lowers adverse selection and moral hazard problems, thereby leading to a rise in investment spending and aggregate output as in the schematic below:

$$M\uparrow \Rightarrow \text{unanticipated } P\uparrow \Rightarrow \text{adverse selection}\downarrow \text{ and moral hazard}\downarrow$$
$$\Rightarrow \text{lending}\uparrow \Rightarrow I\uparrow \Rightarrow Y\uparrow$$

The view that unanticipated movements in the price level have important effects on aggregate demand has a long tradition in economics: it is the key feature in the debt-deflation view of the Great Depression espoused by Irving Fisher (1933).

Bank Lending Channel

Beyond its impact on borrowers' balance sheets, monetary policy may also affect the external finance premium by shifting the supply of intermediated credit, particularly loans by commercial banks. This is the Bank Lending Channel.

Banks, which remain the dominant source of intermediated credit in most countries, specialize in overcoming informational problems and other frictions in credit markets. If the supply of bank loans is disrupted for some reason, bank-dependent borrowers (small and medium-sized businesses, for example) may not be literally shut off from credit, but they are virtually certain to incur costs associated with finding a new lender, establishing a credit relationship and so on. Therefore, a reduction in the supply of bank credit, relative to other forms of credit, is likely to increase the external finance premium and to reduce real activity.

That banks play an important role in overcoming informational problems in credit markets, and that as a result many borrowers are substantially bank-dependent, seems well established. The more controversial question about the bank lending channel is whether monetary policy can significantly affect the supply or relative pricing of bank loans. Bernanke and Blinder's (1988) model of the bank lending channel suggested that open market sales by the Fed, which drain reserves and hence deposits from the banking system, would limit the supply of bank loans by reducing banks' access to loanable funds. Schematically, the monetary policy effect is

$$M\downarrow \Rightarrow \text{bank deposits}\downarrow \Rightarrow \text{bank loans}\downarrow \Rightarrow I\downarrow \Rightarrow Y\downarrow$$

A key assumption of the Bernanke and Blinder model is that banks cannot easily replace lost deposits with other sources of funds, such as certificates of deposit or new equity issues. That loans and securities are not perfect substitutes in bank portfolios is empirically plausible. Arguably, this assumption was a good one for the United States prior to 1980, for several reasons. However, since about 1980, as emphasized by Romer (1990), banks' ability to raise funds on the margin has become

less restricted. Markets for bank liabilities have greatly deepened, regulation Q deposit rate ceilings were removed and reserve requirements have been eliminated for most bank liabilities. Clearly, the Bernanke and Blinder model is a poorer description of reality than it used to be, at least for the United States.

Nevertheless, the existence of a bank lending channel does not require banks to be totally incapable of replacing lost deposits. As emphasized by Kashyap and Stein (1994) among others, it is sufficient that banks do not face a perfectly elastic demand for their open-market liabilities, so that an open-market sale by the Fed also increases banks' relative cost of funds. An increase in the cost of funds should shift the supply of loans inward, squeezing out bank-dependent borrowers and raising the external finance premium.

In summary, because of financial deregulation and innovation, the importance of the traditional bank lending channel has most likely diminished over time. Furthermore, the decline of the traditional bank lending business which is occurring worldwide means that banks are playing a less important role in credit markets, rendering the bank lending channel again less potent.

Additionally, since a tightening of monetary policy leads to a worsening of both borrowers' and banks' balance sheet, balance sheet effects alone could conceivably explain why borrowing becomes more expensive and difficult. Because it is extremely difficult to carry out an empirical test that would conclusively separate the bank lending channel from the balance sheet channel, we are more confident in the existence of a credit channel in general.

Household Balance Sheet Effects

It is true that most of the literature on the credit channel focuses on the behavior of business firms. However, the credit market frictions that affect firms should also be relevant to the borrowing and spending decisions made by households, particularly spending on costly durable items such as automobiles and houses. Thus the workings of the credit channel may help explain the high sensitivity of residential investment and consumer durables to monetary policy shocks.

It is likely that both the bank lending channels and the balance sheet channel have at various times affected residential investment, which accounts for the largest fraction of the decline in economic activity immediately following a monetary

tightening. Declines in bank lending induced by a monetary contraction should cause a decline in durables and housing purchases by consumers who do not have access to other sources of credit. Similarly, increases in interest rates cause a deterioration in household balance sheets because consumers' cash flow is adversely affected.

Another way of looking at how the balance sheet channel may operate through consumers is to consider *liquidity effects* on consumer durable and housing expenditure, which were found to have been important factors during the Great Depression. In the liquidity-effects view, balance sheet effects work through their impact on consumers' desire to spend rather than on lenders' desire to lend. Consumer durables and housing are very illiquid assets. If as a result of a bad income shock consumers needed to sell their consumer durables or housing to raise money, they would expect a big loss because they could not get the full value of these assets in a distress sale. On the contrary, if consumers held financial assets, such as money in the bank, stocks or bonds, they could easily sell them quickly for their full market value and raise the cash. Hence, if consumers expect a higher likelihood of finding themselves in financial distress, they would rather be holding fewer illiquid consumer durable or housing assets and more liquid financial assets.

But what influences the consumers' estimate of the probability of suffering financial distress? Among others, a consumer's balance sheet is an important factor. Specifically, when consumers have a large amount of financial assets relative to their debts, their estimate of the likelihood of financial distress is low, and they will be more willing to purchase consumer durables or housing. When stock prices rise, the value of financial assets rises as well. Therefore, consumer durable expenditure will also rise because consumers have a more secure financial position and a lower estimate of the likelihood of suffering financial distress. This leads to another transmission mechanism for monetary policy operating through the link between money and equity prices:

$$M\uparrow \Rightarrow P_e\uparrow \Rightarrow \text{financial assets}\uparrow \Rightarrow \text{likelihood of financial distress}\downarrow$$
$$\Rightarrow \text{consumer durable and housing expenditure}\uparrow \Rightarrow Y\uparrow$$

The illiquidity of consumer durable and housing assets provides another reason why a monetary contraction which raises interest rates and thereby reduces cash flow to consumers leads to a decline in spending on consumer durables and

housing. A decline in consumer cash flow increases the likelihood of financial distress, which reduces the desire of consumers to hold durable goods or housing, thus reducing spending on them and hence aggregate output. The only difference between this view of cash flow effects and that of the balance sheet channel is that it is not the unwillingness of lenders to lend to consumers that causes expenditure to decline, but the unwillingness of consumers to spend.

Section B. The Cost-Channel of Monetary Transmission

Traditional economic models posit that changes in monetary policy exert an effect upon the economy through a demand channel of transmission. Alternatively, some researchers have proposed that there may be important supply-side, or cost-side effects of monetary policy, e.g. Galbraith (1969, Shapiro (1981), Christiano & Eichenbaum (1992).

Research on industry level suggests that supply-side channels are powerful collaborators in transmitting the real effects of monetary policy in the short-run. In fact, for many important manufacturing industries, a cost channel is the primary mechanism of transmission for the first couple of years after a monetary shock.

A problem with the demand-only view of monetary transmission, noted be Bernanke and Gertler (1995), is the degree of amplification. Empirical evidence suggests that monetary policy shocks that induce relatively small movements in market interest rates have large effects on output. Bernanke and Gertler use this result to support their argument that a credit channel working in tandem with the traditional monetary channel explains this data better. An alternative or complementary means to explain this smplification is to allow monetary policy shocks to have both supply-side and demand-side effects. If this is the case, then a shock to monetary policy could be viewed as shifting both the aggregate supply and aggregate demand curves in the same direction, leading to a large change in output accompanied by a small change in prices.

The literature offers several theoretical foundations for monetary policy as a cost shock. For example, Bernanke and Gertler (1989) model contains both a demand and supply component of balance sheet effects. Several other credit channel papers suggest that there might be a cost-side channel of monetary policy, e.g. Kashyap,

Lamont and Stein (1994). There are several other examples of general equilibrium macroeconomic models that explicitly analyze the supply-side effects of monetary policy through working capital. Blinder (1987), Christiano and Eichenbaum (1992) and Farmer (1988) all begin with the assumption that firms must pay their factors of production before they receive revenues from sales, and must borrow to finance these payments. In most of these models, an increase in the nominal interest rate serves to raise production costs. Thus, a monetary contraction leads to a decline in output through an effect on supply. It is important to note that some type of rigidity is still required for money to be non-neutral. If prices and portfolios adjust immediately, then monetary policy has no initial effect on interest rates, so that neither aggregate demand nor aggregate supply shifts.

A clear framework for distinguishing between supply and demand-side effects of monetary policy is necessary. Blinder (1987), in exploring the theoretical foundations of this issue, notes that a very simple test of supply-side effects is the following: rising prices in the presence of falling output is taken as indication of a cost shock, while a positive correlation between these two quantities identifies a demand shock.

These results imply that if a monetary contraction affects an industry primarily by lowering the demand for its product, we should observe a fall in both output and prices relative to wages. In contrast, if a monetary contraction affects an industry primarily by raising its production costs, we should observe a fall in output accompanied be a rise in prices relative to wages. If a monetary contraction affects both supply and demand, then we should observe an amplified fall in output and an ambiguous effect on the price-wage ratio. Therefore, where this test becomes uninformative is in the presence of both demand and cost shocks of relatively equal strength. In this case, the effect on prices is indeterminate.

It should be clear that this model makes no predictions about the aggregate price level since it does not specify the source of non-neutrality of money. The only price implications are for relative prices.

The important implication is not that unexpected changes in monetary policy are a cost shock or a demand shock, but rather that it is a combination of a cost and a demand shock. This hypothesis is not only better able to explain the stylized facts than sticky price models, but it is able to explain the evidence on a microeconomic level, where traditional models fail.

Part II. The Term Structure of Interest Rates

It is commonly believed that while central banks closely control short-term interest rates, long-term interest rates, which are heavily influenced by expectations, play a more important role in affecting aggregate demand. While monetary authorities conduct monetary policy by setting or influencing very short term money-market interest rates, the interest rates that firms and consumers face when they borrow or lend tend to be tied to longer-term interest rates outside the direct control of the central bank. Central banks thus face the problem of setting short-term interest rates in such a way as to move longer-term interest rates in the desired direction. An understanding of the relationship between short and long-term interest rates, the *term structure of interest rates* is thus of central importance for the conduct of monetary policy.

The term structure of interest rates plays an important role in much of modern financial economics and in the design of monetary policy for another reason, also: the term structure may provide useful information concerning the private sector's expectations of future monetary policy, inflation and economic activity. As a result, the determination of interest rates of various maturities constitutes an important research area for central banks. The fact that the term spread is useful for predicting future macroeconomic conditions suggests that it is a good candidate for use as a monetary policy indicator.

Unfortunately, the standard conclusion in the extensive empirical literature on the term structure suggests that current theoretical models do not explain movements in short- and long-term interest rates very well. Robert Shiller, for instance, in his exhaustive survey on the term structure of interest rates concludes that: "Empirical work on the term structure has produced consensus on little more than that the rational expectations model...can be rejected". While this conclusion comes as no surprise to monetary policy-makers, it suggests that, if correct, little use can be made of the term structure in the design of monetary policy.

Section A. The Expectations Hypothesis

Theoretical Presentation

A great deal of research has explored the information in the term structure of interest rates. This research examines whether the relationship between interest rates at different maturities helps to explain future movements in interest rates.

An attractive theory of the term structure of interest rates is the *expectations hypothesis* that relates the yield on long-term bonds to expected future short rates. Three versions of expectations theory exist in the literature: the pure expectations hypothesis, which does not allow for a risk premium, the constant risk premium expectations hypothesis, which allows for a maturity-specific risk premium that is constant over time, and a generalization of the expectations hypothesis in which the maturity-specific risk premium is time varying. Consistent with evidence from time series studies, both the pure and the constant risk premium expectations theories are strongly rejected. However, once allowance is made for a time varying risk premium, the expectations hypothesis cannot be rejected.

We will briefly review the expectations theory for the two-period case, i.e. for the relationship between yields on one-period and two-period bonds denoted r_t and R_t respectively. The expectations theory of the term structure posits that the "long" rate R_t is related to r_t and the expected future short rate $E_t r_{t+1}$ as follows:

$$R_t = 0.5(r_t + E_t r_{t+1}) + \xi_t \qquad \text{(II.1)}$$

Here

$$E_t r_{t+1} = E(r_{t+1} | \Omega_t)$$

with $\Omega_t = \{r_t, r_{t+1}, \ldots, R_t, R_{t+1}, \ldots\}$ so we are assuming rational expectations. The term ξ_t is a "term premium" that is often assumed constant. Equation (II.1) states that the current two-period yield is an average of the current one-period yield and the expected one-period yield, plus a constant term premium. The return from investing in a two-period bill equals, up to a constant, the expected return from investing sequentially in two one-period bills.

Defining the expectational error $\varepsilon_{t+1} = r_{t+1} - E_t r_{t+1}$, equation (II.1) implies

$$\frac{1}{2}(r_{t+1} - r_t) = (R_t - r_t) - \xi_t + \frac{1}{2}\varepsilon_{t+1} \qquad \text{(II.2)}$$

Then if ξ_t is assumed constant, $\xi_t = \xi$, the orthogonality of ε_{t+1} with R_t and r_t implies that the slope coefficient β in a regression of the form

$$\frac{1}{2}(r_t - r_{t-1}) = \alpha + \beta(R_{t-1} - r_{t-1}) + disturbance \qquad \text{(II.3)}$$

should have a probability limit of 1.0. An estimated value significantly different from 1.0 is inconsistent either with the expectations theory or one of the maintained hypotheses. Equation (II.3) implies that when the long rate rises relative to the short rate, future short rates tend to increase. Such predictive power is consistent with the expectations hypothesis of the term structure.

In fact, it has been documented by many researchers that slope coefficients tend to be well below 1.0 in post – 1914 data for the United States. The only exceptions are of Mankiw and Miron's (1986) value for 1890-1914 and of Campbell and Shiller's (1991) value for 1952-1987, for an exceedingly long short-term rate.

One possible explanation for these findings is, of course, that the expectations theory is simply untrue. Another possibility is invalidity of the rational expectations (RE) hypothesis. This possibility has been explored, using survey data on expectations, by Froot (1989). However, it seems unlikely that the same general type of systematic expectational error would prevail over different sample periods. In any event, B. T. McCallum's proposed explanation is that ξ_t is not constant, i.e., there is a variable term premium and that monetary policy is conducted in a manner to be explained momentarily.

B. MacCallum's explanation of the expectations hypothesis

McCallum suggests that the process generating ξ_t is assumed to be covariance stationary but not necessarily white noise. For specificity, the ξ_t process will be taken to be autoregressive of order one [AR (1)]:

$$\xi_t = \rho\xi_{t-1} + u_t \qquad \text{(II.4)}$$

Here u_t is white noise and $|\rho| < 1.0$. To this writer it seems implausible that there would not be some period-to-period variability in the discrepancy term ξ_t in (II.1), a term that reflects changes in tastes regarding the need for financial flexibility and a myriad of other disturbing influences, none major enough to justify separate recognition.

Regarding monetary policy, McCallum begins with the observation that actual policy behavior in the US and many other nations involves manipulation of a short-term interest rate "instrument" or "operating variable". Specifically he assumes that

$$r_t = \sigma r_{t-1} + \lambda(R_t - r_t) + \zeta_t \qquad \text{(II.5)}$$

where $\sigma \geq 0$ is presumed to be close to 1.0 and $\lambda \geq 0$ to be smaller than 2. Thus there is a considerable element of interest rate "smoothing" – keeping r_t close to r_{t-1} – and also a tendency to tighten policy (by raising r_t) whenever the spread $R_t - r_t$ is higher than normal. This reaction to $R_t - r_t$ occurs because the central bank views it either as a good predictor of future output growth or as a good indicator of recent policy laxity. The final term ζ_t reflects other components of policy behavior and is assumed to be white noise.

It is necessary to briefly consider the rationale for the specification of policy behavior in (II.5). Regarding the r_{t-1} term, there exists some controversy regarding the reason behind central banks' proclivity for interest rate smoothing. But there is virtually no disagreement with the proposition that the Fed and other major central banks have in fact employed such practices during most, if not all, of the last 40 years. In addition (II.5) reflects the assumption that the central bank tends to tighten policy when the spread $R_t - r_t$ is large. One possible rationalization is that the spread is an indicator of monetary policy expansiveness, as suggested by Laurent (1988), so that an unusually high value indicates the need for corrective action. A different idea is that the spread provides an indicator of the state of the economy from a cyclical perspective. Various investigators, including Estrella and Hardouvellis (1991) and Hu (1993), have documented that spread measures have predictive value for future real GNP growth rates. Also, Mishkin (1990) has shown that a spread variable has some predictive content for future inflation rates. Thus an attempt by the central bank to conduct a forward-looking counter-cyclical policy would call for a response of the type indicated in (II.5), i.e. a tightening when $R_t - r_t$ is high. Goodfriend (1993) also

suggests a policy behavior pattern in his pattern which has a substantial degree of similarity with formulation (II.5): both call for an increase in the short rate in response to a ceteris paribus rise in the long rate.

Relations (II.1) and (II.5) constitute only a portion, of course, of a macroeconomic system. But if we assume that the disturbances ξ_t and ζ_t are independent of those in the remaining relations, the system will be recursive and the subsystem (II.1)-(II.5) will determine r_t and R_t without reference to the other variables or shocks. MacCallum considers then a rational expectations solution to the system (II.1)-(II.5) and finds that the spread obeys the equation:

$$R_t - r_t = \frac{1}{2}\left(E_t r_{t+1} - r_t\right) + \xi_t = \left(1 - \frac{\lambda\rho}{2}\right)^{-1} \xi_t \qquad \text{(II.6)}$$

Finally, the obtained solution is:

$$\frac{1}{2}\left(r_t - r_{t-1}\right) = \frac{\lambda\rho}{2}\left(R_{t-1} - r_{t-1}\right) + \frac{\lambda/2}{1 - \rho\lambda/2} u_t + \frac{1}{2}\zeta_t \qquad \text{(II.7)}$$

Here u_t and ζ_t are uncorrelated with $R_{t-1} - r_{t-1}$, so (II.6) represents a population version of the regression described in (II.3). Thus the slope coefficient in (II.3) is a consistent estimator of $\lambda\rho/2$, so the analyst should anticipate a slope well below 1.0. Indeed, if ξ_t were white noise, with ρ=0, a slope coefficient of zero would be implied, even though relation (II.1) is the main behavioral relation of the system. This result demonstrates, as Bennett McCallum would suggest, not only that the usual regression test is inappropriate but also that it is misleading to think of the expectations theory in terms of the "predictive content" of the spread for future changes of the short rate. Such predictive content is not a necessary implication of that theory.

In addition, a zero slope coefficient would be implied if λ=0, i.e., if the central bank did not respond to the current value of the spread but simply set r_t equal to r_{t-1} (plus, perhaps, ζ_t). This special case, of the special case with σ = 1, represents the hypothesis of Mankiw and Miron (1986) – that the Federal Reserve has practiced interest rate smoothing and thereby induced short rates to approximate a random walk process in their behavior. The result of B. MacCallum strongly supports the general idea of the Mankiw and Miron hypothesis, but shows that it holds even if r_t behavior is not that of a random walk.

In the explanation provided above, there is, however, a substantial difficulty that needs to be mentioned. It is the case that actual central banks do not respond only to term spreads in deciding upon changes in r_t. Thus equation (II.5) represents a simplification relative to actual behavior of the Fed and other major central banks, which almost certainly respond to recent inflation and output or employment movements as well as the spread. So, if one were to attempt to econometrically estimate actual reaction functions, then measures of inflation and output gaps would need to be included. In short, this type of study would require specification and estimation of a complete dynamic macroeconometric model. Because of the simplified nature of policy equation (II.5), McCallum' proposed explanation might be regarded as more of a parable than a fully – worked - out quantitative model. However, the main issue is whether a proposed parable is fruitful in understanding important economic phenomena.

As a conclusion, B.T.McCallum in his study of the apparent failure of the expectations theory of the term structure of interest rates (1994) argues that the failure of short-rate (and long-rate) changes to be related as predicted to prevailing long-short spreads is shown to be plausible consequence of monetary policy behavior that features interest rate smoothing in combination with policy responses to movements in the long-short spread. This explanations is entirely consistent with, but more general and more fully developed than, the one proposed in a notable study by Mankiw and Miron (1986).

Empirical Evidence on the Expectations Hypothesis

One problem that has hampered empirical work on the *expectations theory of the term structure* is that in fact there are many different versions of the expectations theory, as emphasized by Cox, Ingersoll and Ross (1981) and others. However, these different expectations theories are in important respects very similar, and are all closely approximated by a single linear expectations theory (Shiller (1979), Campbell (1986)). Almost all studies statistically reject the expectations theory of the term structure. Recently, however, a considerable amount of empirical evidence has been presented suggesting that the explanatory power of the *expectations hypothesis* may be greater than previously thought.

Fama (1984) studies the behavior of interest rates with maturities of 1 to 6 months (specifically, one- to six-month Treasury bills) from 1959 through 1982 and, although he rejects the expectations hypothesis, he finds predictive power in forward rates that lasts about three to five months during the first half of his sample and one month during the second half of his sample. In addition, forward rates can also contain premiums for the risks in the returns on multi-period bonds. Fama (1976) suggests that variation through time in these premiums can obscure the power of forward rates as predictors of future spot rates. On the contrary, the existing literature, as for example Shiller, Campbell and Schoenholtz (1982), generally finds that forward rates, unadjusted for variation in premiums, are poor forecasts of future spot rates.

In addition, *Mankiw and Miron* (1986) study the behavior of 3 and 6-month interest rates over the period 1890-1979 and show that the expectations hypothesis receives considerable support in data from the period before the founding of the Federal Reserve System in 1913, but not thereafter. Indeed it was mentioned above that the slope coefficient (β) for the years 1890-1914 was closer (than for more recent periods) to the value of 1.0. Mankiw and Miron emphasize that, as those years precede the founding of the Federal Reserve System they pertain to a period during which interest rate smoothing behavior would be absent. Furthermore, Mankiw and Miron show that changes in short-term rates were more predictable before the founding of the Fed and argue that, for econometric reasons, regression tests of the expectations hypothesis are likely to be more informative, the more predictable short rates are. If changes in interest rates are more predictable, spreads between long and short interest rates contain more information about future short rates, and the econometric power of term structure regressions increases. This argument suggests that the poor performance of the expectations hypothesis is an econometric problem: with short rates difficult to predict, tests of the expectations hypothesis are always likely to yield poor results.

In a similar vein, *Kugler* (1988, 1990) finds that slope coefficients are closer to 1.0 for Germany and Switzerland than for the United States during recent years. This result he attributes to a smaller degree of interest smoothing behavior by the Bundesbank and the Swiss National Bank, in comparison with the Fed.

A second line of inquiry is followed by *Campbell and Shiller* (1987, 1991), who study the behavior of short and long rates and find that, although the expectations

hypothesis is rejected on *statistical grounds*, on *economic* grounds the rejection is less obvious. Estimating bivariate VARs for short rates and term spreads, Campbell and Shiller show that, although statistically rejected, long rates behave in a way strikingly similar to that implied by the expectations hypothesis. The expectations hypothesis thus remains a useful benchmark.

In their paper (1991) Campbell and Shiller provide a summary characterization of term structure behavior that is consistent with their battery of empirical findings. The two summary characterizations considered are (for the two-period case):

$$R_t - r_t = 0.5E_t(r_{t+1} - r_t) + c + v_t \qquad \text{(II.8)}$$

where v_t is added noise that is orthogonal to $E_t r_{t+1} - r_t$, and

$$R_t - r_t = k0.5E_t(r_{t+1} - r_t) + c \qquad \text{(II.9)}$$

where $k>1$. The latter "could be described as an overreaction model of the yield spread" according to Campbell and Shiller (1991, p.513). A rise in the long rate relative to the short rate is due to the expectation of higher short rates in the future. Long rates overreact to future short rates (or underreact to current short rates). This explanation assumes that risk premia are constant and that the spread between long and short rates correctly incorporates the information about expectations of future short rates.

They explore the implications of these two summary characterizations of ways in which the expectations theory might fail and conclude that (II.9) is consistent with the data but that (II.8) is not. Considering how these characterizations compare with McCallum's explanation, presented above, we see that equation (II.6) is of a similar form to that of (II.8), with the only difference that ξ_t in (II.6) is not orthogonal to $E_t r_{t+1} - r_t$. Furthermore, eliminating ξ_t from (II.6) results in

$$R_t - r_t = \frac{1}{\rho\lambda} E_t(r_{t+1} - r_t) \qquad \text{(II.10)}$$

for McCallum's model. But with $0<\lambda<2$ and $|\rho| < 1$, (II.10) implies that $k>1$ in (II.9) if ρ is positive. So Campbell and Shiller's summary characterization is consistent with the MacCallum's rationalization.

Finally, the expectations hypothesis receives considerable support in *Hardouvelis* (1994), who studies the behavior of 3-month and 10-year rates in the G-7 countries using data between 20 and 40 years of data depending on the country. One finding of particular interest is that the expectations hypothesis appears to do a much better job of accounting for the behavior of interest rates outside the United States. Furthermore, Hardouvelis shows that multi-period regressions – in which cumulative future returns on 3-month investments are regressed on the current spread between long and short rates – are supportive of the expectations hypothesis when longer investment horizons are considered. To the extent that it is easier to predict changes in short rates over longer forecast horizons, this evidence is also compatible with the Mankiw-Miron hypothesis that the expectations hypothesis may fare badly in empirical work because short rates are difficult to predict, and not because the theory is fundamentally flawed.

Since the predictability of short rates is likely to vary between currencies, one would expect that using data for several currencies with potentially substantial differences in the time series behavior and predictability of short rates would shed further light on the expectations hypothesis. Stefan Gerlach and Frank Smets (1995) pursue this reasoning and use Euro-rates to study the behavior of 1, 3, 6 and 12-month rates for a sample of 17 currencies. The data provide considerable support for the expectations hypothesis of the term structure at the short end (less than one year) of the maturity spectrum. They find that for all 17 currencies in their data set the term spread does predict future movements in the short rate. Furthermore, the results reinforce the conclusion in Mankiw and Miron (1986) that the easier it is to forecast future short rates, the better the data conform to the implications of the expectations hypothesis. They also show that the failure of the expectations theory to hold in a number of countries, most notably the United States, is due to a lack of predictability of the short-term rate which introduces a downward bias in the estimates in the presence of a time-varying term premium.

Section B. Instruments of Monetary Policy and Their Effect on the Term Structure

In this section we will examine the relationship between the term structure spread and a direct instrument of monetary policy. An efficacious tightening of monetary policy should have differential effects on short- and long-term interest rates. At the short end, the predominant effect is a tightening in the supply of credit, leading to a rise in interest rates. The long end is driven to a much greater extent by changes in expected inflation and in the real ex ante long-term rate. If the tightening is viewed as credible and effective, reduced long-term inflationary expectations should moderate the effect of tighter initial credit conditions. The combined result is that long-term rates tend to rise by less than short-term rates (they could conceivably decline), and the spread between long and short term rates declines. In other words the yield curve "flattens".

Nevertheless, it is also possible that further increases in the short-term interest rate may be expected over a long future period or that current increases are viewed as insufficient to control inflation and drive down inflationary expectations. In either case, the long-term rate may rise as much or more than the short-term rate and spread between the long and the short would therefore not decline. Arturo Estrella and Frederic S. Mishkin (1995) performed an analysis exploring these types of phenomena.

In their analysis, they identify a short-term instrument as one that is clearly under the central bank's control and that is customarily (or simply frequently) used to implement changes in monetary policy stance. In each of the five countries they studied (United States, United Kingdom, Germany, France and Italy), there is a short-term interest rate that serves these purposes over the observation period 1973-1995. Their empirical results are clearly supportive of a relationship along the lines described earlier. Specifically, an unexpected change in the central bank rate leads to a flattening in the yield curve for domestic government securities. The extent of the flattening, however and the explanatory power of the results vary from country to country.

It should be noted that Estrella and Mishkin are essentially examining the relationship between two endogenous variables. In setting the central bank rate, the

monetary authority is influenced by past economic indicators as well as by expectations of future economic variables. For these and other reasons, it would not be appropriate to think of the central bank rate as a purely exogenous variable. To partially deal with the endogeneity problem the authors adopt a vector autoregressive (VAR) formulation for the central bank, short- and long-term interest rates. The empirical estimates show that although the average direction of the effect is very consistent, results in specific cases are bound to vary considerably. In view of these results, it is hard to argue that the central bank can "control" the term structure spread with operations at the very short end.

The term structure as a predictor of future real activity

The significance of the informational content of the term structure in predicting real activity has been documented for the United States in papers by Harvey (1988), Laurent (1988, 1989), Chen (1991), Estrella and Hardouvelis (1991), Bomhoff (1994), Davis and Henry (1994) and Barran (1995).

Why should such a predictive relationship exist? One possibility – the common factor's explanation – is that both the term structure and future real activity are determined by current monetary policy. Tight monetary policy would tend both to flatten the yield curve and lead to a slowdown in activity. However, even though the yield curve spread seems to be influenced by current monetary policy, to say that it is determined is a gross overstatement.

In addition, if the common factor were the only explanation, the predictive power of the term structure should dissipate when variables representing current policy are added. As seen below, this is generally not the case. It is interesting nevertheless, to examine the significance of the yield curve when alternative monetary variables are added to the predictive equation.

From a purely theoretical point of view, the yield curve spread may be related positively or negatively to future real output. The common factor explanation suggests a positive relationship, which is typical in previous empirical research. Explanations based on real demand shocks are also consistent with a positive relationship, the flavor of which may be conveyed by thinking of simple future shifts in the IS curve. A more elaborate but suggestive formal model based on the consumption capital asset

pricing model is found in Harvey (1988). On the other hand, expectations of future monetary tightening could be associated both with higher interest rates and lower output, especially in the short run, and this could be thought of as future shifts in the LM curve.

The general strategy, A. Estrella and F. S. Mishkin (1995) use, is to estimate a regression equation in which the contemporaneous value of the SPREAD (= 10-year government bond – 3-month government bill rate) is used to forecast the change in real economic activity over the following k periods. Their analysis refers to five countries: United States, United Kingdom, Germany, France and Italy for the period 1973-1995. The basic equation is

$$y_t^k = \alpha_0 + \alpha_1 \, SPREAD_t + \varepsilon_t$$

where y_t^k is one of various measures of the change in economic activity, such as a log change in real GDP. Lagged values of quarterly GDP growth are omitted from this equation because, empirically, such lags are generally insignificant.

There is very consistent evidence that the above relationship is positive. Furthermore, with the exception of Italy, the results tend to be very significant, especially for horizons of 4 to 8 quarters ahead.

A somewhat different approach involves the prediction of whether or not the economy will be in a recession k quarters ahead. The *SPREAD* is the only explanatory variable in a probit regression in which the dependent variable (*RECESSION*) is a dummy that equals 1 if the economy is in recession four quarters ahead and equals 0 otherwise. Because of the limited dependent variable, the probit regression is nonlinear and has the form:

$$P\,(RECESSION_t = 1) = F(\alpha_0 + \alpha_1 \, SPREAD_{t-1})$$

where F is the normal cumulative distribution function. As was the case with the quantitative dependent variables, the results are generally very good. The estimates are statistically significant. The economic significance is harder to gauge, since the relationship between the linear combination with which the parameter is associated and the probability of a recession is highly nonlinear. Even though the coefficient is

constant, the effect of an increase of 1 percentage point in the *SPREAD* will be very different for different levels of the *SPREAD*.

The role of other monetary policy variables

We referred above to the possibility that the predictive power of the term structure for real activity may be attributable to the influence of monetary policy on both. We now take a look at whether there is predictive power of the term structure over and above that provided by other variables that reflect the stance of monetary policy. A. Estrella and F. S. Mishkin (1995) introduce those other monetary policy variables in the equation for predicting future growth in real GDP and examine the significance of both the monetary variable and the *SPREAD*. The form of the regression equation is

$$y_t^k = \alpha_0 + \alpha_1 \, SPREAD_t + \alpha_2 \, \xi_t + \varepsilon_t$$

where ξ_t is the contemporaneous measure of monetary policy. In very general terms, the predictive power of the term structure *SPREAD* remains when other monetary policy variables are introduced.

In summary, the term structure *SPREAD* by itself is useful in predicting real economic activity, especially between 4 and 8 quarters ahead, independently of which measure of activity is used. Moreover, the predictive power does not seem to be attributable solely or primarily to known information about other monetary policy variables.

Section C. Co-movement of Interest Rates: Recent Evidence

The deregulation of financial markets in the 1980s led to a movement away from the targeting of monetary growth towards the targeting of interest rates as the principal conduit of monetary policy. By controlling the level of interest rates, authorities hope to affect the level of business investment and the flow-on to the real economy. While this theoretical relationship between a "representative" interest rate and the real economy is well understood, the role of central banks in influencing this

rate has typically been assumed and has not been examined a lot in the empirical literature. It is generally recognized that economic activity responds to movements in long-term interest rates while monetary authorities use a short-term interest rate as an operating instrument for monetary policy. If this is the case, the term and risk structures of interest rates become important in the transmission process.

Graham Elliot and Ronald Bewley (1994) study the relationship between two key Australian short-term interest rates: the official and the unofficial overnight cash rates. They examine the extent to which innovations in the official rate are transmitted to the unofficial rate and the extent to which this transmission is permanent. Furthermore, they test whether there is any feedback from the unofficial rate to the official. The results of their research support the existence of a stable long-run relationship between the official and unofficial interest rates. Moreover, the change in the official rate is found to react to the past differential between official and unofficial rates implying the existence of some feedback relationship. The finding of such a relationship between the two interest rates is important given the crucial role it plays in the transmission mechanism. The official rate is fully controlled by the Reserve Bank of Australia; the long-run stable relationship between the official and unofficial rates means that the authorities can effectively target unofficial interest rates.

However, this is a link between interest rates only at the shortest end of the maturity spectrum. In order monetary policy to affect economic activity, it is necessary that the targeted interest rates have predictable influence on longer term interest rates. Karfakis and Moschos (1995) examine the links between the official cash rate, the 13-week Treasury note rate and two-, five-, and ten-year Treasury bond rates for the case of Australia. They reach the conclusion that because the expectations theory linking short and long rates seems to hold, the Reserve Bank of Australia can influence the Treasury note rate by intervening in the official cash market.

Because this research is emphasizing on the relationships between the official cash rate and rates on government paper of varying maturities, it neglects the links between the rates on government paper and those on private paper, the rates at which the private sector can borrow. A paper written by Lim and Martin (1994) explores exactly the dynamic interrelationships between short-term interest rates on both private and government paper. Their conclusions support such relationships again for the case of Australia.

Part III. Econometric Part

Section A. Methodology

- **The model**

This section of the essay is focused on the investigation of the relationship between two interest rates, with different duration and possibly belonging to a different market. Our aim is to explore whether there exists a stable long-run relationship between a short-term and a medium-term interest rate and a medium-term and a long-term rate. In other words, the field of interest is whether changes induced to a short-term interest rate, which is used as an instrument in the conduct of monetary policy, are transferred to other interest rates with longer duration. The validity of such a linkage between short- and longer-term interest rates is very important in the transmission process and it can be exploitable by monetary authorities in achieving their goals.

In particular, we will examine

a) the relationship between the one-day rate (R_1) and the Athibor 3-month interest rate (R_2) both of the inter-bank market, and
b) the linkage between the 3-month Athibor interest rate (R_2) and the one-year government bond rate (R_3).

Specifically, we will estimate empirically the following equations:

$$R_2(t) = \alpha + b\, R_1\,(t) + u_t \qquad \text{(III.1)}$$

where R_1 is the one-day interest rate and R_2 the 3-month Athibor rate, both of inter-bank market, while u_t is an error term.

In a similar way, we examine the relationship between the 3-month Athibor inter-bank rate (R_2) and the one-year government bond rate (R_3) by estimating the equation:

$$R_3(t) = \alpha + b\, R_2\,(t) + u_t \qquad \text{(III.2)}$$

Where R_3 is the one-year government bond interest rate. The existence of a long-run relationship between R_1 and R_2, and R_2 and R_3 interest rates is explored by means of

cointegration techniques. More specifically, the Johansen's method of maximum likelihood estimation of cointegrated systems is implemented.

- **Theoretical Issues on Integration Analysis**

Any time series data can be though of as being generated by a *stochastic* or *random process*. A type of stochastic process that has received a great deal of attention and scrutiny by time series analysts is the so – called *stationary stochastic process*.

Broadly speaking, a stochastic process is said to be stationary if its mean and variance are constant over time and the value of covariance between two time periods depends only on the distance or lag between the two time periods and not on the actual time at which the covariance is computed. Let Y_t be a stochastic time series with the following properties:

$$E(Y_t) = \mu \qquad \text{(III.3)}$$

$$Var(Y_t) = E[(Y_t-\mu)^2] = \sigma^2 \qquad \text{(III.4)}$$

$$\gamma_k = E[(Y_t - \mu)(Y_{t+k} - \mu)] \qquad \text{(III.5)}$$

where γ_k, the covariance (or autocovariance) at lag k, is the covariance between the values of Y_t and Y_{t+k}, that is between two Y values k periods part. If a time series is stationary, its mean, variance and autocovariances of Y_{t+k} must be the same as those of Y_t.

If a time series is not stationary in the sense just defined it is called a *nonstationary time series*.

A common practice to examine whether a time series is stationary is based on the so – called *autocorrelation function (ACF)*. The ACF at lag k, denoted by ρ_k, is defined as

$$\rho_k = \gamma_k / \gamma_0 \qquad \text{(III.6)}$$

The plot of the sample autocorrelation function ($\rho_k = \gamma_k / \gamma_0$) against k is known as the *sample correlogram*. For a stationary time series, the autocovariances (and of course

the correlogram) dies down rapidly. This is because in a purely random stochastic process the autocorrelation at any lag greater than zero is zero. However, correlogram output is not used as a formal test for stationarity in literature.

The most formal and popular tests for stationarity are known as the *unit root tests*. In order to introduce this test we consider the following model:

$$Y_t = Y_{t-1} + u_t \qquad \text{(III.7)}$$

where u_t is the stochastic error term that follows the classical assumptions, namely, it has zero mean, constant variance σ^2 and is nonautocorrelated. If the coefficient of Y_{t-1} is in fact equal to 1, we face what is known as the *unit root problem*, i.e. a nonstationarity situation. Therefore, if we run the regression

$$Y_t = \rho Y_{t-1} + u_t \qquad \text{(III.8)}$$

and actually find that $\rho = 1$, then we say that the stochastic variable Y_t *has a unit root*, or in other words is a *random walk*.

The simplest and most widely used tests for unit roots were developed by Fuller (1976) and Dickey and Fuller (1979). These tests are generally referred to as *Dickey – Fuller*, or *(DF)*, *tests* and are based on regressions of the following form:

$$\Delta Y_t = \delta Y_{t-1} + u_t \qquad \text{(III.9)}$$

$$\Delta Y_t = \beta_0 + \delta Y_{t-1} + u_t \qquad \text{(III.10)}$$

$$\Delta Y_t = \beta_0 + \beta_1 t + \delta Y_{t-1} + u_t \qquad \text{(III.11)}$$

where $\delta = (\rho - 1)$, Δ is the first – difference operator and t is the time or trend variable. In each case the null hypothesis is $\delta = 0$, that is there is a unit root. The difference between (III.9) and the other two regressions lies in the inclusion of the constant (intercept) and the trend term.

If δ is in fact zero, we can write (III.9), and in a similar way (III.10) and (III.11), as

$$\Delta Y_t = (Y_t - Y_{t-1}) = u_t \qquad \text{(III.12)}$$

What (III.12) says is that the first differences of a random walk time series (= u_t) are a stationary time series because by assumption u_t is purely random.

Now if a time series is differenced once and the differenced series is stationary, we say that the original (random walk) series is *integrated of order 1*, denoted by *I(1)*. Similarly, if the original series has to be differenced twice (i.e. take first difference of the first difference) before it becomes stationary, the original series is *integrated of order 2*, or *I(2)*. In general, if a time series has to be differenced *d* times, it is *integrated of order d* or *I(d)*.

To find out if a time series is nonstationary, we can run the regression (III.8) and find out if ρ is statistically equal to 1 or, equivalently, estimate (III.9) and find out if $\delta = 0$ on the basis of the *t* statistic. Unfortunately the *t* value thus obtained does not follow Student's *t* distribution even in large samples. Under the null hypothesis that $\rho = 1$, the conventionaly computed *t* statistic is known as the τ *(tau) statistic*. In the literature the *tau test* is known as the *Dickey – Fuller (DF) test*, in honor of its discoverers. If the null hypothesis that $\rho = 1$ is rejected (i.e. the time series is stationary), the usual Student's *t* test can be used.

The unit root tests that we have discussed so far are valid only under the assumption that the error terms in the test regressions (III.9) to (III.11) are serially uncorrelated. This assumption is often untenable, because the regression functions for the test regressions do not depend on any economic variables. This makes it very likely that the error terms will display serial correlation. Therefore, we need unit root tests that are asymptotically valid in the presence of serial correlation. The simplest of such unit root tests are modified versions of the Dickey – Fuller τ tests. These are often called *augmented Dickey – Fuller tests*, or *ADF tests*.

So, if the error term u_t is autocorrelated, regression (III.11), and similarly (III.9) and (III.10), can be modified as following:

$$\Delta Y_t = \beta_0 + \beta_1 t + \delta Y_{t-1} + \alpha_i \sum_{i=1}^{m} \Delta Y_{t-i} + u_t \qquad \text{(III.13)}$$

where for example, $\Delta Y_{t-1} = Y_{t-1} - Y_{t-2}$, $\Delta Y_{t-2} = Y_{t-2} - Y_{t-3}$, etc. that is lagged difference terms are used. The number of lagged difference terms to include is often determined empirically, the idea being to include enough terms so that the error term in (III.13) is serially independent. The null hypothesis is still that $\delta = 0$ or $\rho = 1$, that is, a unit root

exists in Y. The ADF test statistic has the same asymptotic distribution as the DF statistic, so the same critical values can be used.

- **Theoretical Issues on Cointegration Analysis**

Economic theory often suggests that certain pairs of economic variables should be linked by a long – run equilibrium relationship. From a theoretical point of view, the power of economic equilibrium as an attractor should force different variables to move together in the long run even if not in the short run and even if they are individually non-stationary. In other words, although the variables may drift away from equilibrium for a while, economic forces may be expected to act so as to restore equilibrium. An example of such variables might include interest rates on assets of different maturities. However, there is no reason to restrict attention to pairs of variables, although it is easiest to do so. The analysis can also be extended to include as well groups of three, four, or more variables that can be expected to be linked by some long – run equilibrium relationship.

An equilibrium relationship is expressed through a function f $(y_1, y_2, \ldots y_n) = 0$, which describes the relationships that hold among the n variables x_1 to x_n, when the system is in equilibrium. The phrase "long – run equilibrium" is used in this case to denote the equilibrium relationship to which a system converges over time.

The concept of cointegration is fundamental to the understanding of long – run relationships among economic time series. If two or more variables are cointegrated, they must obey an equilibrium relationship in the long – run, although they may diverge substantially from equilibrium in the short-run. Expressed differently, cointegration entails a systematic co-movement among economic variables, which an economic system exemplifies precisely in the long run.

Suppose, to keep matters simple, that we are concerned with just two variables, y_{t1} and y_{t2}, each of which is known to be $I(1)$. Variables which are $I(1)$ tend to diverge as $n \rightarrow \infty$, because their unconditional variances are proportional to n. Thus it might seem that such variables could never be expected to obey any sort of long - run equilibrium relationship. But in fact it is possible for two or more variables to be $I(1)$ and yet for certain linear combinations of those variables to be $I(0)$. If that is the case, the variables are said to be *cointegrated*. Returning to the case of our two

variables, in the simplest case, y_{t1} and y_{t2} would be cointegrated if there exists a vector $\boldsymbol{\eta} \equiv [1 \; -\eta_2]^T$ such that, when the two variables are in equilibrium,

$$[\boldsymbol{y}_1 \;\; \boldsymbol{y}_2]\, \boldsymbol{\eta} \equiv \boldsymbol{y}_1 - \eta_2 \boldsymbol{y}_2 = 0 \qquad \text{(III.14)}$$

Here $\boldsymbol{y}_1$ and $\boldsymbol{y}_2$ denote n - vectors with typical elements y_{t1} and y_{t2}, respectively. The 2-vector $\boldsymbol{\eta}$ is called a *cointegrating vector*. It is clearly not unique, since it could be multiplied by any nonzero scalar without affecting the equality in (III.14).

Realistically, one might well expect y_{t1} and y_{t2} to be changing systematically as well as stochastically over time. Thus one might expect (III.14) to contain a constant term and perhaps one or more trend terms as well. If we write $\boldsymbol{Y} = [\boldsymbol{y}_1 \;\; \boldsymbol{y}_2]$, (III.14) can be rewritten to allow for this possibility as

$$\boldsymbol{Y}\boldsymbol{\eta} = \boldsymbol{X}\boldsymbol{\beta} \qquad \text{(III.15)}$$

where $\boldsymbol{X}$ denotes a non-stochastic matrix that may or not have any elements. If it does, the first column will be a constant, the second, if it exists, will be a linear time trend, the third, if it exists, will be a quadratic time trend, and so on. Since $\boldsymbol{Y}$ could contain more than two variables, (III.15) is actually a very general way of writing a cointegrating relationship among any number of variables.

At any particular time t, of course, an equality like (III.14) or (III.15) cannot be expected to hold exactly. We may therefore define the *equilibrium error* v_t as

$$v_t = \boldsymbol{Y}_t \boldsymbol{\eta} - \boldsymbol{X}_t \boldsymbol{\beta} \qquad \text{(III.16)}$$

where $\boldsymbol{Y}_t$ and $\boldsymbol{X}_t$ denote the tth rows of $\boldsymbol{Y}$ and $\boldsymbol{X}$, respectively. In the special case of (III.14), this equilibrium error would simply be $y_{t1} - \eta_2 y_{t2}$. The m variables y_{t1} through y_{tm} are said to be cointegrated if there exists a vector $\boldsymbol{\eta}$ such that v_t in (III.16) is $I(0)$.

The concept of cointegration brings with it two obvious econometric questions. The first is how to estimate the cointegrating vector $\boldsymbol{\eta}$, and the second is how to test whether two or more variables are in fact cointegrated. These questions are of course closely related; the answer to the second depends on the answer to the first. We begin to discuss the first one now.

The easiest way to estimate the cointegrating vector is to rewrite equation (III.16) as a regression and then to use OLS. This approach is associated with Engle and Granger (1987). Thus, if the coefficient on y_1 were arbitrarily normalized to unity, we could run the regression

$$y_1 = X\beta + Y \times \eta^* + v \qquad \text{(III.17)}$$

where Y^* is an n^* $(m-1)$ matrix with columns y_1 through y_m, and the parameter vector η^* is equal to minus the $m-1$ free elements of the parameter vector η that appears in (III.16).

There are apparently two serious problems with running a regression like (III.17). The first problem is that if the y_{it}'s are cointegrated, they are surely determined jointly, which implies that the error term will almost certainly not be independent of the regressors. Considering the following bivariate model:

$$\begin{aligned} \lambda_1 y_{t1} - y_{t2} &= u_{t1}, \qquad & (1-\rho_1 L)u_{t1} &= \varepsilon_{t1} \\ y_{t1} - \lambda_2 y_{t2} &= u_{t2}, \qquad & (1-\rho_2 L)u_{t2} &= \varepsilon_{t2} \end{aligned} \qquad \text{(III.18)}$$

where λ_1 and λ_2 are parameters, in the case of $\rho_1 = 1$ and $\rho_2 < 1$, for example, the relationship between y_{t1} and y_{t2} is

$$y_{t1} = \lambda_2 y_{t2} + \rho_2 (y_{t-1,1} - \lambda_2 y_{t-1,}) + \varepsilon_{t2} \qquad \text{(III.19)}$$

Thus, if we regress y_{t1} and y_{t2}, the error term is implicitly

$$\rho_2 (y_{t-1,1} - \lambda_2 y_{t-1,2}) + \varepsilon_{t2} \qquad \text{(III.20)}$$

and both terms here are correlated with y_{t2}. The second problem is that, in a regression like (III.17) we are regressing a variable which is $I(1)$ on one or more other $I(1)$ variables. This seems like a most undesirable thing to do, since it is a situation in which spurious regressions are very likely to arise.

Despite these apparent problems, when y_{t1} through y_{tm} are in fact cointegrated, the OLS estimates from regression (III.17) will be consistent. The first apparent problem does not matter asymptotically because y_{t2} is $I(1)$ and the two components of the error term in (III.20) are $I(0)$ (the first component is $I(0)$ only if y_{t1} and y_{t2} are in

fact cointegrated). Thus terms that involve the error term will be asymptotically negligible relative to terms that involve y_{t2}. The second apparent problem does not arise asymptotically for a similar reason, namely, that the (true) cointegrating relationship among the y_{it}'s creates terms that dominate any terms which might ordinarily cause spurious regressions. Another consequence of this is that the R^2 from (III.17) will tend to unity as $n \rightarrow \infty$.

Testing for cointegration

The most popular tests for cointegration, which are very closely related to unit root tests, were suggested by Engle and Granger (1987). The basic idea is very simple. If y_{t1} through y_{tm} are in fact cointegrated, the true equilibrium error term v_t must be $I(0)$. If they are not cointegrated, however, v_t must be $I(1)$. Thus one can test the null hypothesis of noncointegration against the alternative of cointegration by performing a unit root test on v_t.

However, in almost all cases, v_t will not be observed, because at least some elements of $\boldsymbol{\eta}$ will be unknown. It is therefore necessary to estimate $\boldsymbol{\eta}$. This could in principle be done in several ways, but the simplest approach is to apply OLS to regression (III.17). This procedure yields a vector of residuals, or estimated equilibrium errors, $\boldsymbol{v}$. If the variables y_{t1} through y_{tm} are in fact not cointegrated, regression (III.17) is a spurious one, and the vector $\boldsymbol{v}$ should have a unit root. Conventional unit root test statistics may be calculated using this vector of residuals. For obvious reasons, such tests are often called *residual-based cointegration tests*. Because $\boldsymbol{v}$ depends on one or more estimated parameters, which under the null hypothesis are the parameters of a spurious regression, the asymptotic distributions of residual-based cointegration test statistics are not the same as those of ordinary unit root test statistics.

Residual-based cointegration tests may be adapted from any of the usual unit root tests discussed above, provided the right critical values are used. The simplest procedure, sometimes called the *Engle-Granger test*, or the *EG test*, involves first estimating the cointegrating regression (III.17) and then using an ordinary Dickey-Fuller τ test, based on the regression

$$\Delta v = (\alpha - 1)\, v_{t-1} + e_t \qquad \text{(III.21)}$$

Since serial correlation is very often a problem, it is more common to use an *augmented Engle-Granger test* or *AEG test*. Because the cointegrating regression includes the columns of X among the regressors, it is not necessary to include X in the test regression (III.21).

Although tests based on the residual vector v are by far the most popular ones, numerous other cointegration tests have been proposed. References include Stock and Watson (1988), Johansen (1988, 1991) and Johansen and Juselious (1990, 1992). Johansen's approach will be discussed below.

Model - building with cointegrated variables.

Many economic time series are integrated of order one. Because of the possibility of a spurious regression, regressing the levels of a series which is $I(1)$ on the levels of one or more other series which are also $I(1)$ is generally not a good thing to do. At worst, we may "discover" an entirely spurious relationship. At best, we may consistently estimate the elements of some cointegrating vector, but standard asymptotic theory will not apply to our estimates, and we may therefore be led to make incorrect inferences about the parameters we have estimated. The specification and estimation of models for $I(1)$ variables is a very demanding task. In this section we will discuss some simple special cases and some simple results. In the next section we will discuss the estimation of vector autoregressions involving cointegrated variables.

The classical approach to dealing with integrated variables, especially in the time – series literature, has been to difference them as many times as needed to make them stationary. This approach has the merit of simplicity. Once all series have been transformed to stationarity, dynamic regression models may be specified in the usual way, and standard asymptotic results apply. The problem with this approach is that differencing eliminates the opportunity to estimate any relationships between the *levels* of the dependent and independent variables. But cointegration implies that such relationships exist, and they are often of considerable economic interest. Thus simply using differenced data is often not an appropriate strategy.

A second approach is to estimate some sort of *error - correction model*, or *ECM*. A simple but widely applicable single – equation ECM can be written as

$$\Delta y_t = z_t \alpha + \beta (y_{t-1} - \lambda x_{t-1}) + \gamma \Delta x_t + u_t , \qquad u_t \sim \text{IID}(0, \sigma^2) \qquad \text{(III.22)}$$

The dependent variable here is y_t, and the principal independent variable is x_t. These two variables are assumed to be $I(1)$ and cointegrated, which implies that the error – correction term $\beta (y_{t-1} - \lambda x_{t-1})$ is $I(0)$. The row vector z_t includes a constant term and any other independent variables, all of which are assumed to be either nonstochastic or $I(0)$. If (III.22) does not allow sufficiently rich dynamics, it can easily be extended by including more lags of Δx_t and increasing the lag on the error – correcting term.

If λ were known, there would clearly be no problem estimating (III.22) by least squares. The regressand and all the regressors would be either nonstochastic or $I(0)$. Thus the estimates of α, β, and γ would be root – n consistent and asymptotically normal. But in most cases λ is not known. There are several ways to proceed. The simplest is the *Engle – Granger two – step method* proposed by Engle and Granger (1987). The first step is to regress y_t on x_t, including a constant term and possibly a trend if the latter appears in z_t. This will yield a super – consistent estimate of λ, say λ. The second step is to replace λ by λ in (III.22) and then estimate that equation using OLS. Because of the super-consistency of λ, Engle and Granger are able to show that the resulting estimates of the other parameters are asymptotically the same as they would be if λ were known.

The principal merit of the Engle – Granger two step procedure is simplicity. However, there is evidence that it often does not work well in finite samples. The problem is that λ often seems to be severely biased. This bias then causes the other parameters to be biased as well. The problem appears to be least severe when the R^2 of the cointegrating regression is close to 1, as it must be when the sample size is sufficiently large. Thus a relatively low value of the R^2 from the cointegrating regression should be taken as a warning that the two – step procedure may not work well.

The simplest alternative to the Engle – Granger two step procedure is to estimate a model like

$$\Delta y_t = z_t \alpha + \beta y_{t-1} + \delta x_{t-1} + \gamma \Delta x_t + u_t , \qquad \text{(III.23)}$$

in which the new parameter δ is implicitly equal to $-\beta\lambda$. This regression looks rather odd, since the regressand is $I(0)$ and two of the regressors are $I(1)$. One might

therefore expect that standard asymptotic distribution theory would not apply to some or all of the parameter estimates. The asymptotic distribution theory for this equation is indeed nonstandard, but the practical problems turn out to be much less severe than one might expect. The key results for regressions like (III.23) were proved by Sims, Stock and Watson. (1990). Estimating equation (III.22) directly by nonlinear least squares is equivalent to estimating equation (III.23) by OLS. The fits of both equations will be the same, as will be the estimates of parameters that appear in both.

The estimation techniques that have been discussed in this section are all single – equation ones, and they are not in general efficient. Although the Engle – Granger two-step procedure is always super – consistent for λ, it is not asymptotically efficient. Other authors, including Johansen (1988, 1991), have proposed various full – system estimation methods. The approach of Johansen will be discussed in the next section.

Vector Autoregressions and Cointegration

One of the most interesting approaches to the full – system estimation of models involving cointegrated variables was developed in Johansen (1988, 1991) and Johansen and Juselius (1990, 1992). It is based on the estimation of a *vector autoregression*, or *VAR*, by maximum likelihood. We briefly discuss this approach.

Consider the following VAR in the levels of a set of variables:

$$\boldsymbol{Y}_t = \boldsymbol{Y}_{t-1}\boldsymbol{\Pi}_1 + \ldots + \boldsymbol{Y}_{t-p}\boldsymbol{\Pi}_p + \boldsymbol{U}_t \qquad \text{(III.24)}$$

where $\boldsymbol{Y}_t$ and $\boldsymbol{U}_t$ are $1 \times m$ row vectors, and $\boldsymbol{\Pi}_1$ through $\boldsymbol{\Pi}_p$ are $m \times m$ matrices of coefficients. For simplicity, there are no constant terms, although this assumption is rarely realistic. The VAR (11) can be reparametrized as

$$\Delta\boldsymbol{Y}_t = \Delta\boldsymbol{Y}_{t-1}\boldsymbol{\Gamma}_1 + \ldots + \Delta\boldsymbol{Y}_{t-p+1}\boldsymbol{\Gamma}_{p-1} - \boldsymbol{Y}_{t-p}\boldsymbol{\Pi} + \boldsymbol{U}_t \qquad \text{(III.25)}$$

where $\boldsymbol{\Gamma}_1 = \boldsymbol{\Pi}_1 - \mathbf{I}$, $\boldsymbol{\Gamma}_2 = \boldsymbol{\Pi}_2 + \boldsymbol{\Gamma}_1$, $\boldsymbol{\Gamma}_3 = \boldsymbol{\Pi}_3 + \boldsymbol{\Gamma}_2$, and so on. Thus the matrix $\boldsymbol{\Pi}$ is related to the $\boldsymbol{\Pi}_i$'s of (11) by

$$\boldsymbol{\Pi} = \mathbf{I} - \boldsymbol{\Pi}_1 - \ldots - \boldsymbol{\Pi}_P .$$

By stacking the n observations in (III.25), we can write the full system as

$$\Delta \boldsymbol{Y} = \Delta \boldsymbol{Y}_{-1} \boldsymbol{\Gamma}_1 + \ldots + \Delta \boldsymbol{Y}_{-(p-1)} \boldsymbol{\Gamma}_{p-1} - \boldsymbol{Y}_{-p} \boldsymbol{\Pi} + \boldsymbol{U} \qquad \text{(III.26)}$$

in obvious notation. Each term in (III.26) is an $n \times m$ matrix.

If all the variables in $\boldsymbol{Y}$ are nonstationary, (III.25) implies that there must be cointegration, and any vector in the range of $\boldsymbol{\Pi}$ must be a cointegrating vector. Suppose that $\boldsymbol{\Pi}$ has a rank of r, with $0 < r < m$. If that is the case, we can write $\boldsymbol{\Pi}$ as

$$\boldsymbol{\Pi} = - \boldsymbol{\eta}\boldsymbol{\alpha}^{\mathrm{T}} \qquad \text{(III.27)}$$

where $\boldsymbol{\alpha}$ and $\boldsymbol{\eta}$ are $m \times r$ matrices, and the minus sign is introduced here for later convenience. From (III.27), we see, that $\boldsymbol{Y}_{-p} \boldsymbol{\Pi} = - \boldsymbol{Y}_{-p} \boldsymbol{\eta}\boldsymbol{\alpha}^{\mathrm{T}}$. The cointegrating vectors are proportional to the columns of the matrix $\boldsymbol{\eta}$. Thus, for each column, say $\boldsymbol{\eta}_i$, $\boldsymbol{Y}\boldsymbol{\eta}_i$ is a stationary random variable. When $r = 1$, there is a single cointegrating vector, which is proportional to $\boldsymbol{\eta}_i$. When $r = 2$, there is a two – dimensional space of cointegrating vectors, spanned by $\boldsymbol{\eta}_1$ and $\boldsymbol{\eta}_2$, and so on. The two extreme cases are those in which $r = 0$, when there are no cointegrating vectors at all, and $r = m$, when any linear combination of the y_i's will be stationary, because each y_i will be $I(0)$.

The approach of Johansen (1988, 1991) is to estimate the VAR (12) subject to the constraint (III.27) for various values of r, using maximum likelihood. This estimation is based on the assumption that the error vector $\boldsymbol{U}_t$ is multivariate normal for each t and independent across observations. This assumption is not as restrictive as it may seem, since if there are enough lagged differences of $\boldsymbol{Y}$ included in (III.25), they should remove any evidence of serial correlation in the residuals. According to Johansen 's method, it is possible to maximize the loglikelihood function analytically conditional on any value of r.

Often, we are not specially interested in the parameters of the VAR (13). The focus of our interest is more likely to be testing the hypothesis of noncointegration against an alternative of cointegration of some chosen order. Should the null hypothesis that $r = 0$ be rejected, we may then wish to test the hypothesis that $r = 1$ against the alternative that $r = 2$, and so forth. The eigenvalues λ_i, $i = 1, \ldots, m$, provide a very convenient way to do this, in terms of a likelihood ratio test. Likelihood ratios for different values of r are therefore just products of the eigenvalues, raised to the

power $n/2$. If we take *logs* and multiply by 2 in order to obtain an *LR* statistic, we obtain $-n$ times a product of *logs* of the eigenvalues.

Specifically, to test the null that $r = r_1$, $0 \leq r_1 \leq m$, against the alternative that $r = r_2$, $r_1 < r_2 \leq m$, the *LR* test statistic is

$$LR = -n \sum_{i=r_1+1}^{r_2} \log \lambda_i \qquad \text{(III.28)}$$

Under, the various nulls that can be tested, the *LR* statistics (III.28) will have nonstandard asymptotic distributions that depend on the number of degrees of freedom $r_2 - r_1$ and on whether the VAR includes a constant or a trend term. Conditional on a given value of r, inference about the elements of cointegrating vectors can also be performed by means of *LR* statistics, which will then have their standard chi-squared asymptotic distributions under the null hypothesis being tested. This is a convenient property of the VAR approach.

Section B. Empirical Analysis

- **Data**

At this part of the paper we will estimate empirically equations (III.1) and (III.2) from section A of the econometric part. These equations will be estimated for the case of Greece, while the period under study is March 1994 to November 2000. Monthly data are used as they are considered to be more appropriate for this analysis. Daily data were not chosen because they are not available for a quite long period, and a longer period is certainly necessary when examining possible long-run relationships. As it was mentioned in the previous section, our three variables constitute of the daily inter-bank interest rate, the 3-month Athibor inter-bank interest rate and the one-year rate of government bonds. The source of our data was the "Economic Bulletin" published by the Central Bank of Greece. Finally, all the results were generated using the Microfit 4.0 for Windows econometric program.

- **Graphical presentation of time series data**

We explicitly explore whether there existed long-run co-movements between the one-day and the 3-month Athibor inter-bank rates, and between the last one and the one-year government bond rate by observing the graphs of their time series. In the following section the same object is tested by means of the cointegration techniques.

The raw data of the three interest rates and the relationships between them are illustrated in figures 1 and 2. To begin with, we first study the relationship between the daily and the 3-month Athibor inter-bank interest rates. From their plot in figure 1 it can be easily observed that these two rates clearly move together. Their relationship seems to be particularly closer from November 1994 to September 1997 and from January 1999 to November 2000.

Similar results are also obtained when exploring the linkage of the 3-month Athibor rate with the one-year government bond rate. The two interest rates appear to be approximately coincident. The coherence between them is very high except for the periods March 1994 to September 1994 and September 1997 to March 1998. From figure 2, which illustrates the time series of the two interest rates, we can observe their high co-movement during the whole sample.

In summary, the graphical presentation of a short-, a medium- and a long-term interest rate seems to support the argument that there are close long-run relationships between the daily and 3-month inter-bank interest rates, and the 3-month Athibor and one-year government bond rates. Although the series are trended and generally move with one another, from just depicting and observing their graphs it is not possible to say whether the above variables are cointegrated. A further examination by means of co-integration techniques is needed.

- **Results of Integration Analysis**

Before proceeding to the cointegration techniques, it is necessary to determine the order of integration of our variables. For this purpose, we implement the Dickey-Fuller (DF) and augmented Dickey-Fuller (ADF) tests for each one of the three interest rates. By this method, all our variables are tested for a unit root and the test is

repeated in first differences only for the variables that exhibit a unit root at their levels.

We begin the analysis with the first of our variables, the one-day inter-bank interest rate. Both the DF and ADF statistics are presented in Table 1. The Akaike Information and Schwarz Bayesian criteria indicate that the Augmented Dickey-Fuller tests of order 1 or 2 should be used in order to conclude about stationarity of our variable. Both the values of ADF(1) and ADF(2) tests are lower from the critical value at 5% significance level, hence supporting the H_0 hypothesis for a unit root. The next table (Table 2) presents the results of both DF and ADF tests for a second unit root, i.e., for a unit root on the first-differenced series of the one-day inter-bank rate. As can be easily seen, the results are in favor of stationarity on the first differences, and hence, the series of R_1 is best characterized as a $I(1)$ procedure, i.e., having a stochastic trend.

We now proceed to the next of our variables, the 3-month inter-bank interest rate. As can be observed in Table 3, the ADF(2) and ADF(3) test statistics, chosen by the Akaike Information and Schwarz Bayesian criteria provide significant evidence in favor of the unit root hypothesis. In consequence, we test for a unit root on the first difference of the R_2 interest rate. The results of both the DF and ADF tests, presented in Table 4, at all levels reject the hypothesis of a unit root. Hence, the R_2 time series is also characterized as a $I(1)$ procedure.

Finally, we apply the DF and ADF tests for the third variable, the one-year government bond rate. It can be deduced from the results of Table 5 that in no case is there significant evidence against the unit root hypothesis. When the tests are applied for the case of the first difference of the R_3 rate, the results (presented in Table 6) are in favor of stationarity. Under this evidence it is reasonable to conclude that the R_3 interest rate is a $I(1)$ process.

- **Results of Cointegration Analysis**

The above analysis confirmed the existence of a stochastic trend in each of the series, meaning that the observed time-series possess trends, which can be removed by differencing once. The question now is whether the long – run movements of the interest rates are determined by some common driving fundamentals, or in other

words, whether there exists cointegration between the one-day and the 3-month inter-bank rates, and the 3-month inter-bank and the one-year government bond rates.

We test for cointegration, and estimate any cointegrating vector, using the Johansen's method of maximum likelihood estimation of cointegrated systems. This procedure provides a unified framework for estimation and testing of cointegrating relations in the context of vector autoregressive (VAR) error correction models.

We first apply this method to the case of the one-day and the 3-month Athibor inter-bank interest rates. Since both interest rates are two jointly dependent, stochastic variables with $I(1)$ processes, we have considered a VAR model in order to capture any long-run relationship between them. However, before using the Johansen method, it is important to select the order of the VAR model, that is the number of lags in the involved VAR system. For this purpose we use both the Akaike Information Criterion (AIC) and the Schwarz Bayesian Criterion. In Table 7 there are presented the results of these two tests. As we can observe, the Akaike Criterion suggests a VAR of order 5, while the Schwarz Bayesian Criterion of order 2. Since we have a short time-series (75 observations), we cannot take the risk of over-parameterization and therefore choose 2 as the order of the VAR.

The choice of intercepts and trends included in the model is very important in testing for cointegration. In the present application, while an intercept is considered necessary, a trend is not included in the model. The reason is that although the underlying variables are trended, they move together, and it seems unlikely that there will be a trend in the cointegrating relation between the R_1 and R_2 interest rates.

Now, we can test for the null hypothesis of no cointegration against the alternative of cointegration, using Johansen's likelihood ratio test. The relevant statistics on this test are the maximum eigenvalue test statistic and the trace statistic. Table 8 reports the results of the cointegration tests.

According to these results, both the maximum eigenvalue test and trace test statistics strongly reject the null hypothesis that there is no cointegration between the R_1 and R_2 interest rates (namely that $r=0$) both at the five and ten per cent significance level. In consequence, we observe that the null hypothesis of one cointegrating vector is also rejected by both tests in favor of the alternative that there are two such vectors (again at five and ten per cent significance level). A similar result also follows from the values of the various model selection criteria reported in the third part of Table 8.

This complete agreement between the three procedures for testing the number of cointegrating relations is very rare. However, in practice, these three methods often result in conflicting conclusions, and the decision concerning the choice of r, the number of cointegrating relations, must be made in view of other information, perhaps from economic theory. Therefore, it can be strongly supported that a cointegration relationship exists between the one-day and the 3-month Athibor inter-bank interest rates for the period May 1994 to November 2000. The cointegrating vectors are provided in Table 9, and are given by the following equations:

$$0.057R_2(t) = -0.263 + 0.082\,R_1(t) + u_t \qquad \text{(III.29)}$$

and

$$0.014R_2(t) = -0.32 + 0.393R_1(t) + u_t \qquad \text{(III.30)}$$

Normalizing on the 3-month Athibor inter-bank rate (R_2), we obtain the following vectors:

$$R_2(t) = -4.57 + 1.42\,R_1(t) + u_t \qquad \text{(III.31)}$$

and

$$R_2(t) = -22.38 + 2.75\,R_1(t) + u_t \qquad \text{(III.32)}$$

Similar results are also obtained when performing the same tests in the case of the one-year government bond rate and the 3-month Athibor inter-bank interest rate. Akaike and Schwarz Bayesian criteria select an order of 1 for the VAR model (Table 11). Furthermore, both the maximum eigenvalue test and trace test statistics reject the null hypothesis of no cointegration between R_2 and R_3, but do not reject the hypothesis that there is one cointegrating relation between these variables (i.e. $r = 1$), at both five and ten percent significance level (Table 12). This cointegrating relation, reported in Table 13a, is given by the following equation:

$$0.0168R_3(t) = 0.0168 + 0.0112R_2(t) + u_t \qquad \text{(III.33)}$$

After normalizing on the one-year government bond rate (R_3) we obtain:

$$R_3(t) = 0.999 + 0.663R_2(t) + u_t \qquad \text{(III.34)}$$

Now that the existence of a long-run equilibrium relationship between the above interest rates has been established, it is appropriate to examine the associated Error Correction Mechanisms. We first consider the relationship between the daily inter-bank interest rate (R_1) and the 3-month Athibor rate (R_2). The usefullness of the Error Correction Mechanism comes from the fact that most of the times the relationship between the two rates is out of equilibrium, and it usually takes times to be corrected, that is to come back to equilibrium. The ECM term represents exactly this deviation and hence, it can be used as the "equilibrium error". The ECM for this case has the following form:

$$DR_2(t) = a_0 + a_1 DR_1(t) + a_2 DR_2(t\text{-}1) + a_3 DR_1(t\text{-}1) + a_4 ecm_{t\text{-}1} \qquad \text{(III.35)}$$

Regression (III.35) relates the change in R_2 to the change in R_1, the past values of their differences, and the "equilibrating error" in the previous period. In this regression DR_2 captures the short-run disturbances in R_2 whereas the error correction term $ecm_{t\text{-}1}$ captures the adjustment toward the long-run equilibrium. The value of this error correction term reflects the speed with which the 3-month Athibor inter-bank interest rate approaches its long-run equilibrium.

These error correction mechanisms are derived from the residuals of the long-run relationships (III.29) or (III.30) of the Johansen's procedure. They are given by the following equations:

$$ecm_{1t} = -\,0.057R_{2t} + 0.819R_{1t} - 0.263 \qquad \text{(III.36)} \quad \text{and}$$

$$ecm_{2t} = 0.0143R_{2t} - 0.039R_{1t} + 0.32 \qquad \text{(III.37)}$$

The sign of this term is expected to be positive and statistically significant.

The estimation procedure for equation (III.35) yielded:

$$DR_2(t) = -\,1.105\,DR_1(t\text{-}1) + 0.553\,DR_2(t\text{-}1) + 37.02\,ecm1(t\text{-}1) \qquad \text{(III38)}$$
$$(0.575) \qquad (0.375) \qquad (10.648)$$

and

$$DR_2(t) = -\,1.105\,DR_1(t\text{-}1) + 0.553\,DR_2(t\text{-}1) + 32.26\,ecm2(t\text{-}1) \qquad \text{(III.39)}$$
$$(0.575) \qquad (0.375) \qquad (10.648)$$

and R^2=0.383, DW=1.2915, $F(3, 75) = 15.5516$.

The results of the estimation are reported in Table 10. Examining this evidence, we observe that all the regressors (the past changes of the daily inter-bank interest rate, $DR_1(t-1)$, and the 3-month Athibor rate, $DR_2(t-1)$, as well as the two error correction mechanisms) are statistically significant terms in determining the behavior of R_2 interest rate. Therefore, it can be inferred that the change of the 3-month Athibor interest rate is determined by the short-run disturbances in its past value, the short-run changes induced in the one-day inter-bank interest rate, and by the adjustment process towards its long-run equilibrium. The fact that the coefficients of the error correction mechanisms have a quite high value leads to the conclusion that it would take only a short time for the equation to return to its equilibrium once it has been shocked. In particular, the values of 37.02 and 32.26 for the *ecm1* and *ecm2* respectively, imply that about 37.02 (and 32.26 in the second case), of the discrepancy between the actual and the long - run, or equilibrium, value of R_2 interest rate is eliminated or corrected each month.

We proceed the analysis for the second case, i.e. the relationship between the one-year government bond rate (R_3) and the 3-month Athibor inter-bank interest rate (R_2). The error correction mechanism is derived from the residuals of the long-run relationship (III.33) and it is given by the following equation:

$$ecm_{1t} = 0.0168R_{3t} - 0.0112R_{2t} - 0.0168 \qquad \text{(III.40)}$$

The estimation procedure yielded:

$$DR_3(t) = -\underset{(0.551)}{1.1376}\, ecm1(t-1) \qquad \text{(III.41)}$$

Table 14 presents the results of this estimation. We observe that the error correction mechanism is not only statistically significant, but also that it is the only variable, which determines the behavior of the R_3 interest rate. In other words, the one - year government bond rate is exclusively determined by its adjustment process towards its long-run equilibrium value. The short – run disturbances in either its past value or the 3-month Athibor inter-bank rate do not influence its level.

Conclusions

The main field of interest of this essay was whether changes induced to a short-term interest rate, which is used as an instrument in the conduct of monetary policy, are transferred to other interest rates with longer duration. The revealment of such linkages is very important in the transmission process, as it can be exploitable by monetary authorities in achieving their goals. In particular, our aim was to explore whether there exists a stable long-run relationship between (a) the 3-month Athibor interest rate and the one-day rate both of the inter-bank market and (b) the one-year government bond rate and the 3-month Athibor interest rate. For this purpose the means of cointegration techniques were used.

The application of the Johansen's method of maximum likelihood revealed that indeed a cointegration relationship exists between the 3-month Athibor and the one-day inter-bank interest rates, and the one-year government bond rate and the 3-month Athibor rate for the period May 1994 to November 2000. In consequence, the Error Correction Models were used in order to explore more precisely the dynamics that determine the path of the 3-month rate, as well of the one-year government bond rate. It was found that the behavior of the 3-month Athibor interest rate is determined not only by the short-run disturbances in its past value and the short-run changes induced in the one-day inter-bank rate, but also by the adjustment process towards its long-run equilibrium. In addition, in the equation of the one-year bond rate it was also found that any discrepancy between the actual and the long - run, equilibrium, value is corrected rapidly.

It is worth - while to mention the policy implications flowing from the results of the above analysis. The finding of a stable long-run relationship between the one-day and the 3-month inter-bank rates, and the latter and the one-year bond rate is important given the crucial role it plays in the transmission process. The daily inter-bank rate can be directly controlled by the Central Bank of Greece. The stable long-run relationships between the daily and 3-month inter-bank rates and the latter and the one-year bond rate mean that the authorities can effectively target the one-year government bond rate. This becomes most significant for those conducting monetary policy when we consider the important role that long-term rates play in affecting aggregate demand and, generally, magnitudes of the real sector of the economy.

APPENDIX

Figure 1

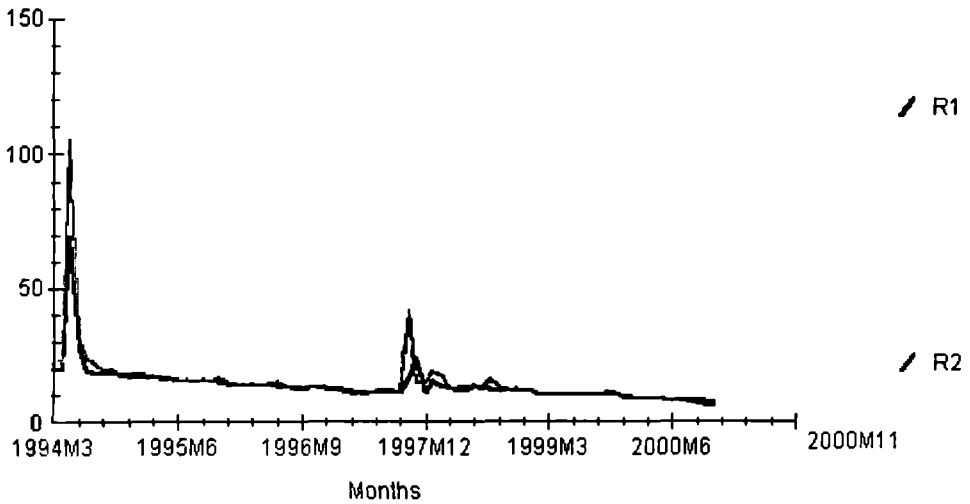
150
100
50
0
R1
R2
1994M3
1995M6
1996M9
1997M12
1999M3
2000M6
2000M11
Months

Figure 2

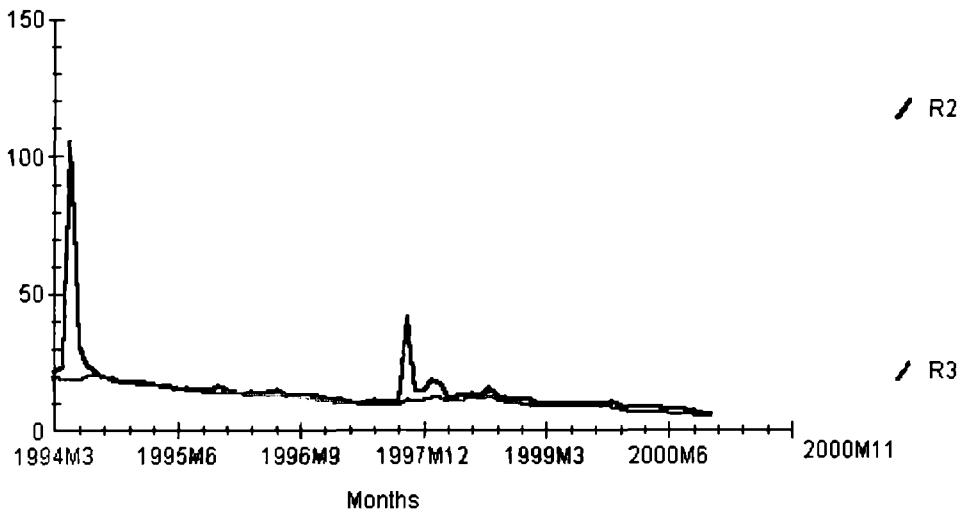
150
100
50
0
R2
R3
1994M3
1995M6
1996M9
1997M12
1999M3
2000M6
2000M11
Months

Table 1

Unit root tests for variable R_l

The Dickey-Fuller regressions include an intercept but not a trend

68 observations used in the estimation of all ADF regressions.

Sample period from 1995M4 to 2000M11

	Test Statistic	LL	AIC	SBC	HQC
DF	-3.3237	-139.0822	-141.0822	143.3017	-141.9617
ADF(1)	-2.2422	-135.7841	-138.7841	-142.1134	-140.1033
ADF(2)	-1.6782	-133.7063	-137.7063	-142.1453	-139.4651
ADF(3)	-1.4306	-133.0034	-138.0034	-143.5522	-140.2020
ADF(4)	-1.3337	-132.8930	-138.8930	-145.5515	-141.5313
ADF(5)	-1.3567	-132.8312	-139.8312	-147.5995	-142.9093

95% critical value for the augmented Dickey-Fuller statistic = -2.9042

LL = Maximized log-likelihood AIC = Akaike Information Criterion

SBC = Schwarz Bayesian Criterion HQC = Hannan-Quinn Criterion

Table 2

Unit root tests for variable DR_1

The Dickey-Fuller regressions include an intercept but not a trend

67 observations used in the estimation of all ADF regressions.

Sample period from 1995M5 to 2000M11

	Test Statistic	LL	AIC	SBC	HQC
DF	-12.3584	-136.7769	-138.7769	-140.9816	-139.6493
ADF(1)	-9.1034	-133.6780	-136.6780	-139.9851	-137.9866
ADF(2)	-7.1298	-132.6116	-136.6116	-141.0210	-138.3564
ADF(3)	-5.7217	-132.3769	-137.3769	-142.8886	-139.5579
ADF(4)	-4.5074	-132.3648	-138.3648	-144.9789	-140.9821
ADF(5)	-4.7421	-131.0013	-138.0013	-145.7178	-141.0548

95% critical value for the augmented Dickey-Fuller statistic = -2.9048

LL = Maximized log-likelihood AIC = Akaike Information Criterion

SBC = Schwarz Bayesian Criterion HQC = Hannan-Quinn Criterion

Table 3

Unit root tests for variable R_2

The Dickey-Fuller regressions include an intercept but not a trend

68 observations used in the estimation of all ADF regressions.

Sample period from 1995M4 to 2000M11

	Test Statistic	LL	AIC	SBC	HQC
DF	-5.5717	-193.6615	-195.6615	-197.8810	-196.5409
ADF(1)	-3.4977	-191.5578	-194.5578	-197.8870	-195.8769
ADF(2)	-2.3886	-189.1612	-193.1612	-197.6002	-194.9201
ADF(3)	-1.8618	-188.0664	-193.0664	-198.6152	-195.2650
ADF(4)	-1.8321	-188.0506	-194.0506	-200.7091	-196.6889
ADF(5)	-1.7519	-188.0488	-195.0488	-202.8171	-198.1268

95% critical value for the augmented Dickey-Fuller statistic = -2.9042

LL = Maximized log-likelihood AIC = Akaike Information Criterion

SBC = Schwarz Bayesian Criterion HQC = Hannan-Quinn Criterion

Table 4

Unit root tests for variable DR_2

The Dickey-Fuller regressions include an intercept but not a trend

67 observations used in the estimation of all ADF regressions.

Sample period from 1995M5 to 2000M11

	Test Statistic	LL	AIC	SBC	HQC
DF	-13.7865	-195.0063	-197.0063	199.2110	197.8787
ADF(1)	-10.3259	-189.7220	-192.7220	196.0291	194.0306
ADF(2)	-8.0907	-187.5759	-191.5759	195.9853	193.3207
ADF(3)	-5.8241	-187.5325	-192.5325	198.0442	194.7135
ADF(4)	-4.9368	-187.4058	-193.4058	200.0199	96.0230
ADF(5)	-4.4931	-187.1795	-194.1795	201.8959	197.2329

95% critical value for the augmented Dickey-Fuller statistic = -2.9048

LL = Maximized log-likelihood

AIC = Akaike Information Criterion

SBC = Schwarz Bayesian Criterion

HQC = Hannan-Quinn Criterion

Table 5

Unit root tests for variable R_3

The Dickey-Fuller regressions include an intercept but not a trend

68 observations used in the estimation of all ADF regressions.

Sample period from 1995M4 to 2000M11

	Test Statistic	LL	AIC	SBC	HQC
DF	-1.0503	-53.7663	-55.7663	-57.9858	-56.6457
ADF(1)	-1.0313	-53.7528	-56.7528	-60.0820	-58.0719
ADF(2)	-1.0262	-53.7483	-57.7483	-62.1873	-59.5072
ADF(3)	-1.0332	-53.5714	-58.5714	-64.1202	-60.7700
ADF(4)	-1.0503	-52.7354	-58.7354	-65.3939	-61.3737
ADF(5)	-1.1777	-45.1787	-52.1787	-59.9470	-55.2567

95% critical value for the augmented Dickey-Fuller statistic = -2.9042

LL = Maximized log-likelihood AIC = Akaike Information Criterion

SBC = Schwarz Bayesian Criterion HQC = Hannan-Quinn Criterion

Table 6

Unit root tests for variable DR_3

The Dickey-Fuller regressions include an intercept but not a trend

67 observations used in the estimation of all ADF regressions.

Sample period from 1995M5 to 2000M11

	Test Statistic	LL	AIC	SBC	HQC
DF	-8.2891	-53.9902	-55.9902	-58.1949	-56.8626
ADF(1)	-5.7084	-53.9894	-56.9894	-60.2964	-58.2980
ADF(2)	-4.3011	**-53.8278**	**-57.8278**	-62.2372	-59.5726
ADF(3)	-3.2213	-53.0469	-58.0469	-63.5587	-60.2279
ADF(4)	-4.7586	-45.6286	-51.6286	-58.2426	-54.2458
ADF(5)	-3.1301	-43.0904	-50.0904	-57.8068	-53.1438

95% critical value for the augmented Dickey-Fuller statistic = -2.9048

LL = Maximized log-likelihood

AIC = Akaike Information Criterion

SBC = Schwarz Bayesian Criterion

HQC = Hannan-Quinn Criterion

Table 7

Test Statistics and Choice Criteria for Selecting the Order of the VAR Model

Based on 75 observations from 1994M9 to 2000M11. Order of VAR = 6

List of variables included in the unrestricted VAR:

R_2 R_1

List of deterministic and/or exogenous variables:

CNST

Order	LL	AIC	SBC	LR test	Adjusted LR test
6	-242.7210	-268.7210	-298.8483	------	------
5	-242.9620	-264.9620	-290.4543	CHSQ(4)=.48195 [.975]	.39841[.983]
4	-249.5818	-267.5818	-288.4392	CHSQ(8)=13.7216 [.089]	1.3432[.183]
3	-259.6284	-273.6284	-289.8508	CHSQ(12)=33.8148 [.001]	27.9535[.006]
2	-261.4022	-271.4022	-282.9897	CHSQ(16)=37.3625 [.002]	0.8863[.014]
1	-301.2586	-307.2586	-314.2111	CHSQ(20)=117.0753 [.000]	96.7822[.000]
0	-382.3161	-384.3161	-386.6336	CHSQ(24)=279.1902 [.000]	230.7973[.000]

AIC=Akaike Information Criterion SBC=Schwarz Bayesian Criterion

Table 8

Cointegration with restricted intercepts and no trends in the VAR

Cointegration LR Test Based on Maximal Eigenvalue of the Stochastic Matrix

79 observations from 1994M5 to 2000M11. Order of VAR = 2.

List of variables included in the cointegrating vector:

R_2 R_1 *Intercept*

List of eigenvalues in descending order:

.39301 .14014 .0000

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	39.4399	15.8700	13.8100
r<= 1	r = 2	11.9276	9.1600	7.5300

Use the above table to determine r (the number of cointegrating vectors).

Cointegration with restricted intercepts and no trends in the VAR

Cointegration LR Test Based on Trace of the Stochastic Matrix

79 observations from 1994M5 to 2000M11. Order of VAR = 2.

List of variables included in the cointegrating vector:

R_2 R_1 *Intercept*

List of eigenvalues in descending order:

.39301 .14014 .0000

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r>= 1	51.3675	20.1800	17.8800
r<= 1	r = 2	11.9276	9.1600	7.5300

Use the above table to determine r (the number of cointegrating vectors).

Table 8 (Continue)

Cointegration with restricted intercepts and no trends in the VAR

Choice of the Number of Cointegrating Relations Using Model Selection Criteria

79 observations from 1994M5 to 2000M11. Order of VAR = 2.

List of variables included in the cointegrating vector:

R_2 R_1 *Intercept*

List of eigenvalues in descending order:

.39301 .14014 .0000

Rank	Maximized LL	AIC	SBC	HQC
r = 0	-481.4947	-485.4947	-490.2336	-487.3933
r = 1	-461.7748	-469.7748	-479.2526	-473.5719
r = 2	-455.8110	-465.8110	-477.6582	-470.5573

AIC = Akaike Information Criterion SBC = Schwarz Bayesian Criterion

HQC = Hannan-Quinn Criterion

Table 9

Estimated Cointegrated Vectors in Johansen Estimation

(Normalized in Brackets)

Cointegration with restricted intercepts and no trends in the VAR

79 observations from 1994M5 to 2000M11. Order of VAR = 2, chosen r =2.

List of variables included in the cointegrating vector:

R2 R1 Intercept

	Vector 1	Vector 2
R_2	-.057717	.014302
	(-1.0000)	(-1.0000)
R_1	.081912	-.039396
	(1.4192)	(2.7546)
Intercept	-.26380	.32017
	(-4.5705)	(-22.3861)

Table 10

ECM for variable R_2 estimated by OLS based on cointegrating VAR(2)

Dependent variable is dR_2

79 observations used for estimation from 1994M5 to 2000M11

Regressor	Coefficient	Standard Error	T-Ratio[Prob]
DR_{21}	.55434	.37598	1.4744[.145]
DR_{11}	-1.1055	.57533	-1.9215[.058]
$ecm_1(-1)$	37.0205	10.6480	3.4768[.001]
$ecm_2(-1)$	32.2629	10.6480	3.0299[.003]

List of additional temporary variables created:

$DR_2 = R_2 - R_2(-1)$

$DR_{21} = R_2(-1) - R_2(-2)$

$DR_{11} = R_1(-1) - R_1(-2)$

$ecm_1 = -.057717*R_2 + .081912*R_1 - .26380;$

$ecm_2 = .014302*R_2 - .039396*R_1 + .32017$

R-Squared	.38350	R-Bar-Squared	.35884
S.E. of Regression	10.6480	F-stat. F(3, 75)	15.5516[.000]
Mean of Dependent Variable	-.22025	S.D. of Dependent Variable	13.2980
Residual Sum of Squares	8503.5	Equation Log-likelihood	-296.9082
Akaike Info. Criterion	-300.9082	Schwarz Bayesian Criterion	-305.6471
DW-statistic	1.2915	System Log-likelihood	-455.8110

Table 10 (continue)

Test Statistics	Diagnostic Tests	
	LM Version	F Version
A:Serial Correlation	CHSQ(12) = 9.0507[.699]	F(12,63)=.67929[.765]
B:Functional Form	CHSQ(1) = 9.6620[.002]	F(1,74)=10.3116[.002]
C:Normality	CHSQ(2) = 10408.7[.000]	Not applicable
D:Heteroscedasticity	CHSQ(1) = .9739E-3[.975]	F(1,77)= .9493E-3[.976]

A:Lagrange multiplier test of residual serial correlation

B:Ramsey's RESET test using the square of the fitted values

C:Based on a test of skewness and kurtosis of residuals

D:Based on the regression of squared residuals on squared fitted values

Table 11

Test Statistics and Choice Criteria for Selecting the Order of the VAR Model

Based on 75 observations from 1994M9 to 2000M11. Order of VAR = 6

List of variables included in the unrestricted VAR:

R_3 R_2

List of deterministic and/or exogenous variables:

CNST

Order	LL	AIC	SBC	LR test	Adjusted LR test
6	-225.1439	-251.1439	-281.2712	------	------
5	-235.4025	-257.4025	-282.8949	CHSQ(4)=20.5173[.000]	16.9610[.002]
4	-239.1337	-257.1337	-277.9911	CHSQ(8)=27.9796[.000]	23.1298[.003]
3	-240.9290	-254.9290	-271.1514	CHSQ(12)=31.5702[.002]	26.0981[.010]
2	-243.8163	-253.8163	-265.4037	CHSQ(16)=37.3448[.002]	30.8717[.014]
1	-244.6480	-250.6480	-257.6005	CHSQ(20)=39.0083[.007]	32.2469[.041]
0	-397.3247	-399.3247	-401.6422	CHSQ(24)=344.3616[.000]	284.6723[.000]

AIC=Akaike Information Criterion SBC=Schwarz Bayesian Criterion

Table 12

Cointegration with restricted intercepts and no trends in the VAR

Cointegration LR Test Based on Maximal Eigenvalue of the Stochastic Matrix

80 observations from 1994M4 to 2000M11. Order of VAR = 1.

List of variables included in the cointegrating vector:

R_3 R_2 *Intercept*

List of eigenvalues in descending order:

.48677 .082095 .0000

Null	Alternative	Statistic	95% Critical Value	90%Critical Value
r = 0	r = 1	53.3624	15.8700	13.8100
r<= 1	r = 2	6.8529	9.1600	7.5300

Use the above table to determine r (the number of cointegrating vectors).

Cointegration with restricted intercepts and no trends in the VAR

Cointegration LR Test Based on Trace of the Stochastic Matrix

80 observations from 1994M4 to 2000M11. Order of VAR = 1.

List of variables included in the cointegrating vector:

R_3 R_2 *Intercept*

List of eigenvalues in descending order:

.48677 .082095 .0000

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r>= 1	60.2153	20.1800	17.8800
r<= 1	r = 2	6.8529	9.1600	7.5300

Use the above table to determine r (the number of cointegrating vectors).

Table 12 (continue)

Cointegration with restricted intercepts and no trends in the VAR

Choice of the Number of Cointegrating Relations Using Model Selection Criteria

80 observations from 1994M4 to 2000M11. Order of VAR = 1.

List of variables included in the cointegrating vector:

R_3 R_2 *Intercept*

List of eigenvalues in descending order:

.48677 .082095 .0000

Rank	Maximized LL	AIC	SBC	HQC
r = 0	-386.7662	-386.7662	-386.7662	-386.7662
r = 1	-360.0851	-364.0851	-368.8491	-365.9951
r = 2	-356.6586	-362.6586	-369.8047	-365.5237

AIC = Akaike Information Criterion SBC = Schwarz Bayesian Criterion

HQC = Hannan-Quinn Criterion

Table 13

Estimated Cointegrated Vectors in Johansen Estimation

(Normalized in Brackets)

Cointegration with restricted intercepts and no trends in the VAR

---

80 observations from 1994M4 to 2000M11. Order of VAR = 1, chosen r =1.

List of variables included in the cointegrating vector:

R_3 R_2 *Intercept*

	Vector 1
R_3	.016886
	(-1.0000)
R_2	-.011208
	(.66374)
Intercept	-.016884
	(.99985)

Table 14

ECM for variable R_3 estimated by OLS based on cointegrating VAR(1)

Dependent variable is DR_3

80 observations used for estimation from 1994M4 to 2000M11

Regressor	Coefficient	Standard Error	T-Ratio[Prob]
$ecm_1(-1)$	-1.1376	.55118	-2.0640[.042]

List of additional temporary variables created:

$DR_3 = R_3 - R_3(-1)$

$ecm_1 = .016886*R_3 - .011208*R_2 - .016884$

R-Squared	-.038230	R-Bar-Squared	-.038230
S.E. of Regression	.55118	F-stat.	NONE
Mean of Dependent Variable	-.16500	S.D. of Dependent Variable	.54094
Residual Sum of Squares	24.0006	Equation Log-likelihood	-65.3571
Akaike Info. Criterion	-66.3571	Schwarz Bayesian Criterion	-67.5481
DW-statistic	1.9712	System Log-likelihood	-360.0851

Diagnostic Tests

Test Statistics	LM Version	F Version
A:Serial Correlation	CHSQ(12) = 21.0872[.049]	F(12,67)= 1.9985[.038]
B:Functional Form	CHSQ(1) = 5.9362[.015]	F(1,78) = 6.2516[.015]
C:Normality	CHSQ(2) = 40.8598[.000]	Not applicable
D:Heteroscedasticity	CHSQ(1) = .010660[.918]	F(1,78) = .010395[.919]

A:Lagrange multiplier test of residual serial correlation

B:Ramsey's RESET test using the square of the fitted values

C:Based on a test of skewness and kurtosis of residuals

D:Based on the regression of squared residuals on squared fitted values

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