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WHY DOES MONEY MATTER? A STRUCTURAL ANALYSIS OF MONETARY POLICY, CREDIT AND AGGREGATE SUPPLY EFFECTS IN ITALY

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WHY DOES MONEY MATTER? A STRUCTURAL ANALYSIS OF MONETARY POLICY, CREDIT AND AGGREGATE SUPPLY EFFECTS IN ITALY*

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Abstract

The 1990s witnessed the resurgence of the view that "money matters" in the explanation of *real* macroeconomic fluctuations and in their stabilization. However, robust and persuasive explanations of the impact of monetary policy on economic activity are still matter of research. Drawing on the modern literature on the monetary transmission mechanisms and capital market imperfections, we put forward a theoretical and empirical analysis of the "credit-cost channel" of monetary policy. The thrust of the model is that firms' reliance on bank loans ("credit channel") may make *aggregate supply* sensitive to bank interest rates ("cost channel"), which are in turn driven by the official rate controlled by the central bank and a credit risk premium charged by banks on firms. Shocks in either variable, produce a pattern of macroeconomic relationships that fit and explain the observed empirical regularities in major industrial countries, with no recourse to additional non-competitive hypotheses. Moreover, it is shown that by controlling bank interest rates, monetary policy can exert *permanent* effects on the output and inflation equilibrium paths. By means of the Johansen-Juselius structural cointegration approach, we find statistical support for this view in Italian data for the pre-EMU years.

Keywords: Macroeconomics and monetary economics, Monetary transmission mechanisms, Structural cointegration models, Italian economy

JEL codes: E51, C32

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WHY DOES MONEY MATTER? A STRUCTURAL ANALYSIS OF MONETARY POLICY, CREDIT AND AGGREGATE SUPPLY EFFECTS IN ITALY

1. Introduction

The 1990s witnessed the resurgence of the view that "money matters" in the explanation of *real* macroeconomic fluctuations and in their stabilization. Most economists now share, with minor theoretical or empirical nuances, a core of common ideas that can be summarized as follows. 1) Monetary policy impulses have persistent real effects. 2) The typical observed pattern is one where policy interventions (mainly activated by changes in administered rates and money-market rates) are followed by quick and large responses in short-term interest rates, monetary aggregates, total credit, and different measures of real economic activity, and by slow and delayed adjustment of different price indexes. 3) Real wages and profits are also procyclical with output after a monetary shock[1]. However, robust and persuasive explanations of the impact of monetary policy on economic activity are still matter of research. In a sense, we have evidence in search of theory. The old neoclassical synthesis focused on the effects of monetary policy on aggregate demand through the sensitivity of households' and firms' expenditure to long-term interest rates and real money balances. None of these links ever appeared to be particularly strong and significant empirically as an exhaustive explanation. The prevailing view today is that money matters because different mechanisms co-operate in *amplifying* the impact of monetary policy on economic activity.

[1]Good overall presentations of this view can be found in Mishkin (1995), Christiano et al. (1996, 1997), Goodfriend-King (1997), AEA (1997), Clarida et al. (1999). Goodfriend and King (1997) have aptly called this view "the new neoclassical synthesis" since its basic tenets are apparently akin to the macroeconomic consensus of the 1960s. The qualification "new" refers to the deeper microeconomic foundations of two key factors in the transimission mechanism from monetary to real variables: i) wage and price rigidities, ii) the functioning of financial markets especially as regards firms' spending capacity.

A prominent line of investigation has recently pointed to the so-called "*credit channel*" of the transmission of monetary policy[2]. This channel helps explain the large impact that monetary interventions are observed to exert on economic activity by way of financial markets imperfections. The most important source of imperfection is seen in asymmetric information generating agency problems between the firm and its external financial suppliers. As a consequence, and in contrast with the Modigliani-Miller approach underlying the old neoclassical synthesis, firms are not indifferent among different financial resources but follow a "pecking order". They rely primarily on internal funds, whereas bank credit is the first choice among external sources, most likely for small firms with poor capacity of internal accumulation and with limited access to open markets. When embedded in imperfect financial markets, a monetary restriction that lowers asset prices, diverts bank funds from loans to bonds and raises bank interest rates worsens almost all possible sources of investments (Bernanke and Blinder (1988)). In spite of extensive research at the micro and macro-level in various industrialized countries, however, the identification and strength of an independent credit channel of monetary policy are still matters of discussion (e.g. Bernanke and Gertler (1995), Eichenbaum (1994) Angeloni et al. (2002)).

Another research track follows the theoretical argument that limiting the link between monetary policy and economic activity to aggregate-demand effects is an over-simplification of microeconomic relationships. There are, in fact, several possible links with *aggregate supply* too. First, investment decisions determine future production capacity; if imperfect financial markets in some way transmit monetary policy impulses through constrained investment decisions, the effects should also manifest themselves in current production decisions which must be consistent with the overall intertemporal production path of each firm (e.g. Stiglitz (1992))[3].

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[2]In truth, this is a rather heterogeneous collection of views, dating far back in history, sharing the idea that changes in *banks' assets* (i.e. total credit to the economy), rather than in banks' liabilities (i.e. money balances in the economy), are the key mechanism linking economic activity to monetary policy (Trautwein (2000)). In this paper we cannot offer a thorough survey of the modern credit view. The interested reader is referred to Greenwald and Stiglitz (1993b), Gertler and Gilchrist (1993) Bernanke and Gertler (1995).

[3]A formal model can also be found in Myatt and Scarth (1995) who relate current labour demand to investment through equipment production and installment.

Second, besides fixed capital, also working capital may need financial resources as current inputs should be paid before output can be sold, and these resources (liquidity, inventories, credit, etc.) carry a financial cost. Consequently, the interest rate paid on working capital affects production costs – a view largely shared by businessmen (e.g. Goodhart (1986)) – and monetary policy, by altering interest rates, can influence aggregate supply. Barth and Ramey (2001) testify to the growing interest in this further "*cost channel*" of monetary policy and provide evidence of its importance for monetary transmission[4].

The supply-side effects of monetary policy have several interesting implications. To begin with, as stressed by Greenwald and Stiglitz (1987) co-movements of demand and supply after a monetary shock can provide a straightforward explanation for the observed pattern of large adjustments in quantities and small ones in prices even in competitive markets. Moreover, the empirical regularity that real wages and profits are procyclical with output after a monetary shock is at variance with the traditional view of the demand effects of monetary policy (real wages should be countercyclical or remain unaffected depending on the degree of price/wage stickiness), whereas it indicates that supply effects do play a significant explanatory role (Christiano et al. (1997)). If, say, a monetary restriction raises firms' variable costs and/or forces them to cut production, then, for a given monetary wage, prices may well increase and real wages fall (Blinder (1987), Barth and Ramey (2001)). Alternatively, firms may respond by cutting back labour demand, thus forcing real wages to fall directly (Greenwald and Stiglitz (1988, 1993a), Christiano and Eichenbaum (1992), Christiano et al. (1997)). Finally, these effects also call into question the general presumption that real effects of monetary shocks can only arise

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[4]Like the credit channel, also the cost channel of monetary policy has antecedents. As is well known, Keynes prior to the *General Theory* envisaged a theory of production based on firms' access to liquid means of payment of inputs (1933), and later (1937) he added firms' "finance motive" to the aggregate demand for money. Subsequent developments have led to "money-in-the-production-function" models (e.g. Simos (1981)), either by inserting money balances in the production function as a complementary factor or by adding a separable monetary counterpart of (or constraint on) demand for specific inputs (e.g. Vickers (1981), Mitchell (1984), Ramey (1989)). This latter approach is also common to production theories where firms' demand for money arises from the time mismatch between purchases of inputs and sales of output (e.g. Hicks (1973), Farmer (1984), Amendola and Gaffard (1998)).

as short-lived consequences of sticky nominal variables, whereas the empirical persistence of these real effects arises from underlying shifts in the long-run paths of output and employment (Greenwald and Stiglitz (1993b)). The aim of this paper is to contribute to this line of research by proposing a theoretical model and then by providing econometric evidence drawn from the Italian economy.

The paper is organized as follows. Section 2 presents our theoretical model and a comparative analysis with other models. Various macroeconomic models have recently been produced which link monetary policy to aggregate supply via firms' working capital (see Barth and Ramey (2001) for an overview). Our own approach is characterized by the following features. First of all, we believe that important insights can be gained by blending the cost channel of monetary transmission with the credit channel. In fact, for working capital finance to be consistently treated as an extra cost firms should be forced to resort to external sources (otherwise, in a perfect capital market firms would obtain all the liquidity they need from owners at the capital rental rate (Greenwald and Stiglitz (1993a), Holmstroem and Tirole (1998)). The capital market imperfections adopted in the credit-channel literature provide a consistent framework for this role of credit in the production process[5]. Bank credit is a close substitute for internal liquidity as well as for direct lending and as such it is largely used by firms to finance working capital (Blinder and Stiglitz (1983)). Thus the credit channel may act as amplifier of the supply-side impact of monetary shocks. These may trigger cost effects via bank interest rates (e.g. Greenwald and Stiglitz (1988, 1993a)) as well as rationing effects on credit availability (e.g. Blinder (1987), Stiglitz and Weiss (1992))[6]. Both effects were indeed detected in early empirical microeconomic investigations of the credit channel (Kashyap et al. (1993, 1994), Gertler and Gilchrist (1993, 1994)).

In the second place, we do not wish to add non-competitive conditions in labour and goods markets (staggered contracts, sticky prices, monopolistic competition, etc.). Though these are important features of real

[5] For instance Christiano et al. (1997) have both intermediaries and direct lending from households to firms in their model, but this co-existence is not motivated theoretically.

[6] Other relevant macro-models with credit-dependent firms are those by Bernanke and Gertler (1990), Hahn and Solow (1995), Kiyotaki and Moore (1997).

economies present in comparable models (e.g. Barth and Ramey (2001), Christiano et al. (1997), Ravenna and Walsh (2003)), we believe that an important theoretical issue is establishing that for the supply-side effects of monetary policy to be significant imperfect capital markets are a sufficient prime principle, as suggested by Greenwald and Stiglitz (1993b).

We therefore present a competitive macroeconomic model with bank-dependent firms where monetary policy affects economic activity through changes in the bank interest rate with a compound "credit-cost channel" (CCC)[7]. The economy consists of three competitive markets – labour, credit, and output – and three classes of agents – households, firms and banks – with a central bank. Following the relevant literature, the key feature of the model is that firms need external funds in advance to finance working capital. This consists of labour only, but bank loans are the sole financial resource to which firms have access due to an ex-post verification problem by individual lenders. Banks raise funds from households and the central bank. When firms plan production and demand bank loans, they are uncertain about their future revenue from output sales. A positive probability exists that a firm is unable to repay its initial debt. Banks anticipate this risk when granting loans, and in consideration of their commitment towards depositors they can insure themselves by borrowing reserves from the central bank at a given official interest rate. As a result, credit risk and the official rate determine the bank interest rate charged to firms, providing the link between monetary policy and aggregate supply[8].

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[7]Our model draws on Greenwald and Stiglitz (1988, 1993a) and elaborates on versions provided by Fiorentini and Tamborini (2001, 2002). The model distinguishes from, and integrates, Greenwald and Stiglitz (1988), where non-competitive wage setting was assumed and monetary policy was simply the exogenous control over the quantity of money, and (1993a), where aggregate demand was not modelled, labour market relations were postulated without explicit demand-supply analysis, and the central bank was missing altogether. With respect to Fiorentini and Tamborini (2002), where specific functional forms of the production functions and of the households' utility function were assumed, we propose here a generalization of the model and a different estimation procedure.

[8]In our view this is a simpler and more natural link than the one proposed by Christiano et al. (1997) where the central bank is supposed to control the money growth rate, and this affects the bank interest rate by way of the "limited participation" hypothesis. That is to say, households cannot re-adjust their bank deposits after a monetary shock, thereby producing a large impact of the monetary shock on bank loans. In our model this hypothesis in unncessary: indeed, given the official rate set by the central bank, money creation is endogenous via credit

This theoretical part of the paper yields a pattern of relationships consistent with the observed empirical regularities produced by monetary policy and highlights some key explanations. 1) The CCC affects firms by changing their real unit cost (the current real wage rate increased by the gross real (expected) bank rate) *vis-à-vis* their marginal product. Thus, for example, a higher official rate induces a higher bank rate which raises firms' unit cost and reduces labour demand and aggregate supply. This is typical of models with borrowed working capital. 2) In addition we also show that depending on (reasonable) characteristics of the production function and of households' behaviour, the overall effect of the monetary restriction is to reduce the real wage rate and aggregate demand. Inflation falls too, though to a limited extent. 3) A higher nominal bank rate after a monetary restriction is translated into a higher real rate not because the price level is sticky but because firms correctly anticipate a lower inflation rate. Consequently, the result is a *permanent* change in the output and inflation levels (i.e. corresponding to rational expectations equilibrium).

In section 3 we show the results of an econometric investigation of the CCC of monetary policy in Italy based on our theoretical model. Our aim is not to argue that the CCC is the sole monetary transmission mechanism. We instead wish to show that statistical analysis of the data yields a pattern of macroeconomic relationships consistent with the hypothesis that the CCC operates in the economy. All variables are monthly observations over the hard-EMS pre-EMU years 1986:1-1998:12. The time period ensures uniformity of data and the abandonment by the Bank of Italy of pervasive and recurrent administrative interventions that characterized the previous decades. As regards the two theoretical explanatory variables, credit risk and official rate, we have had to resort to proxies: the former has been obtained by an independent estimation procedure (see below) while for the latter we have conformed to the practice that takes the inter-bank rate as a policy driven variable.

We have chosen Italy as a case study because there are strong priors to believe that the CCC is prominent in this country, where, as a matter of fact, the contribution of marketable assets to corporate finance is much less

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market, and deposits are always in line with money creation as households leave their wage bill with banks.

important than elsewhere, while firms' bank-dependence is much more pervasive[9]. The role of the credit market in the monetary transmission mechanism has always been carefully monitored by the Italian monetary authorities and is explicitly included in the Bank of Italy's econometric model (1997a). A number of empirical studies, using various approaches and methodologies, have offered substantial support for the idea that credit is an important component in the monetary transmission mechanism in Italy[10]. However, to our knowledge, empirical research on the Italian case has taken no account of possible supply-side effects. A recent exception is Gaiotti and Secchi (2004), who find evidence of a cost channel of monetary policy at industry-level data. Yet they follow the Barth-Ramey (2001) approach, that is, industry partial equilibrium, with no explicit modelling of the credit market. Moreover, they assume imperfect competition in such a way that the cost channel is identified by a *positive* pass-through of the interest rate on prices. Yet, as clarified by our model, these are unnecessary conditions for the cost channel to operate.

Our econometric analysis, operating at the general-equilibrium macro-level, may complement the micro-approach with two advantages. The first is that it provides manageable aggregate models for macroeconomic and policy analysis, with a clear identification of the various links between monetary policy and economic activity at the same time. The second is that it may overcome major identification problems in traditional macro-models of monetary policy. As will be explained in section 3, our null hypothesis about the CCC implies a unique pattern of correlation signs between the official rate, on one side, and the real wage rate, output and inflation on the other, to the exclusion of alternative hypotheses such as “monetary ineffectiveness”, “nominal rigidity” or “monopolistic competition”.

Since the theoretical model also predicts that the CCC generates permanent effects, our econometric methodological choice has been to use

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[9]A long-standing body of empirical literature is available. See e.g. Vicarelli (1974), and among recent contributions Angeloni et al. (1997).

[10]Buttiglione and Ferri (1994), Bagliano and Favero (1998b), Chiades and Gambacorta (2000) have worked on macroeconomic models and aggregate data. Angeloni et al. (1995), Conigliani et al. (1997), Rondi et al. (1998) have examined microeconomic data from firms and banks. Angeloni et al. (2002) have used both micro and macro information. Bertocco (1997), who takes a "narrative approach", and Favero et al. (1999), who use balance-sheet data from a sample of European banks, are less supportive. For more details see Fiorentini-Tamborini (2002).

the structural cointegration approach suggested by Johansen and Juselius (2003). Firstly, whereas the traditional VAR models and the impulse-response exercises currently employed in the econometrics of monetary policy are mute about structural relationships and are fraught with arbitrariness in their imposition of identification structures (e.g. Bernanke and Mihov (1995), Bagliano and Favero (1998a), Rudebusch (1998)), the Johansen-Juselius methodology enables us to identify rigorously, through a testing procedure, the determinants of the long-run equilibrium paths along which the variables are moved, and around which short-run dynamics gravitate. Hence this methodology is best suited to running econometric analyses when a fully specified system of structural relationships is available.

Moreover, in view of policy analysis, Johansen and Juselius, in the same framework, also provide a rigorous definition of "control variable". The control rule that they define and demonstrate is based on an evaluation of the long-run impact of shocks on an instrument variable: shocks must have a significant long-run impact on the variables of the system.

Therefore, we have applied the Johansen-Juselius methodology first for the theory-based construction of a proxy for credit risk, second for the estimation and identification of the long-run structural relationships among the theoretically relevant variables, and third for the evaluation of the policy variable as a control variable of the system. Whereas the short-run dynamics after a monetary shock is transitory as argued by current common wisdom, the identified long-run structural relationships do not reject the CCC theoretical prediction that a monetary shock also permanently shifts the long-run path along which the real wage rate, output and inflation evolve. Moreover, we have also found that the inter-bank rate does qualify as control variable of the system according to the Johansen-Juselius criterion.

Section 4 concludes the paper with a summary of our main findings.

2. A model of the "credit cost channel" of monetary policy

We consider an economy with the following structure:

- the economy consists of three competitive markets, for labour, credit and goods (the single output of the economy), and hosts three representative classes of agents, households, firms and banks, and a central bank as the sole policy authority

- the economy operates sequentially along discrete time periods indexed by t, t+1, ..., where production takes 1 period of time regardless of the scale of production; firms can start a new production round only after "closing accounts" (i.e. the whole output has been sold out and all various claimants paid)[11]
- as firms plan production at the beginning of period t, they face uncertainty about revenue from output sales at t+1
- households sell labour and receive their income at the beginning of each period t, they can save to spend for consumption in t+1, whereas consumption for period t is bought out of saving carried over from t-1, etc.
- banks offer deposit services to households and debt contracts to firms, and can borrow reserves from the central bank.

To gain full understanding of the model, the reader should bear in mind the flow chart of the time structure of transactions illustrated in Figure 1

[Figure 1]

2.1. Firms

Firms are indexed with j, each producing a homogeneous output by means of a common labour technology with decreasing marginal returns and one-period production time, such that

(2.1) $$Q(t)_{jt+1} = Q(N_{jt}), \qquad Q_N > 0,\ Q_{NN} < 0$$

Let $Q(t)_{t+1}$ be total output in the economy and P_{t+1} its market-clearing price level. Each firm is price taker, hence it expects to sell its whole profit-maximizing output at the market-clearing price. Thus the firm has to plan output upon a forecast of the market-clearing price. As will be seen, P_{t+1} is a function of a set of parameters and variables observable at time t. We assume that at time t firms form their forecasts of P_{t+1} conditional on the available information set, but that owing to some

[11]In their original papers, Greenwald and Stiglitz considered firms that can rely on internal funds (past profits), which are a way to generate endogenous dynamics for business cycle analysis. We are not interested in this aspect here. Moreover, retained profits are not contemplated by the standard assumption that the competitive firm's revenue is exhausted by all claimants' incomes.

informational noise (e.g. "parameter uncertainty" or "measurement error" *à la* Pesaran (1987)) individual price forecasts P^e_{jt+1} differ across firms according to the function

(2.2) $$P^e_{jt+1} = P_{t+1}u_{jt}$$

where u_{jt} is a i.i.d. random variable with cumulative function F, unit expected value, $\mathrm{E}(u_{jt}) = 1$, and zero correlation across firms, $\mathrm{Cov}(u_{jt}, u_{it}) = 0$, $i \neq j$. This implies that each firm is subject to an individual forecast error $P^e_{jt+1} - P_{t+1} = P_{t+1}(u_{jt} - 1)$[12]. The individual values u_{jt} and the true realization of the firm's revenue are private information and cannot be observed by agents external to the firm[13].

Since production takes time, the firm has to hire labour in advance. Knowing that its revenue is subject to forecast error, with true realization being inside information, any external lender faces an ex-post verification problem of the firm's ability to pay which precludes efficient direct lending by households[14]. Thus the firm cannot promise to pay for labour from realized revenue, which would imply that workers lend in kind the value of labour to the firm, nor can it borrow the funds in advance from other households. Funds can instead be borrowed in the credit market by means of "standard debt contracts" with banks[15]. Hence, given the nominal wage

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[12]The key point in the model is that firms may make different errors, so that some, but not all, may default on debt (for instance, Greenwald and Stglitz introduce an ex-post noisy market-clearing process with different selling prices around the commonly expected market mean value P_{t+1}). We have chosen ex-ante different forecasts *vis-à-vis* a single market-clearing price beacuse this assumption is consistent with asymmetic information on the loan market (see below). A large literature has explained the origins and importance of heteregeneous price forecasts (see e.g. Frydman and Phelps (1983), Visco (1984), Pesaran (1987))

[13]Although P_{t+1} is freely observable by all market participants, the firm's actual revenue $P_{t+1}Q(N_{jt})$ depends on the level of output $Q(N_{jt})$, which, as is intuitive at this stage, in turn depends on the privately held price forecast $P_{t+1}u_{jt}$.

[14]As shown by Diamond (1984) and others, the ex-post verification problem may not have an efficient solution in a private financial market for two reasons: a free-riding problem (each private lender has no incentive to bear the cost of monitoring the actual state of the borrower if that information is to be immediately revealed to all the other lenders), and magnification of costs (it is not socially efficient for all lenders to monitor the same borrower separately). Intermediaries as “delegated controllers” provide an efficient solution to this problem.

[15]See Greenwald and Stiglitz (1993a). Freixas and Rochet (1998), explain the concept of standard debt contract in detail. Fiorentini and Tamborini (2002) have proved that under the assumptions made above, standard debt contracts are indeed the optimal financial contract.

rate w_t, the amount of bank loans demanded by a firm at time t is the value of its wage bill

(2.3) $$L^d_{jt} = W_t N_{jt}$$

against which the firm is committed to paying in t+1

- $L^d_{jt}R_t$ if the solvency state $P_{t+1}Q(t)_{jt+1} \geq L^d_{jt}R_t$ is declared
- $P_{t+1}Q(t)_{jt+1}$ if the default state $P_{t+1}Q(t)_{jt+1} < L^d_{jt}R_t$ is declared, with deterministic monitoring[16]

where $R_t \equiv (1 + r_t)$ is the gross nominal interest rate charged by banks.

Under these conditons, the firm's expected one-period profit is simply:

(2.4) $$Z^e_{jt+1} = P^e_{jt+1}Q(t)_{jt+1} - W_t N_{jt} R_t$$

Now let us assume that the firm wishes to maximize its expected stream of future real profits, Z^e_{jt+s}/P^e_{jt+s} , s = 1, ... Given the assumed operational conditions of the firm, the solution to this programme is simply the sequence of labour inputs N_{jt} that maximizes the one-period profits Z^e_{jt+1}. Denote with $w_t \equiv W_t/P_t$ the current real wage rate, with $\Pi^e_{jt+1} \equiv P^e_{jt+1}/P_t \equiv (1 + \pi^e_{t+1})$ the one-period expected price growth factor (expected inflation for short), and with $\underline{R}^e_{jt+1} \equiv R_t/\Pi^e_{jt+1} \equiv (1 + \underline{r}^e_{jt+1})$ the gross (expected) real interest rate. Hence, the first order condition for a real profit's maximum is

$$Q_N = w_t \underline{R}^e_{jt+1}$$

This condition states that the firm in each period t employs labour up to the point where its marginal product equals its *expected real unit cost*, which is the compound real cost of labour and credit. Under standard assumptions concerning the production function, and expanding the expression of $\underline{R}^e_{jt+1}$, the labour demand function can be written in the generic form

(2.5) $$N^d_{jt} = N^d(w_t, r_t, \pi^e_{jt+1}) \qquad N^d_{\underline{w}} < 0, N^d_{r} < 0, N^d_{\pi} > 0$$

We have thus obtained the typical labour demand function of the bank-dependent firm. In addition to the usual negative relationship with

[16]Deterministic monitoring means that the firm is monitored and its true state is observed with certainty. We exclude here that the default state has pecuniary or non-pecuniary extra-costs for the firm. These costs are assumed to be present and play an important role in the Greenwald-Stiglitz model because they force the firm to take account of the expected default cost in profit maximization, which thus appears in the aggregate supply function. Yet this procedure requires auxiliary assumptions on the nature of these costs. We propose a simpler solution: there are no such costs, but, as we shall see, default probability enters the model via bank interest rates in the form of credit risk premium.

the real wage rate, the main features of labour demand due to bank debt are: 1) it is decreasing in the real interest rate, 2) it is systematically lower than the standard labour demand for any positive interest rate.

Output supply is derived from labour demand by means of the production function (2.1), i.e.:

(2.6) $$Q(t)_{jt+1} = Q(N^d(w_t, r_t, \pi^e_{jt+1}))$$

Moreover, from the previous equations we can obtain an implicit probability measure of default. In fact, default occurs in all states such that

$$P_{t+1}Q(t)_{jt+1} < L^d_{jt}R_t$$

or, after little manipulation,

$$P_{t+1} < V_{jt}$$

where $V_{jt} \equiv L^d_{jt}R_t / Q(t)_{jt+1}$ is the firm's debt-output ratio at time t. According to equations (2.3), (2.5) and (2.6), it turns out that, in the optimal production plan,

$$V_{jt} = V(w_t, r_t, \pi^e_{jt+1}) \qquad V_{\underline{w}} < 0,\ V_r < 0,\ V_\pi > 0$$

Note that $V_\pi > 0$ because higher expected inflation *cet. par.* lowers the firm's expected real unit cost, then the firm increases planned output, employment and borrowing, and its debt-output ratio is eventually increased as well[17]. Since $\pi^e_{jt+1} \equiv (P_{t+1}/P_t)u_{jt} - 1$, given the observed P_t, and for any P_{t+1}, the level of π^e_{jt+1} depends on the price forecast noise of the firm u_{jt}. Therefore, the higher u_{jt} and π^e_{jt+1}, the higher V_{jt} relative to P_{t+1}. Consequently, while the firm operates on the profit-maximizing expectation that $V_{jt} < P^e_{jt+1}$, there will be a critical u^*_{jt} such that $V_{jt} = P_{t+1}$. Moreover, since all firms operate under the same technological and market conditions, this critical value will be the same for all firms, u^*_t.[18] Hence, knowing the function $F(u_{jt})$, the default probability of a firm is the probability of a forecast noise exceeding u^*_t, i.e.

(2.7) $$\text{Prob}(u_{jt} \geq u^*_t) = 1 - F(u^*_t) \equiv \phi_t, \text{ all } j$$

Labour demand (2.5), output supply (2.6), and default probability (2.7) form the core of the supply-side economics of the credit channel of monetary policy. They show that firms' employment and production decisions are affected by credit supply conditions to the extent that 1) these

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[17]This outcome follows from the assumption that the production function is concave. In fact, $V_\pi > 0$ if $Q(N^d)/N^d > Q_N$.

[18]Let us re-write the equation of V_{jt} in the extended form $V(w_t, r_t, P_{t+1}, u_{jt})$. The critical value of u_{jt} is the solution to $V(w_t, r_t, P_{t+1}, u_{jt}) = P_{t+1}$, i.e. $u^*_t = u(w_t, r_t, P_{t+1})$, which is the same for all j because so are w_t, r_t, P_{t+1}.

entail changes in the bank interest rate, *and* 2) these changes are to some extent transmitted to the expected real unit cost, i.e. to the extent that movements in the real wage rate, in the nominal interest rate and in expected inflation do not exactly offset each other. This a crucial aspect of the model, and its importance will become clearer once aggregate supply has been embedded in the rest of the economic system.

2.2. Households

Households perform three activities: labour supply, output demand and saving. As a counterpart of firms' equity rationing in the manner that has been explained in paragraph 2.1, the scope of households' assets is limited to bank deposits at zero interest rate[19]. According to the time structure of the economy, at the beginning of each production period t each household h owns the previous period's deposits $D_{ht\text{-}1}$, can sell labour $N^s{}_{ht}$ at the market wage rate W_t, earns $W_t N^s{}_{ht}$ and saves D_{ht} for consumption in t+1 $C(t)_{ht+1}$, on the expectation of a price $P^e{}_{ht+1}$. For simplicity, households price expectations are assumed to be regulated by the same law (2.2) as those of firms.

Owing to the assumed time structure of transactions, consumption for period t, $C(t\text{-}1)_{ht}$, can only be satisfied by t-1 production $Q(t\text{-}1)_t$ by spending $D_{t\text{-}1}$ deposits at the price P_t before the new production round starts, whereas consumption for t+1 $C(t)_{ht+1}$ will be satisfied by $Q(t)_{t+1}$, and so on. Therefore, the general representation of the household's problem at the beginning of any production period t is a sequence of choices $\{C_h$: $C(t)_{ht+1}$, $C(t+1)_{ht+2}$, ...$\}$, $\{N_h$: N_{ht}, N_{ht+1}, ...$\}$ such that

(2.8) $$\max_{C,N} U_{ht} = U(C_h, N_h), \qquad U_C > 0,\ U_N < 0$$
$$\text{s.t. } P_t C(t\text{-}1)_{ht} \leq D_{t\text{-}1}$$
$$P^e{}_{ht+1} C(t)_{ht+1} \leq D_t$$
$$D_t = D_{t\text{-}1} - P_t C(t-1)_{ht} + W_t N_{ht}$$

Using P_t as numeraire, the generic form of the solution to the foregoing problem includes a labour supply function of the form:

(2.9) $$N^s{}_{ht} = N^s(w_t, \pi^e{}_{ht+1}) \qquad N^s{}_{\underline{w}} > 0,\ N^s{}_{\pi} < 0$$

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[19]This is of course a strong limitation. However, the inclusion of interest-bearing bonds (as e.g. in Bernanke and Blinder (1988)) would complicate the transmission mechanism from monetary policy to credit supply with no substantial gain for our purposes.

This function reflects the household's choice between working time and consumption in t, *and* between consumption in t and t+1. Note that in this model, the real wage rate w_t, i.e. the relative value of W_t with respect to P_t, measures the incentive to work in period t in view of consumption in t+1, given the amount of consumption obtained for period t. In fact, for any given value of D_{t-1}, P_t acts as wealth effect, determining the amount of consumption goods $C(t\text{-}1)_{ht}$ with which the household enters the production period. *Cet. par.*, high $C(t\text{-}1)_{ht}$, i.e. low P_t relative to W_t, requires a parallel increase in $C(t)_{ht+1}$ to match the intertemporal marginal rate of substitution along the optimal consumption path, and therefore the household should also increase labour supply. At the same time, the cross-elasticity between *present* working time and *future* consumption is not nil: higher π^e_{ht+1}, i.e. higher P^e_{ht+1} relative to P_t, shifts resources from future to current consumption of goods (i.e. output and leisure) so that the labour supply is decreasing in π^e_{ht+1}[20].

Looking at the constraints of the household's problem we can also deduce a generic function for consumption, which at the end of each period can be equal to, or less than, the real value of deposits:

(2.10) $$C(t)_{ht+1} = C(D_t)\ /P_{t+1} \qquad\qquad C(D_t) \leq D_t$$

The possibility that $C(D_t) < D_t$ arises because at the end of each period households cannot spend more than previous period's deposits, but may choose to spend less and carry more resources to the next period depending on their intertemporal preferences. In any case, the result is a simple demand function determined by real money balances, and, for algebraic simplicity, we shall set $C(D_t) = D_t$

2.3. Banks and central bank

Banks collect deposits from households at zero rate, can borrow from the central bank at the given official rate, and offer standard debt contracts to firms in a competitive credit market. Since firms are ex-ante homogenous, banks face no screening problems. However they bear monitoring costs whenever a firm defaults on payments (see paragraph 2.1). Credit recovery

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[20]The fact that households, too, decide under price uncertainty (the actual price P_{t+1} that the household will pay may differ from its expected price P^e_{ht+1} with positive probability) may mean that their decisions depend on the specific form of the utility function (e.g. a higher level of precautionary saving in case of risk aversion), but it does not substantially modify the model's main properties.

in the event of bankruptcy of any firm j is $P_{t+1}Q(t)_{jt+1}$. Since the incentive to monitor firms exists up to equality between credit recovery and monitoring cost, without loss of generality we can set the net revenue from defaulting firms to zero[21]. As to the cost of funds, in the absence of the interest rate on deposits, we introduce a kind of cost which is important in bank's risk management and gives the central bank an explicit role to play by way of the official rate.

In consideration of the time structure of the economy, banks' balance sheets evolve intertemporally over production periods. At the beginning of each t, a bank b can grant loans $L^s{}_{bt}$. Loans finance the wage bill for period t and are therefore redeposited by households as savings for period t+1 (see paragraphs 2.1, 2.2). Hence the resulting balance sheet is

(2.11) $$L^s{}_{bt} = D_{bt}$$

In view of the fact that households will claim on D_{bt} one period later, the bank should secure itself a sufficient amount of liquid resources. This requirement acts as a liquidity constraint on the bank's decision problem[22]. The bank expects a gross return from loans $Z^e{}_{bt+1}$. As is clear from (2.11), if all firms repaid capital, the bank would be certain that its liquidity constraint would be satisfied. Yet, as explained in paragraph 2.1, each loan at time t embodies a default risk of the firm. All firms face the same default probability ϕ_t given by (2.7), and if the loans portfolio of the bank is well diversified it will only bear the undiversifiable risk ϕ_t. Therefore, recalling that the bank expects zero net revenue at time t+1 from each defaulting firm, it anticipates a liquidity risk (the probability of capital repayments falling short of deposits) equal to $L^s{}_{bt}\phi_t$ associated with its loans portfolio. The bank can insure itself against this risk by borrowing reserves BR_t from the central bank at the gross official interest rate $K_t \equiv (1 + k_t)$, i.e. it can cover all illiquidity states $L^s{}_{bt}\phi_t$ under the obligation to repay $L^s{}_{bt}\phi_t K_t$ in t+1.

Hence the bank's expected gross return on the loans portfolio at the market gross rate R_t is $Z^e{}_{bt+1} = L^s{}_{bt}R_t(1 - \phi_t)$, while its expected net profit is

(2.12) $$L^s{}_{bt}R_t(1 - \phi_t) - L^s{}_{bt}K_t\phi_t - L^s{}_{bt} \geq 0$$

---

[21]See Fiorentini-Tamborini (2002) for detailed analysis.

[22]The amount of liquid resources is generally given by the statistical expectation of withdrawals, i.e. the aggregation of the individual withdrawals $C(D_{ht}) \leq D_{ht}$ from our households' model. Having assumed $C(D_{ht}) = D_{ht}$ for all h, the bank's liquidity constraint should strictly hold as equality.

Competitive pressure will drive this expression to equality, with the bank interest rate equal to[23]

(2.13) $$R_t = \frac{1+\phi_t K_t}{1-\phi_t}$$

The result is that the bank interest rate is determined as an increasing risk-adjusted function of the official rate, and is also increasing in risk. A simple algebraic manipulation allows a more transparent interpretation of this result. The actual interest rates can be approximated by $r_t \approx \log R_t$ and $k_t \approx \log K_t$. In addition, if k_t is a small fractional number, i.e. around $K_t = 1$, the logarithm of expression (2.13) is closely approximated by

(2.14) $$r_t \approx \rho_t + k_t$$

where $\rho_t = \log\frac{1+\phi_t}{1-\phi_t}$ is a proxy for credit risk increasing in ϕ_t, so that r_t can be interpreted as the sum of the official rate plus a credit risk premium.

At the rate r_t all credit demand is satisfied, while the amount of borrowed reserves, $BR_t = L^s{}_{bt}\phi_t$, is created. Borrowed reserves also appear in the central bank's balance sheet as monetary base. Since in this model the central bank pegs the interest rate, the creation of monetary base is endogenous, being equal to the risk-adjusted fraction of bank loans.

2.4. Macroeconomic equilibrium and the effects of monetary policy

The complete model is given by the intertemporal equilibrium conditions of the labour and credit markets in each period *t,* and of the output market in the subsequent period *t*+1. In the previous paragraphs we have obtained the behavioural functions of a representative firm, household and bank. To move to macroeconomic equilibrium we need aggregate functions. The functions of individual firms (households) only differ as a consequence of different price forcast errors, but, given the assumptions on price expectations formation in paragragh 2.1, aggregation across firms (households) ensures that the rational expectations hypothesis holds in the aggregate.

Therefore we can summarize the complete macroeconomic equilibrium as follows.

---

[23] Note that the following result is independent of expected inflation. In fact, if we take the bank's expected real net profit and deflate (2.12) by expected inflation, we still obtain (2.13).

Labour market

(2.15) $N^d(w_t, r_t, \pi_{t+1}) = N^s(w_t, \pi_{t+1})$

Credit market

(2.16) $L_t = W_t N_t$

(2.17) $r_t \approx \rho_t + k_t$

Output market

(2.18) $Q(N^d(w_t, r_t, \pi_{t+1})) = D_t / P_{t+1}$

$\pi_{t+1} = P_{t+1}/P_t - 1$

We shall examine the comparative statics of the system after two exogenous impulses from what we may call the "credit variables": $\mathrm{d}k_t$, which represents a monetary policy intervention, and $\mathrm{d}\rho_t$, which represents an exogenous shock to credit supply. We shall give the results of three endogenous variables: the real wage rate w_t, output $Q(t)_{t+1}$ and the inflation rate π_{t+1}. The remaining variables, the interest rate r_t, employment N_t and bank loans L_t can easily be derived from equations (2.15) (2.16), and (2.17). Following previous notation, variables typed as subscripts denote first derivatives and are expressed in absolute values[24].

(2.19)
$$\begin{bmatrix} \mathrm{d}w_t \\ \mathrm{d}Q(t)_{t+1} \\ \mathrm{d}\pi_{t+1} \end{bmatrix} = \frac{1}{\Delta} \begin{bmatrix} N^d_w(1 + N^s_\pi(1 - Q_N)) \\ N^d_w Q_N (N^s_w - N^s_\pi) \\ N^d_w(1 + N^s_w(1 - Q_N)) \end{bmatrix} (\mathrm{d}k_t + \mathrm{d}\rho_t)$$
$$\Delta = (1 + N^d_w(1 - Q_N))(N^s_\pi - N^s_w)$$

As can be seen from system (2.19), the thrust of our model is that *both* k_t and ρ_t can exert non-zero permanent effects on the rational expectations equilibrium level of w_t, $Q(t)_{t+1}$ and π_{t+1}. In particular, if $N^s_\pi < N^s_{\underline{w}}$ and $Q_N <$ 1, changes in k_t and ρ_t have negative effects on w_t, $Q(t)_{t+1}$ and π_{t+1} which fit the empirical regularities observed in major industrialized countries. The transmission mechanism of the two credit variables is the same, and can be explained with the help of Figure 2.

[Figure 2]

Assume the economy is in equilibrium at points A, and consider the case that the central bank raises the official rate $\mathrm{d}k_t > 0$. As a consequence,

[24]Form equation (2.5), the obvious restrictions $N^d_w = N^d_r = N^d_\pi$ applies.

the nominal bank rate r_t increases (left-hand panel) and generates a higher real rate for firms (due to higher r_t *and* the rational expectation of a lower inflation π_{t+1}). Hence labour demand shifts downwards (right-hand panel). As the nominal wage rate W_t falls relative to P_t, households are induced to supply less labour (along the previous supply schedule), while the expected lower inflation induces them to supply more (shifting the supply schedule downwards), placing further competitive pressure on w_t. The condition $N^s{}_\pi < N^s{}_{\underline{w}}$ implies that this latter effect is weaker than the former, so that the overall fall of the real wage rate does not compensate for the rise in the real interest rate, thus leaving firms with higher real unit cost[25]. The consequence is a net cut in employment and output. Less employment at a lower wage rate generates fewer bank loans and deposits: hence, the next period's aggregate demand and supply are both reduced. $Q_N < 1$ entails that the demand shift is larger than the supply one, so that the inflation rate actually falls[26].

We can summarize our theoretical findings in two main points. First, inclusion of the supply-side effects of monetary policy yields a pattern of relationships consistent with the set of empirical regularities that are today regarded as the *explanandum* of monetary macroeconomics. This result is in line with most of the recent literature. Second, the CCC transmission mechanism also shows that this macroeconomic pattern does not depend on major deviations from competitive behaviour in labour and goods markets,

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[25] This misalignment between real wage and real interest rate was what Keynes strived to show in chapter 13 of the *General Theory*. The problem is *not* (necessarily) one of spot wage stickiness, but is intertemporal in nature (see Leijonhufvud (1968, 1981), Greenwald and Stiglitz (1993b), Hahn and Solow (1995)). In our model, the relative magnitude of $N^s{}_\pi$ and $N^s{}_{\underline{w}}$ measures the responsiveness of labour supply to the intertemporal substitution effect relative to the current consumption effect. As is well known, this factor has played a major role in the development of modern business cycle theory. Theories that, in order to fit observable comovements between real or nominal impulses and output (employment), postulate large intertemporal substitution effects (as in the standard versions of real business cycle models) have been impaired by their inability to detect that condition in the data. By contrast, the credit transimission mechanism identified by our model does yield the observable correlations $\mathrm{d}Q(t)_{t+1}/\mathrm{d}k_t < 0$, $\mathrm{d}\pi_{t+1}/\mathrm{d}k_t < 0$ *thanks to a relatively small intertemporal substitution effect.*

[26] The condition $Q_N < 1$ is consistent with non-increasing returns in a large class of production functions (the Cobb-Douglas function is the typical example). Hence we can consider this condition as sufficiently general according to current theoretical standards.

but hinges on two simple primitive conditions on the capital market (and reasonable elasticities in labour demand and aggregate supply): firms' dependence on bank credit, and the central bank's ability to drive bank interest rates. Consequently, the presumption arises that the CCC may also have permenent, rather than transitory, effects on real variables.

3. Econometric analysis

We now present the results of an econometric analysis aimed at identifying and estimating the supply-side effects of monetary policy in the Italian economy, in the hard-EMS pre-EMU period, i.e. 1986:1-1998:12. Our econometric investigation employed a testable version of the quasi-reduced form (2.19) of the CCC model presented in the previous section. Details on statistical procedures and tests are gathered in a separate Appendix available on request.

3.1. Data and methodology

According to our theoretical model, the variables of interest are:

- the real wage rate w_t, measured by the industrial wage index at the producer cost;
- the monetary policy variable k_t, for which we have taken the inter-bank rate[27];
- the credit risk premium, ρ_t, as a measure of autonomous shifts in credit supply; this variable is not observable and we elaborated a proxy: our procedure is described in the Appendix A2, where further relevant data are the average bank lending rate, r_t, and total loans to the private sector, L_t;
- output Q_{t+s}, given by the industrial production index, seasonally adjusted;
- the inflation rate, π_{t+s}, measured by the consumer price index.

The time lead s should capture the theoretical gestation time of output and the related time-horizon for expected inflation[28]. A time lead of 12

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[27] The inter-bank rate, being highly sensitive to central bank's interventions and a close determinant of bank lending rates (see also Figure 3), is typically taken as the closest market indicator of monetary policy in almost all available empirical studies of the credit channel in Italy (see e.g. De Arcangelis and Di Giorgio (1998) and fn. 10 for references).

months was chosen empirically by means of sensitivity tests. This seems a reasonable figure in consideration of the average length of short-term loans to firms; it is also consistent with values found in other similar works on Italy (Bagliano and Favero (1998b), Fiorentini and Tamborini (2002)).

In consideration of the fact, not made explicit in the theoretical model, that Italy is an open economy we also added a foreign variable to control for possible foreign effects and constraints on domestic monetary policy and interest rates in a period of strong exchange-rate targeting and high capital mobility in the EMS, namely

- the German inter-bank rate, k^*_t.[29]

All variables are log-transformed monthly times series, and are plotted with sources in Figure 3.

[Figure 3]

We applied the structural cointegration methodology developed by Johansen and Juselius (2003) first for the theory-based construction of the proxy for the credit risk premium, second for the estimation and identification of the long-run structural relationships among the theoretically relevant variables, and third for the evaluation of the policy variable as a control variable of the system[30]. This methodology is

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[28]Since we assume flexible prices and rational expectations, we can take the actual inflation rate on the same time lead as output as a proxy for the theoretical expected inflation.

[29] The literature on monetary policy in the EMS (see e.g. De Grauwe (1992)) would predict that a country like Italy could not misalign the domestic interest rate systematically from the uncovered parity with the German one as implied by

$$k_t - k^*_t = \mathrm{E}_t(\dot{e}) \to 0$$

where $\mathrm{E}_t(\dot{e})$ is the expected depreciation rate. However, temporary non-zero interest differentials would still be possible as long as the implied expected change in the exchange-rate remained within the band of the parity. In this view, a monetary policy shock could be identified by a deviation from uncovered interest parity, i.e. a non-zero interest-rate differential. Suppose k^*_t rises in Germany while k_t remains constant in Italy: the interest rate differential in Italy *falls*. Given the commitment to the exchange-rate parity, this is perceived as a *positive* monetary shock. Therefore, we have introduced the two inter-bank rates as two independent variables and let the data say to what extent they actually exerted independent effects. It is worth noting that the introduction of the German inter-bank rate substantially improves the overall quality of the estimates.

[30] All the empirical analysis was performed by using the software CATS and the software MALCOLM. Both need the package RATS to be run. The results are available upon request.

expounded in the Appendix A1, while here we simply summarize the main steps and results.

3.2. The unrestricted cointegrated model

In the first place, we started from a p-variables VAR(n) format of our theoretical model, with a p = 5 variables vector $\mathbf{y'}_t = [w_t, k_t, \rho_t, q_{t+12}, \pi_{t+12}]$, augmented to include 1 exogenous variable $\mathbf{z'}_t = [k^*_t]$, a constant term μ_0, a linear trend t, a vector of seasonal and intervention dummies $\mathbf{D}_t$[31] and a vector of normal disturbances $\boldsymbol{\varepsilon}_t$. Maximum lag analysis led us to choose a lag parameter n = 3. On the basis of multivariate and univariate mis-specification tests the statistical model appeared acceptable.

In the second place, we moved to an "error correction" (VECM) respecification

$$\text{(3.1)} \qquad \Delta\mathbf{y}_t = \Gamma_0\Delta\mathbf{z}_t + \sum_{i=1}^{n-1}\Gamma_i\Delta\mathbf{x}_{t-i} + \Pi\mathbf{x}_{t-1} + \mu_0 + \mu_1 t + \Phi\mathbf{D}_t + \boldsymbol{\varepsilon}_t$$

where Δ is the first difference operator, $\mathbf{x'}_t = [\mathbf{y'}_t, \mathbf{z'}_t]$, and $\mathbf{\Gamma}$, $\mathbf{\Pi}$, $\mathbf{\Phi}$ are matrices of coefficients. Note that $\mathbf{\Pi x}_{t\text{-}1}$ contains the error correction mechanism, where $\mathbf{\Pi}$ is a matrix whose rank, $0 < r < p$, corresponds to the number of stationary relations, or cointegrating relations, between the variables of the model. Testing for the cointegration rank r is connected with determining the appropriate trend polynomial. Given that in the cointegrating relations we initially chose a linear trend[32], based on the 90% critical value for the *trace* statistic, the selected rank would be r = 3. Some caution is needed because we included exogenous and dummy variables in the model and the quantiles of the distribution are not appropriate in this case. We further performed a simple graphical analysis based on the stability of the cointegration rank that supported the decision of r = 3. Then it was possible to factorise $\mathbf{\Pi}$ into $\mathbf{\Pi} = \boldsymbol{\alpha\beta}'$, where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are both $(p\times r)$ matrices which convey important economic information. The columns of matrix $\boldsymbol{\beta}$ are the cointegration vectors, that is the *long-run relationships*

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[31] In order to obtain residuals close to Normality, in our data set we introduced four permanent intervention dummies to account for the exit of Italian Lira from the EMS in 1992 and for a couple of other events. Moreover, there were eleven centered seasonal dummies. The permanent intervention dummies were defined for 1992/VII, 1992/IX, 1991/V and 1993/III.

[32] Given linear trends in the data, this choice is generally the best specification to start with, unless we have a strong prior hypothesis that the trends cancel in the cointegration relations.

that can be dectected among the variables ("attractor set"), whereas the elements of matrix $\boldsymbol{\alpha}$ are the adjustment coefficients of variables *towards* their long-run relationships. Also, we were able to detect $(p-r) = 2$ common trends: these are associated with a matrix, $\boldsymbol{\alpha}_{\perp}$, whose elements measure how cumulated stochastic shocks push the variables *along* their long-run relationships[33].

Table 1 reports the decomposition of vector $\mathbf{x}_t$ into the r cointegrating relations and the $(p-r)$ non cointegrating relations, together with the adjustment coefficients and the common trend coefficients. Upon the finding of 3 cointegrating relations, and upon the information given by the covariances observed in the data, we have normalized the unrestricted relations with respect to the 3 variables that theory indicates as "endogenous" (π_{t+12}, w_t, q_{t+12}) *vis-à-vis* the CCC "explanatory" (k_t, ρ_t, k^*_t) variables and the trend.

Table 1. The I(0) and I(1) components of $\mathbf{x}_t$ (bold α coefficients denote significance at 5%)

	π_{t+12}	w_t	q_{t+12}	k_t	ρ_t	k^*_t	$trend_t$
	The cointegrating matrix $\hat{\boldsymbol{\beta}}$ (transposed)						
$\hat{\boldsymbol{\beta}}_1'$	1.000	0.048	0.693	1.701	3.476	-1.049	0.003
$\hat{\boldsymbol{\beta}}_2'$	4.736	1.000	-1.535	0.445	-1.614	-3.073	0.001
$\hat{\boldsymbol{\beta}}_3'$	-0.663	0.551	1.000	0.779	0.799	-0.252	-0.001
	The adjustment coefficient matrix $\hat{\boldsymbol{\alpha}}$ (transposed)						
$\hat{\boldsymbol{\alpha}}_1'$	-0.016	0.033	-0.031	**-0.151**	0.003		
$\hat{\boldsymbol{\alpha}}_2'$	**-0.012**	**-0.058**	0.050	0.005	0.000		
$\hat{\boldsymbol{\alpha}}_3'$	**0.024**	-0.051	**-0.305**	-0.016	-0.006		

The common stochastic trend matrix $\hat{\boldsymbol{\alpha}}_{\perp}$ (transposed)
(residual standard deviations in brackets)

	$\hat{\varepsilon}_{\pi_{t+12}}$	$\hat{\varepsilon}_{w_t}$	$\hat{\varepsilon}_{q_{t+12}}$	$\hat{\varepsilon}_{k_t}$	$\hat{\varepsilon}_{\rho_t}$
	(0.0018)	(0.0049)	(0.0154)	(0.0037)	(0.0005)
$\hat{\boldsymbol{\alpha}}_{1\perp}'$	-0.8970	0.1157	-0.0898	0.1282	-0.3969
$\hat{\boldsymbol{\alpha}}_{2\perp}'$	-0.3837	0.0363	-0.0592	0.0805	0.9173

[33]Mis-specification tests are detailed in Appendix A3.

The estimated unrestricted vectors forming the matrix $\hat{\boldsymbol{\beta}}$ may not have an economic meaning, since economic interpretation of the cointegrating relations is difficult without imposing identifying restrictions on the vectors. This will be done below, but beforehand it is convenient to draw some other information from table 1. As regards the adjustment coefficients in matrix $\hat{\boldsymbol{\alpha}}$, it should be noted that the risk premium enters no cointegrating relation significantly[34]. As a zero row of $\boldsymbol{\alpha}$ is the condition for the corresponding variable to be weakly exogenous w.r.t. the cointegration relations, our measure of risk premium can indeed be taken as exogenous, as required by the theoretical model[35]. The estimated matrix $\hat{\boldsymbol{\alpha}}_{\perp}$ shows that in the first common trend the estimated cumulated shocks exhibit the largest (absolute) weight for the inflation rate, while in the second common trend the largest weight is associated with the risk premium. However, unless the residuals are normalized the weights are not very informative. We have therefore given the standard deviation underneath each residual. It can be seen that in the second common trend, even though the coefficient associated with the risk premium is almost sixteen times larger than that associated with industrial production, the residual standard deviation of the latter is thirty times larger than that of the former. In other words, the two common trends appear to be driven mainly by unanticipated cumulated shocks to the inflation rate and to industrial production. In other words, the two common trends appear to be driven mainly by unanticipated cumulated shocks to the inflation rate and to industrial production[36].

[34] The other coefficients show that the first cointegration relation is error adjusting only in the inter-bank rate, while the second one is adjusting in the real wage and in the inflation rate; the third one is adjusting in the industrial production whereas it is overshooting in the inflation rate.

[35] The relevant hypothesis to test takes the form $\mathrm{H}^{c}_{\alpha}(r): \mathbf{R}'\boldsymbol{\alpha} = \mathbf{0}$, where the matrix $\mathbf{R}$ becomes the following row vector: $\mathbf{R}' = [0,0,0,0,1]$. The *LR* test statistic, distributed as a χ^2_3, is equal to 4.09, with a *p-value*=0.25; therefore the hypothesis is clearly accepted. According to Garratt et al. (2003), risk is then considered to be a "long-run forcing" variable, and, under this assumption, the cointegrating properties of our model can be analysed without having to specify a risk equation.

[36] Though the cumulated residuals in the equations for these two variables can be considered common trends, this does not imply that these two variables are the common trends. For this to be the case they would have to be weakly exogenous and we would have no significant short run effects in the corresponding equations of the estimated VECM model.

3.3. Identifying the long-run structure

The system of cointegrating relations can be written as follows:

$$
\begin{aligned}
&\pi_{t+12} = \beta_{11}w_t + \beta_{12}q_{t+12} + \beta_{13}k_t + \beta_{14}\rho_t + \beta_{15}k^*_t + \beta_{16}t + u_{\pi_{t+12}} \\
(3.2)\qquad &w_t = \beta_{21}\pi_{t+12} + \beta_{22}q_{t+12} + \beta_{23}k_t + \beta_{24}\rho_t + \beta_{25}k^*_t + \beta_{26}t + u_{w_t} \\
&q_{t+12} = \beta_{31}w_t + \beta_{32}\pi_{t+12} + \beta_{33}k_t + \beta_{34}\rho_t + \beta_{35}k^*_t + \beta_{36}t + u_{q_{t+12}}
\end{aligned}
$$

This system and its underlying theory allow rigorous identification and testing of the hypothesis that monetary policy interventions, made through the inter-bank rate, affect output and inflation by way of supply-side effects triggered by the CCC. We can proceed in two steps: firstly, we discuss the expected signs of the coefficients; secondly, we impose some identifying restrictions.

The CCC transmission mechanism, to which we refer as the null hypothesis, was explained at the end of section 2 (see also Figure 2): it hinges on the signs of the CCC variables k_t, ρ_t and k^*_t, that we interpret as explanatory variables, in the equations for π_{t+12}, w_t, and q_{t+12} , and implies the unique pattern of signs of coefficients in Table 2.

Table 2. The pattern of coefficients' signs of the CCC hypothesis

H_0	k_t	ρ_t	k^*_t
π_{t+12}	$\beta_{13} < 0$	$\beta_{14} < 0$	$\beta_{15} > 0$
w_t	$\beta_{23} < 0$	$\beta_{24} < 0$	$\beta_{25} > 0$
q_{t+12}	$\beta_{33} < 0$	$\beta_{34} < 0$	$\beta_{35} > 0$

The analogous signs of k_t and ρ_t reflect the theoretical prediction that k_t, a policy driven variable, operates on the credit-cost component of firms' plans like ρ_t, an autonomous measure of changes in credit cost. The simultaneous presence of the two variables, as long as they are independent, should improve identification of the theoretical prediction. The opposite signs of k^*_t and k_t instead reflect the empirical specification of the uncovered interest differential with Germany explained above in paragraph 3.1 (where strict uncovered interest parity under fixed exchange rate, $k_t - k^*_t = 0$, would imply equality of the two coefficients in absolute value; see also fn. 29).

Our expected pattern of coefficients excludes different alternative hypotheses. First, "monetary ineffectiveness" would require most of the adjustment to take place in the inflation rate with non-significant effects on the real wage rate and output ($\beta_{23} = \beta_{33} = \beta_{25} = \beta_{35} = 0$)[37]. Second, "nominal rigidity" would yield the negative correlation of output and inflation with the explanatory variables as a result from invariant nominal wage rate *vis-à-vis* the expected change in the price level, which implies that the contemporaneous real wage rate is uncorrelated with the explanatory variables ($\beta_{23} = \beta_{24} = \beta_{25} = 0$). Third, the "monopolistic competition" version of the cost channel of monetary policy would predict the negative correlation of output with the explanatory variables in association with a positive pass-through on the price level ($\beta_{33} > 0$, $\beta_{34} > 0$, $\beta_{35} < 0$).

We now move to the restriction matrices that the estimated cointegration vectors $\hat{\boldsymbol{\beta}}$ should satisfy in order to have unique structural cointegrating relations[38] (Johansen (1995)). To this effect, the forward-looking sequential structure of the theoretical model can be exploited. In fact, the model says that at any point in time future output depends on the current real wage rate and on the CCC variables corrected for the expected inflation rate. The real wage rate in turn responds to the CCC variables and the expected inflation rate via planned output. Eventually, assuming that the inflation expectations are given by the statistical expected value of the actual inflation rate, the latter also comes to depend on the past CCC variables as planned output arrives to the market. Consequently, the unrestricted system (3.2) can be restricted to a quasi-reduced "pyramid" form simply by setting $\beta_{11} = \beta_{12} = \beta_{22} = 0$. In Table 3 we report the identifying structure, and below the final identified long-run relationships (standard errors in brackets) [39].

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[37] On this view, only shocks to credit risk may be regarded as real shocks admitting of real adjustments.

[38] See Appendix A1 for procedure.

[39] We also tested further restrictions in order to identify the strict uncovered interest parity, i.e. $\beta_{i3} = \beta_{i5}$, which was not rejected in some of the i relations but not in all of them.

Table 3. Identified structure on the cointegration vectors (bold coefficients denote significance at 5%)

π_{t+12}	w_t	q_{t+12}	k_t	ρ_t	k^*_t	$trend_t$
1.000	0.000	0.000	**0.650**	**1.477**	**-0.685**	**0.002**
2.449	**1.000**	0.000	**1.254**	0.073	**-2.213**	0.000
-0.731	**0.360**	**1.000**	**0.906**	**0.906**	-0.156	**-0.001**
$\chi^2_3 = 4.09$, p - value $= 0.25$						

$$\pi_{t+12} = -0.650k_t - 1.477\rho_t + 0.685k^*_t - 0.002t + \hat{u}_{\pi_{t+12}}$$
$$(0.089) \quad (0.160) \quad (0.089) \quad (0.000)$$

(3.3)
$$w_t = -2.449\pi_{t+12} - 1.254k_t - 0.073\rho_t + 2.213k^*_t + \hat{u}_{w_t}$$
$$(0.506) \quad (0.350) \quad (0.258) \quad (0.281)$$

$$q_{t+12} = -0.360w_t + 0.731\pi_{t+12} - 0.906k_t - 0.906\rho_t + 0.15k^*_t - 0.001t + \hat{u}_{q_{t+12}}$$
$$(0.106) \quad (0.355) \quad (0.213) \quad (0.213) \quad (0.292) \quad (0.000)$$

Recall that these statistical equations detect the long-run relationships that the normalized variables entertain with the other variables once short-run dynamics has petered out. In other words, they identify the determinants of the long-run equilibrium stochastic paths along which the l.h.s variables are moved, and around which their short-run dynamics gravitate. These relationships are broadly consistent with the theoretical model.

First, the inter-bank rate k_t, and more specifically k_t corrected for uncovered interest parity with k^*_t and for the credit risk premium ρ_t always has the expected, significant sign on inflation, real wage rate and output. As explained previously, this can be considered evidence that this variable operates through the supply side of the economy, that it has long-run real effects, and that these effects are not consistently explained by the nominal rigidity or the monopolistic competition hypotheses.

Second, further independent evidence is provided by ρ_t. It proves consistent with the null hypothesis in the inflation equation and in the output equation. According to these equations, the credit risk premium is indeed a co-determinant, with the inter-bank interest rate, of the long-run equilibrium paths of inflation and output. However, it enters the real wage equation with the correct sign but not significantly. We are aware that measuring credit risk premium on bank loans at high frequencies is a difficult task, and that we have created a highly model-specific proxy measure; hence further empirical analysis on this variable may be necessary.

We now turn to short-term dynamics. The short-run dynamic adjustment structure relative to the formally and empirically identified cointegrating relations is reported in Table 4.

Table 4. Short-run adjustment coefficients (bold coefficients denote significance at 5%)

	π_{t+12}	w_t	q_{t+12}	k_t	ρ_t
	The adjustment coefficient matrix $\hat{\boldsymbol{\alpha}}$ (transposed)				
$\hat{\boldsymbol{\alpha}}_1'$	-0.028	0.082	-0.001	**-0.291**	0.000
$\hat{\boldsymbol{\alpha}}_2'$	**-0.014**	**-0.102**	0.052	**0.033**	0.000
$\hat{\boldsymbol{\alpha}}_3'$	**0.028**	0.067	**-0.403**	**-0.130**	0.000

First of all, the negative sign of the three diagonal coefficients of the cointegrating variables π_{t+12}, w_t, and q_{t+12} , ensures they their short-run dynamics is not explosive. Likewise, it is interesting to note that the first and third relationship, which concern inflation and output respectively, display significant disequilibrium error correction in the inter-bank rate, with quite a high speed of adjustment in case of inflation, suggesting that this variable also ensures non-explosive dynamics, and may contain some information on policy conduct.

3.4. Is the inter-bank rate a control variable in the system?

Overall, the statistical picture is one where changes in the inter-bank rate trigger transitory dynamics as maintained by current conventional wisdom, but the key finding is that transitory dynamics occurs around shifting long-run equilibrium paths of output and inflation. This finding suggests that as long as the central bank can control the inter-bank rate, it can also gain control over long-run inflation and output. In other words, shifts in the values of k_t in system (3.3) trace out a non-vertical long-run Phillips curve in the output-inflation space. Here we wish to put forward a more rigorous statistical analysis of the idea that the inter-bank rate is a means to control the long-run equilibrium path of output and inflation.

As Johansen and Juselius (2003) show, a target variable is controllable if it can be made stationary around a desired target value by using an instrument. This methodology is detailed in Appendix A1. To this end, it is necessary to move to the VMA representation of the DGP of system and, in particular, to the matrix $\mathbf{C} = \boldsymbol{\beta}_{\perp}(\boldsymbol{\alpha}'_{\perp}\boldsymbol{\Gamma}\boldsymbol{\beta}_{\perp})^{-1}\boldsymbol{\alpha}'_{\perp}$, which plays an important role in understanding the I(1) models. This control theory in terms of instrument variables and target variables requires that the long-run impact of a shock to the instrument variable must be significant on the target variable. If we consider the inter-bank rate as the instrument and inflation and output as the targets, we can see from the corresponding rows of the estimated matrix $\hat{\mathbf{C}}$ reported in Table 5 that the target variables can in fact be controlled by the inter-bank rate.

Table 5. The long-run impact on inflation and output of unanticipated shocks to the system (*t-values* in parentheses)

	$\varepsilon_{\pi_{t+12}}$	ε_{w_t}	$\varepsilon_{q_{t+12}}$	ε_{k_t}	ε_{ρ_t}
π_{t+12}	0.578	-0.079	0.052	-0.078	-0.689
	(4.279)	(-1.532)	(2.020)	(-1.681)	(-1.238)
q_{t+12}	1.279	-0.174	0.114	-0.172	-1.238
	(4.341)	(-1.554)	(2.049)	(-1.706)	(-1.019)

The interesting information we get is that the long-run impact of unanticipated shocks to the inter-bank rate have a negative, significant at 10%, impact on inflation and on industrial production.

4. Conclusions

Drawing on the modern literature on the monetary transmission mechanisms and capital market imperfections, in this paper we have put forward a theoretical and empirical analysis of the CCC of monetary policy. This channel combines bank credit supply, as a means whereby monetary policy affects economic activity, and interest rates on loans as a cost to firms. The thrust of the model is that firms' reliance on bank loans make *aggregate supply* dependent on credit variables, namely the official rate controlled by the central bank and a credit risk premium charged by banks on firms. In this way, shocks in either variable produce a pattern of macroeconomic relationships that fit and explain the observed empirical regularities in major industrial countries, but are also candidates to exert *permanent* effects on output and inflation, with no recourse to additional non-competitive hypotheses.

We have presented a statistical analysis of this model applied to Italian data from 1986:1 to 1998:12. The statistical methodology adopted enabled us to treat in a single integrated framework both the identification of structural relationships among the variable of interest – i.e. the determinants of the long-run stochastic equilibrium path of these variables – and their deviations from these paths. Though Italy may be regarded as a paradigmatic case, we think that our main conclusions may be of general interest.

First, statistics support the hypothesis that the inter-bank rate, by way of the CCC transmission mechanism, is a co-determinant, with negative sign, of the long-run stochastic equilibrium paths of the real wage rate, output and inflation around which transitory dynamics takes place. Support for the same hypothesis for the credit risk premium is weaker since its connection with the real wage rate is not significant. This may be due to the lack of a better measure of risk, which is notoriously problematic in applied macro-models with high frequency observations. Second, by exploiting the properties of Johansen's theory of control, we have also provided a statistical test and measure of the proposition that the inter-bank rate is a control variable for output and inflation. We have found that the inter-bank rate qualifies as a control variable of the system since its shocks exert a significant, though at 90%, long-run impact on inflation and ouput, implying controllability of the latter variables.

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Figure 1. Work, production and consumption in the sequence economy

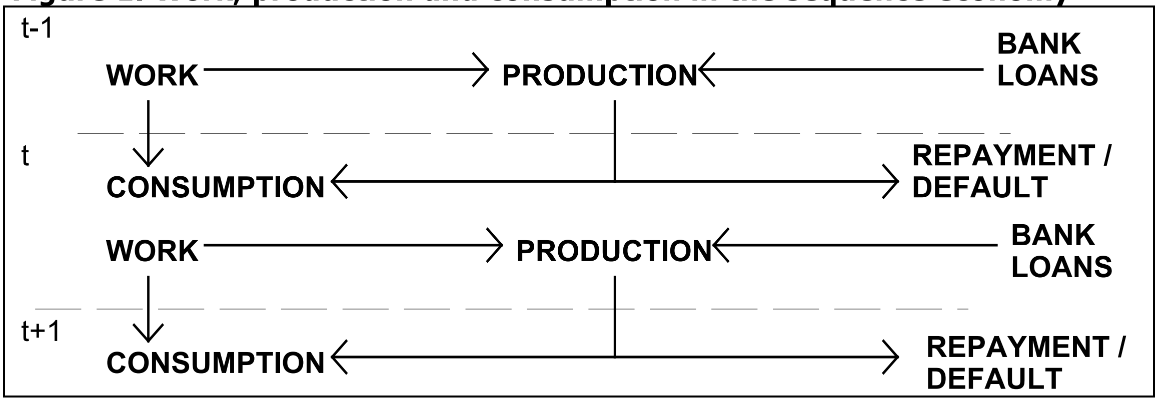


Figure 2. An increase in the official rate

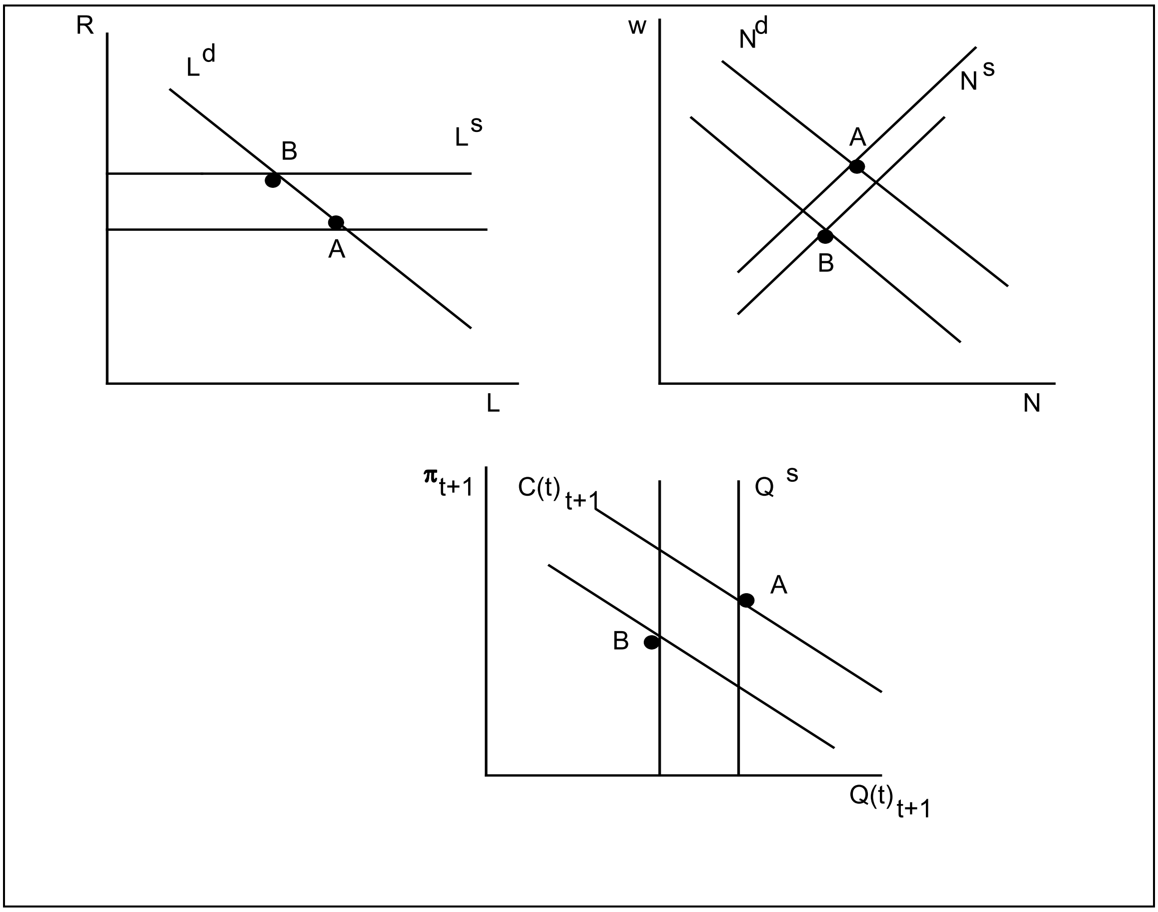
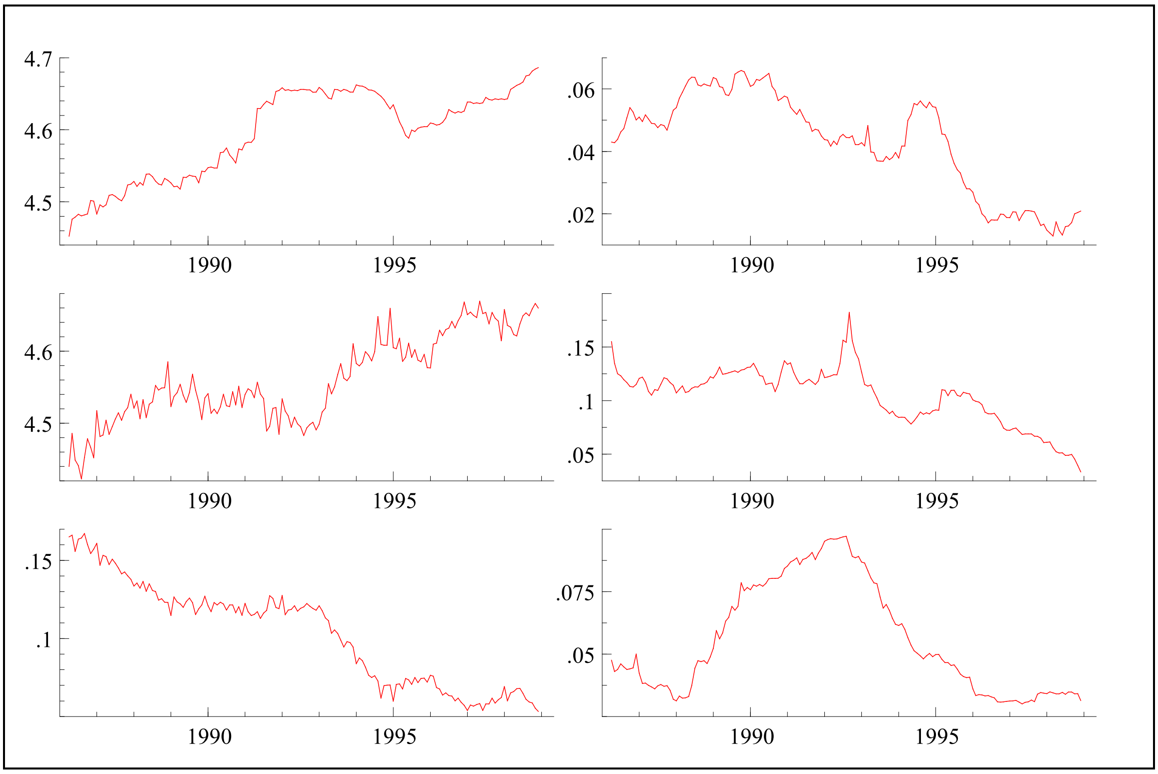


Figure 3. Plots of variables (left to right): index of industrial real wages[1]; inflation rate[1]; index of industrial production (12 months ahead)[1]; inter-bank rate[1]; credit risk premium; German inter-bank rate[1]



Sources: [1]IMF, *International Financial Statistics*; [2]Bank of Italy, *Monetary Statistics*

Appendix

A1. Methodology

The VECM specification of a p-variables VAR(n) model is the following

(A.1) $$\Delta\mathbf{y}_t = \Gamma_0\Delta\mathbf{z}_t + \sum_{i=1}^{n-1}\Gamma_i\Delta\mathbf{x}_{t-i} + \Pi\mathbf{x}_{t-1} + \mu_0 + \mu_1 t + \Phi\mathbf{D}_t + \varepsilon_t$$

where $\mathbf{z}_t$ is a vector of exogenous variables, $\mathbf{x}'_t$ =[$\mathbf{y}'_t$, $\mathbf{z}'_t$], μ_0 is a constant term, t is a linear trend, $\mathbf{D}_t$ a vector of seasonal and intervention dummies, $\boldsymbol{\varepsilon}_t$ a vector of normal disturbances, and $\boldsymbol{\Gamma}$, $\boldsymbol{\Pi}$, $\boldsymbol{\Phi}$ are matrices of coefficients. Let us suppose for semplicity that there are no exogenous variables. For the existence of cointegrating relations, the matrix $\boldsymbol{\Pi}$ must be of reduced rank r, where r is an integer such that $0 < r < p$. This implies that it is possible to factorise $\boldsymbol{\Pi}$ into $\Pi = \alpha\beta'$, where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ can both be reduced to (pxr) matrices. In this case we say that there are r cointegration vectors - that is, r columns of $\boldsymbol{\beta}$ form r linearly independent combinations of the variables in $\mathbf{x}_t$, each of which is stationary - and (p – r) non-stationary vectors, the common trends. The space spanned by the matrix $\boldsymbol{\alpha}$ is the adjustment space and its elements are the adjustment coefficients.

To test the rank of matrix $\boldsymbol{\Pi}$, that there are at most r cointegration vectors, we use what has become known as the *trace* statistic:

(A.2) $$\lambda_{trace} = -2\log(Q) = -T \sum_{i=r+1}^{p} \log(1-\hat{\lambda}_i), \quad r = 0,1,2,\ldots,(p-1),$$

where $\hat{\lambda}_i$ are eigenvalues of a properly defined and estimated matrix (Johansen (1995) p. 93).

Once the matrix $\boldsymbol{\beta}$ has been estimated, it is possible to test linear restrictions on its parameters. Hypotheses about the long-run coefficients $\boldsymbol{\beta}$ can be formulated as follows:

(A.3) $$H_\beta : \beta = (\mathbf{H}_1\varphi_1, \mathbf{H}_2\varphi_2, \ldots, \mathbf{H}_r\varphi_r)$$

where the matrices $\mathbf{H}_1, \mathbf{H}_2, \ldots, \mathbf{H}_r$, expressing the linear economic hypotheses to be tested on each of the r cointegrating vectors, are (pxn_i) matrices and each φ_i is an (n_ix1) vector of parameters to be estimated in the i-th cointegrating vector. The test function is a Likelihood Ratio test which is asymptotically χ^2 distributed with a number of degrees of freedom equal to the number of overidentifying restrictions on the parameters. It is important to check that the restrictions actually identify the system and this can be done by the rank condition.

Hypotheses on the adjustment coefficients $\boldsymbol{\alpha}$, of the form whether the coefficients are zero for a certain subset of equations, can be tested by a Likelihood Ratio test. This means that the variables in the subset are weakly exogenous for the long-run parameters and the remaining adjustment parameters. As Johansen (1995, p. 78) notes: "...zeros in a row of $\boldsymbol{\alpha}$ means that the disturbances from this equation cumulate to a common trend, since $\alpha_{\perp}$ contains the corresponding unit vector...".

If the matrix $\alpha'_{\perp}\Gamma\beta_{\perp}$ is of full rank – where $\alpha_{\perp}$ and $\beta_{\perp}$ are $(p\times(p-r))$ matrices orthogonal to $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$, respectively, and $\Gamma = \mathbf{I} - \sum_{i=1}^{n-1}\Gamma_i$ – then the vector $\mathbf{x}_t$ is I(1) and its VMA representation is the following:

$$\text{(A.2)} \qquad \mathbf{x}_t = \mathbf{C}\sum_{i=1}^{t}(\varepsilon_i + \Phi\mathbf{D}_i) + \mathbf{C}\mu\, t + \mathbf{C}(L)(\varepsilon_t + \mu_t + \Phi\mathbf{D}_t) + \mathbf{B}_0 \ ,$$

where the matrix $\mathbf{C} = \boldsymbol{\beta}_{\perp}(\boldsymbol{\alpha}\boldsymbol{\alpha}_{\perp}\boldsymbol{\Gamma}\boldsymbol{\beta}_{\perp})^{-1}\boldsymbol{\alpha}'_{\perp}$, $\mathbf{C}(L)$ is an infinite-order polynomial in the lag operator L, and $\mathbf{B}_0$ is a function of the initial values (Johansen (1995) p. 49). The matrix $\mathbf{C}$ plays an important for the understanding of the I(1) models. As Johansen (1995, pp. 49-50) explains: "...One can interpret matrix $\mathbf{C}$ as indicating how the common trends $\boldsymbol{\alpha}'_{\perp}\sum_{i=1}^{t}\varepsilon_i$ contribute to the various variables through the matrix $\boldsymbol{\beta}_{\perp}$. Another interpretation is that a random shock to the first equation, say, at time t=1, gives rise to short-run random effects as represented by $\mathbf{C}(L)\varepsilon_t$ which die out over time, and a long-run effect to the stochastic part given by the first column of $\mathbf{C}$. This is orthogonal to $\boldsymbol{\beta}$ such that the new position is a new point on the attractor..."

The estimated matrix $\hat{\boldsymbol{\alpha}}_{\perp}$ plays an important role because it determines the stochastic common trends as given by $\hat{\boldsymbol{\alpha}}'_{\perp}\sum_{i=1}^{t}\hat{\varepsilon}_i$, where $\hat{\varepsilon}_i$ is the vector of estimated errors of the VECM, and the matrix $\tilde{\boldsymbol{\beta}}_{\perp} = \hat{\boldsymbol{\beta}}_{\perp}(\hat{\boldsymbol{\alpha}}'_{\perp}\hat{\boldsymbol{\Gamma}}\hat{\boldsymbol{\beta}}_{\perp})^{-1}$ determines their loadings.

A2. A theory-based proxy for the credit risk premium

According to our model for firms, default of a firm j occurs whenever

$$\text{(A.3)} \qquad P_{t+1}Q(t)_{jt+1} < L^{d}_{jt}R_t$$

or, $L^{d}_{jt}R_t \big/ Q(t)_{jt+1} > P_{t+1}$,

where the variables have already been defined. Since we assume that all firms operate under the same technological and market conditions the default probability will be the same for all firms. In order to measure it, we use a procedure where we consider default risk any deviation from a stable "solvency relation" given by (A.3) taken as an equality for the economy as a whole. That is, we assume that in the long-run a stable linear relation should hold between the non-stationary variables L_tR_t and $P_{t+12}Q_{t+12}$, once they both have been log-transformed, i.e.:

(A.4) $$\log(L_tR_t) = \mu_0 + \beta_1 \log(P_{t+12}Q_{t+12}) + u_t .$$

According to cointegration analysis, the quantities u_t are to be considered as disequilibrium errors controlled by the short run dynamics of the variables in the model. Whenever these errors tend to depart from the common trend of the model, they are pulled towards it by some adjustment coefficients estimated together with the long-run coefficients β (Johansen (1995) p.41).

The empirical cointegration analysis is performed on the vector y_t = [log(L_tR_t), log($P_{t+12}Q_{t+12}$]. The chosen specification shows k = 4 lags and a constant restricted to the cointegration space, that is a constant is allowed in the cointegrating relation. The test results show that the variables cointegrate and thus we have r = 1 cointegrating relation and (p – r) = 1 common trend. The estimated disequilibrium errors have been considered as a time serie measure of the default risk variable ϕ in the theoretical model, and have then been translated into the credit risk premium times series according to relationship (2.14) in the text. As can be seen from figure 3 in the text, the risk premium appears to be characterized by a declining trend in the first period, followed by a period in which it tends to invert this trend and then it declines again, until reaching a stabilization period starting from 1995. Notably, a spike of riskiness occurs around the 1991-92 crisis of the lira in the EMS with a sharp increase in interest rates and in the expected depreciation rate of the lira. This path is consistent with other different measures of default risk on bak loans in Italy: see e.g. Morelli and Pittaluga (1997), Fiorentini and Tamborini (2002)..

A3. Mis-specification tests of the unresctrited cointegrated model

The results of specification tests for the unrestricted VAR(3) model with dummies in paragraph 3.2 take the following values: the *LM*(1) test for first order autocorrelation, asymptotically distributed as a χ^2_{25} variable, is equal to 39.42 with a *p-value* of 0.03; the *LM*(4) test for fourth order autocorrelation, asymptotically distributed as a χ^2_{25} variable, is equal to 27.14 with a *p-value* of 0.35. Given that conditioning on weakly exogenous variables often improves the statistical specification of the model, we have estimated a conditional model in which risk has been assumed weakly exogenous, as resulted from a previous test for weak exogeneity. The output of the residual analysis for the conditional model was the following: the *LM*(1) test, distributed as a χ^2_{16} variable, equal to 14.08 with a *p-value* of 0.59 and the *LM*(4) test, distributed as a χ^2_{16} variable, equal to 23.94 with a *p-value* of 0.09. Therefore, in the conditional model, we cannot reject the hypothesis of no autocorrelation. As concerns residual normality, this is rejected due to excess kurtosis in real wages and in the inter-bank rate. Because VAR estimates are more sensitive to deviations from normality due to skewness than to excess kurtosis, we consider the model chosen as a well specified one.

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