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MONETARY POLICY THROUGH THE "CREDIT-COST CHANNEL". ITALY AND GERMANY

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Monetary policy through the "credit-cost channel". Italy and Germany*

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Abstract

In this paper we wish to extend the empirical content of the "credit-cost channel" of monetary policy that we proposed in Passamani and Tamborini (2005). In the first place, we replicate the econometric estimation of the model for Italy, to which we add Germany. We find confirmation that, in both countries, firms' reliance on bank loans ("credit channel") makes *aggregate supply* sensitive to bank interest rates ("cost channel"), which are in turn driven by the inter-bank rate controlled by the central bank plus a credit risk premium charged by banks on firms. The second extension consists of a formal econometric analysis of the idea that the interest rate is an instrument of control for the central bank. The empirical results of the CCC model that, according to Johansen and Juselius (2003), innovations in the inter-bank rate qualify this variables as a "control variable" in the system. Hence we replicate the Johansen and Juselius technique of simulation of rule-based stabilization policy. This is done for both Italy and Germany, on the basis of the respective estimated CCC models, taking the inter-bak rate as the instrument and the inflation of 2% as the target. As a result, we find confirmation that inflation-targeting by way of inter-bank rate control, grafted onto the estimated CCC model, would stabilize inflation through structural shifts of the "AS curve", that is, the path of realizations in the output-inflation space.

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

JEL codes: E51, C32

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MONETARY POLICY THROUGH THE "CREDIT-COST CHANNEL". ITALY AND GERMANY

1. Introduction

The analysis of the channels through which monetary policy operates affecting real macroeconomic fluctuations has long been and is still matter of research in the economic literature. More recent lines of inquiry have moved in two directions. One is the "credit channel", which refers to the means by which monetary policy affects aggregate demand via banks and credit institutions (Bernanke and Gertler (1990), Gertler and Gilchrist (1993, 1994), Trautwein (2000)). The other direction is the "cost channel", which investigates monetary effects on aggregate supply via firms' production costs, including the financial costs (Barth and Ramey (2001), Ravenna and Walsh (2003)).

In Passamani and Tamborini (2005) (PT henceforth), we proposed to blend these two channels into a single one, the "credit-cost channel" (CCC). The CCC may provide a consistent framework for monetary policy analysis for three main reasons. In the first place, the credit-channel literature offers an explanation for the interest rate to be treated as a production extra cost: due to capital market imperfections (mostly asymmetric information between lenders and borrowers) firms are forced to resort to external sources and pay a premium on them. Secondly, it also explains why credit represents the single source of external funds for some classes of firms and hence plays a "special role" in the production process. Finally, the transmission from policy rates to bank rates is tight and well documented, whereas the transmission to open-market long-term rates is notoriously problematic.

Our CCC model consists of three competitive markets – labour, credit, and output – and three classes of agents – households, firms and banks – with a central bank. Firms are bank-dependent for working capital and face a real unit cost of production given by the current real wage rate plus the (expected) real interest rate. Monetary policy affects economic activity as policy-induced changes in the bank interest rate exert a credit-cost effect on firms that shifts labour demand, output supply and demand, given rational expectations of future inflation. A major implication of the CCC is that the supply-side impact of monetary shocks is amplified. The joint consideration of the credit and cost channels may overcome the weaknesses of the two separate approaches yielding 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, namely:

- monetary policy impulses have persistent real effects
- policy interventions are followed by delayed adjustment of prices
- real wages are also procyclical with output after a monetary shock.

The evidence presented in PT by means of a structural cointegration analysis (Johansen (1996)) of the CCC model for Italy (1986:1-1998:12) supports the view that firms' reliance on bank loans makes *aggregate supply* sensitive to bank interest rates, which are in turn driven by the inter-bank rate controlled by the central bank plus a credit risk premium charged by banks on firms. Moreover, the structural cointegration technique allowed to point out that changes in the inter-bank rate trigger transitory dynamics as maintained by current conventional wisdom, but, as a result of the supply-side effect, transitory dynamics occurs around *shifting long-run equilibrium paths of output and inflation*.

In this paper we wish to extend the empirical content of the CCC approach. In the first place, we replicate the econometric estimation of the model for Italy, to which we add Germany. Germany qualifies as a natural case study for the same reasons as Italy. First, because bank credit is an important element in the transmission mechanism of monetary policy in both countries. As shown in Gambacorta (2001, pp. 12-15), in both countries the business sector was heavily dependent on bank credit in the years between 1986 and 1998, the period chosen for the analysis. In fact, in this period, figures on the composition of financial liabilities of Italian firms were similar to German ones, with a relatively low stock market capitalization. Moreover, the weight of bank credit with respect to total credit was 85% for Italy and 84% for Germany, the share of loans backed by collateral and the availabilitity of non bank finance were very similar. Yet whereas both countries have been object of numerous investigations detecting the traditional demand effects of the credit channel (see Fiorentini and Tamborini (2001) for a survey), to our knowledge investigations of the supply-side effects are very limited[1].

On the other hand, Germany and Italy historically differed as far as the conduct of pre-EMU monetary policy is concerned. The role of the credit market in the monetary transmission mechanism has always been carefully monitored by the Italian monetary authorities and was explicitly included in the Bank of Italy's (BoI) econometric model (1997a). Direct controls over credit supply were also explicitly considered among the BoI's policy instruments. By contrast, the Bundesbank (BB) officially endorsed the money-quantity approach based in the monetarist tradition. However, the official policy style only concealed the importance that the BB attached to the role of credit and bank rates in the transmission mechanism. What is more important, the time periods we have chosen for the two countries (see below) saw a substantial homogeneization of the policy framework in the two countries under the pressure of the exchange-rate constraints of the European Monetary System (EMS) first, and of the convergence towards the single currency later (Angeloni (1994), Visco (1995)). In the second half of the '1980s, the BoI abandoned the pervasive and recurrent administrative interventions that characterized the previous decades. In the 1990s, all major European central banks moved towards a more or less explicit practice of interest-rate control, the well-known "corridor of rates", that was eventually adopted by the European Central Bank (European Monetary Institute (1997)).

The model for Italy is re-estimated over the same time period (1986:1-1998:12) and with a monthly data set including the same variables (the real wage rate, the inter-bank interest rate, industrial production and inflation) except for the proxy of credit risk, which is now given by the spread between the bank lending rate and the medium-term government bond yield (a proxy largely employed in the relevant literature: see Fiorentini and Tamborini

[1] As regards Italy, see Fiorentini and Tamborini (2002). Gaiotti and Secchi (2004) 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.

(2001))[2]. The model for Germany is estimated over a shorter time period (1990:1-1998:12) in order to bypass the reunification shock, and with the same data set as Italy. For both countries the results are consistent with the previous findings in PT.

The second extension of our empirical study consists of a formal econometric analysis of the idea that the interest rate is an instrument of control for the central bank. The above mentioned empirical results found by means of the Johansen technique indicate that innovations in the inter-bank rate shift the stochastic paths of output and inflation. According to Johansen and Juselius (2003) this result qualifies this variables as a "control variable" in the system: a control variable is such that its innovations have a significant long-run impact on the associated target variable of the system, and make it stationary around the desired target. In this paper, we apply the Johansen and Juselius technique of simulation of rule-based stabilization policy whereby the control variable is aimed to the associated target variable. This is done for both Italy and Germany, on the basis of the respective estimated CCC models, taking the inter-bak rate as the instrument and the inflation target of 2% as the target. As a result, we find confirmation that inflation-targeting by way of inter-bank rate control, grafted onto the estimated CCC model, would stabilize inflation through structural shifts of the "AS curve", that is, the path of realizations in the output-inflation space. The simulation can then be interpreted in two ways. As a "counterfactual" exercise, it shows how the history of output and inflation would have differed in the two countries if the two central banks had followed the 2% rule in the past. As a "predictive" *ceteris-paribus* exercise, it shows how output and inflation react in the two countries in response to the common policy rule of 2%.

2. The structural relations identifying the "credit-cost channel"

In light of the relevant literature, the key features of the CCC model are that (Greenwald and Stiglitz (1988, 1993):

- production takes time, typically 1 time period: t, t+1

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[2] In PT (2005) this proxy was given by an independent estimation of the deviations from a cointegrated relationship between output and firms' outstanding bank debt representing their solvency condition.

- firms need external funds in advance to finance working capital
- these funds come from bank credit.

The important element in the cost-channel models of monetary policy is the labour demand function of the bank-dependent firm. When firms plan production at any time *t*, they are uncertain about their future revenue from output sales. The sale price of a firm *j* at time *t*+1 is a random draw from the probability distribution with density $f(\tilde{P}_{jt+1})$, cumulative function *F*, and expected value $E(\tilde{P}_{jt+1}) = P_{t+1}$ for all *j*. Assuming that all firms use the same production technology ($Q(\bullet)$) and face the same current real wage rate ($w_t \equiv W_t/P_t$) and nominal bank interest rate (r_t), in each period *t* along the otpimal production path they will employ labour ($N^d{}_t$) up to the point where the marginal product equals the *expected real unit cost*, which is the compound real cost of labour and credit (Greenwald and Stiglitz (1988, 1993), Christiano and Eichenbaum (1992), Christiano et al. (1997)). In addition to the usual negative relationship with the real wage rate, the main features of labour demand/output supply due to bank debt are that: 1) they are systematically lower than the values for the unconstrained firm for any positive interest rate, 2) they are decreasing in the real interest rate.

As is typical of this class of models, three activities of households are considered: labour supply, output demand and saving (Christiano et al. (1997)). The labour supply function ($N^s{}_t$) displays the usual properties once account is taken of the fact that *current* working time is the means to buy *future* consumption, so that the expected rate of inflation ($\pi^e{}_{t+1}$) affects the working time distribution over time. In an economy where there is no direct lending to firms, the consumption demand function shows that at the end of each period this can be equal to, or less than, the real value of deposits (D_t/P_t) and the result is a simple demand function determined by real money balances. Yet these are endogenous with the amount of loans extended to firms ($L_t = D_t$).

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. 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 via the official rate. In view of the fact that households will claim on their deposits one period later, the bank should secure itself a sufficient amount of liquid resources. As a result, the interest rate on loans charged by the bank is (approximately) given by $r_t \approx \rho_t + k_t$, where ρ_t is a measure of credit risk. Hence, r_t can be interpreted as the sum of the official rate plus a credit risk premium providing the link between monetary policy and aggregate supply.

We can summarize the complete macroeconomic equilibrium as follows:

Labour market

(2.1) $N^d(w_t, r_t, \pi^e_{t+1}) = N^s(w_t, \pi^e_{t+1})$

Credit market

(2.2) $L_t = W_t N_t$

(2.3) $D_t = L_t$

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

Output market

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

(2.6) $\pi^e_{t+1} = \mathrm{E}_t(\widetilde{P}_{jt+1}/P_t - 1) = P_{t+1}/P_t - 1$

The thrust of our model is that variations of k_t (and/or ρ_t) can, under certain conditions, generate a pattern of relationships which is consistent with the empirical regularities observed in major industrialized countries, i.e. $\mathrm{d}w_t/\mathrm{d}k_t < 0$, $\mathrm{d}Q(t)_{t+\mathrm{s}}/\mathrm{d}k_t < 0$, $\mathrm{d}\pi_{t+\mathrm{s}}/\mathrm{d}k_t < 0$ to the exclusion of ancillary hypotheses like monopolistic competition or price stickiness[3].

The CCC transmission mechanism, to which we shall refer as the null hypothesis, hinges on the signs of the variables k_t and ρ_t in the equations for $\pi_{t+\mathrm{s}}$, w_t, and $q_{t+\mathrm{s}}$, and it implies the unique pattern of signs of coefficients in Table 1.

[3] The above-mentioned conditions are, in terms of first derivatives, $N^s_\pi < N^s_{\underline{w}}$ and $Q_N < 1$. 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). As is well known, the relative magnitude of N^s_π and $N^s_{\underline{w}}$ 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 (N^s_π) (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 CCC transimission mechanism identified by our model does yield the observable correlations *thanks to a relatively small intertemporal substitution effect.*

Table 1. The pattern of coefficient signs for the CCC hypothesis

H_0	k_t	ρ_t
π_{t+12}	$\beta_{13} < 0$	$\beta_{14} < 0$
w_t	$\beta_{23} < 0$	$\beta_{24} < 0$
q_{t+12}	$\beta_{33} < 0$	$\beta_{34} < 0$

3. Identification and estimation of the structural relationships

In this section we present the results of the econometric analysis of the CCC model presented above applied to both Italy (1986:1-1998:12) and Germany (1990:1-1998:12). In what follows we summarize the main steps and results. Details on statistical procedures and tests are gathered in a separate Statistical Appendix.

3.1.Data and methodology

According to the CCC model, the variables of interest for both countries 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 used the inter-bank rate[4];
- the credit risk premium ρ_t, not observable, for which we have adopted, as a proxy, an appropriately defined log transformation of the spread between the bank lending rate and the medium-term government bond yield for Italy[5] and of the spread between the short-term bank lending rate and the money market rate for Germany;
- output Q_{t+s}, given by the industrial production index;
- the inflation rate, π_{t+s}, measured by the consumer price index.

Our theoretical focus is on the inter-bank rate k_t. Thus, instead of adding the rate on bank loans as an independent variable, we have considered directly its two components k_t and ρ_t. The latter variable mainly

[4] Since the inter-bank rate is highly sensitive to central bank interventions, it is taken to be 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).

[5] The same spread has been used for Italy by Chiades and Gambacorta (2004).

allows for control for autonomous changes in credit conditions. Hence ρ_t, a problematic variable to measure, is not crucial for the significance of the model.

In consideration of the fact that Italy, in the relevant time period, faced constraints on domestic monetary policy and interest rates due to strong exchange-rate targeting and high capital mobility in the EMS, we have also added

- the German inter-bank rate, k^*_t.[6]

As to Germany, though she was generally regarded as the uncontrained country in the EMS country, we have also added an exogenous foreign variable to control for world monetary conditions, namely

- the three months LIBOR in US dollar, *Lib*.

The time lead s to be applied to Q and π should capture the theoretical gestation time of output and the related time-horizon for expected inflation[7]. A time lead of 12 months has been chosen empirically by means of sensitivity tests.

All variables, except interest rates, are log-transformed and are observed through monthly times series, plotted with sources in Figure 1 for Italy and in Figure 2 for Germany.

Since our aim is to test a fully specified system of structural relationships, for the reasons put forward in the Introduction we have chosen the structural cointegration approach developed by Johansen (1996)

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[6] The literature on monetary policy in the EMS (see e.g. De Grauwe (1992)) would predict that in a country like Italy the domestic interest rate could not deviate systematically from uncovered parity with Germany, 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. On this view, a monetary policy shock can 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. We consequently introduced the two inter-bank rates as two independent variables with *opposite* expected sign, and we 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 improved the overall quality of the estimates.

[7] Since in PT we assumed flexible prices and rational expectations, we could take the actual inflation rate on the same time lead as output as a proxy for the theoretical expected inflation.

and Johansen and Juselius (2003), both for the estimation and identification of the long-run structural relationships among the theoretically relevant variables, and for the evaluation of the policy variable as a control variable of the system[8].

3.2. Econometric results

In the first place, for the p-dimensional (p = 5) observed process $\mathbf{y'}_t$ = [π_{t+12}, w_t, q_{t+12}, k_t, ρ_t,] we have assumed an unrestricted vector autoregressive model written in error correction form (VECM). The model has been augmented, in both countries, to include an exogenous variable and deterministic terms. The resulting equation is the following:

$$\Delta\mathbf{y}_t = \boldsymbol{\Gamma}_0\Delta\mathbf{z}_t + \sum_{i=1}^{n-1}\boldsymbol{\Gamma}_i\Delta\mathbf{x}_{t-i} + \boldsymbol{\Pi}\mathbf{x}_{t-1} + \boldsymbol{\mu}_0 + \boldsymbol{\mu}_1 t + \boldsymbol{\Phi}\mathbf{D}_t + \boldsymbol{\varepsilon}_t \ , \tag{3.1}$$

where $\mathbf{z'}_t$ = [k^*_t], $\mathbf{x'}_t$ =[$\mathbf{y'}_t$, $\mathbf{z'}_t$], $\boldsymbol{\varepsilon}_t$ is a vector of normal disturbances, Δ is the first difference operator and $\boldsymbol{\Gamma}$, $\boldsymbol{\Pi}$, $\boldsymbol{\Phi}$ are matrices of coefficients. The deterministic terms include a vector of constants $\boldsymbol{\mu}_0$, a linear trend t restricted in the analysis to the cointegration space[9], a vector of intervention dummies $\mathbf{D}_t$[10]. The number n of lags in (3.1), n = 3 for Italy and n = 2 for Germany, was determined on the basis of misspecification tests[11].

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[8] The entire empirical analysis was performed using the CATS software which needs the RATS package to be run. The results are available upon request.

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

[10] In order to obtain residuals close to Normality, in the Italian data-set we introduced five permanent intervention dummies and two transitory intervention dummies into our data set to account for the exit of the Italian Lira from the EMS in 1992 and for few other events. The permanent intervention dummies were defined for 1991/I, 1991/V, 1992/VII, 1994/IV and 1995/III. The transitory intervention dummies were defined for shocks of opposite signs in 1992/IX-1992/X, and in 1993/III-1993/IV. In the German data-set we introduced two permanent intervention dummies for 1991/X and 1996/I.

[11]For the Italian data the results of specification tests for the unrestricted VAR(3) model with dummies take the following values: the $LM(1)$ test for first order autocorrelation, asymptotically distributed as a χ^2_{25} variable, is equal to 26.186 with a *p-value* of 0.399; as concerns residual Normality, the test asymptotically distributed as a χ^2_{10} variable, is equal to 19.217, with a *p-value* of 0.038.

For the German data the results of specification tests for the unrestricted VAR(2) model with dummies take the following values: the $LM(1)$ test for first order autocorrelation is equal to 33.576 with a *p-value* of 0.117; the test for residual

Italy

We have then sought for cointegrating relations, first of all for Italy. The procedure has followed closely the previous one in PT (2005) to which we refer the reader for greater details. Here we summarize the main steps:

- by standard procedure, tests indicated 3 cointegrating vectors, and hence 2 nonstationary relations[12]
- the unrestricted relations have been normalized with respect to the 3 variables that the theory indicates as "endogenous" (π_{t+12}, w_t, q_{t+12}) *vis-à-vis* the CCC “explanatory” variables (k_t, ρ_t, k^*_t) and the trend
- identification[13] (two zero restrictions and one normalization on each cointegrating relation in order to satisfy the rank and order condition) has been accomplished by exploting the forward-looking sequential structure of the theoretical model; consequently, the unrestricted system (2.1), extended to include the variable k^*_t in the relations, could be restricted to a quasi-reduced “pyramid” form simply by setting $\beta_{11} = \beta_{12} = \beta_{22} = 0$[14].

Below we report the final just-identified long-run relations (*t*-statistics in parentheses; bold coefficients denote significance at 10%) together with the value of the *LR* test[15]:

Normality is equal to 26.824, with a *p-value* of 0.003. As concerns residual Normality for Germany, this is rejected due to excess kurtosis in real wages. 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.

[12] See the Statistical Appendix A1

[13] As shown in Chapter 10 of Juselius’ “Notes” for “Advanced Econometrics”, www.econ.ku.dk/okokj/, the long-run structure can be identified in the so called reduced form (3.1) of the cointegrated VAR model, so that we can test structural hypotheses on the long-run structure $\boldsymbol{\beta}$ without having jointly to identify the short-run structure.

[14] . In addition, in the third relation we set the coefficient of k^*_t equal to zero, as it did not show up as significant, though correctly signed, in any preliminary analysis.

[15] The degreees of freedom of the *LR* test corresponds to the weak exogeneity restrictions for the variable k_t, which support the finding that the interbank rate can be considered as an instrument policy variable.

$$\text{(3.2)}\quad \begin{aligned} \pi_{t+12} &= \mathbf{-3.412}k_t + \mathbf{2.066}\rho_t + 1.313k^*_t \mathbf{-0.002}t + \hat{u}_{\pi_{t+12}} \\ &\qquad (-3.130) \quad (3.068) \quad (1.495) \quad (-4.808) \\ w_t &= \mathbf{-2.908}\pi_{t+12} - \mathbf{1.045}k_t - \mathbf{0.071}\rho_t + \mathbf{2.043}k^*_t + \hat{u}_{w_t} \\ &\quad (-6.769) \quad (-3.739) \quad (-1.872) \quad (7.884) \\ q_{t+12} &= \mathbf{0.705}\pi_{t+12} - \mathbf{0.655}w_t - \mathbf{0.417}k_t - \mathbf{0.417}\rho_t + \mathbf{0.002}t + \hat{u}_{q_{t+12}} \\ &\qquad (1.677) \quad (-6.265) \quad (-2.673) \quad (-2.673) \quad (8.973) \\ &\qquad \chi^2_3 = 2.318, \quad p\text{-}value = 0.509 \end{aligned}$$

These statistical relationships 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 and with those in PT (2005):

- the inter-bank rate k_t always has the expected signs and significant coefficients on inflation, real wage rate and output
- correction for uncovered interest parity via the German rate k^*_t has also the expected sign (see also fn. 6) (apart from the equation for output where it is constrained to zero), but is significant only in the second relation[16]
- the proxy for the credit risk premium ρ_t also replicates the same results in PT (2005) in spite of the change of measurement: it proves consistent with the null hypothesis in the real wage equation and in the output equation.

As explained previously, the result for the inter-bank rate, in particular that it has a negative effect on the real wage rate, can be considered evidence that this variable operates through the supply side of the economy in a way that cannot be consistently explained by the nominal rigidity or the monopolistic competition hypotheses (see Table 1).

Germany

The search for structural cointegrating relations for Germany has followed the same strategy as for Italy. Instead of the variable k^*_t, we have Lib_t as exogenous variable. Given that Germany had no explicit non-EMS exhange-rate target, we expect that, unlike k^*_t for Italy, Lib_t takes the same

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[16] It should be noted that imposing in the first relation the restriction in order to identify the strict uncovered interest parity, i.e. β_{13} = - β_{15}, would have made significant the coefficient of k^*_t .

sign as the domestic rate (see also fn. 6). The pattern of signs of coefficients reported in Table 1 should still be valid for the endogenous variables k_t and ρ_t . Below we report the final just-identified long-run relations (t-statistics in parentheses; bold coefficients denote significance at 10%) together with the value of the *LR* test[17]:

$$
(3.3)\qquad
\begin{array}{l}
\pi_{t+12} = \mathbf{-0.634}k_t - \mathbf{0.756}\rho_t - 0.\mathbf{360}Lib_t - \mathbf{0.001}t + \hat{u}_{\pi_{t+12}} \\
\qquad\quad (-3.109) \quad (-3.681) \ (-2.821) \qquad (-8.891) \\
w_t = \mathbf{-4.358}\pi_{t+12} - \mathbf{2.888}k_t - \mathbf{1.764}\rho_t - \mathbf{2.973}Lib_t + \hat{u}_{w_t} \\
\qquad (-13.563) \qquad (-3.600) \ (-1.931) \ (-5.936) \\
q_{t+12} = \mathbf{5.720}\pi_{t+12} - 0.321w_t - \mathbf{2.845}k_t - \mathbf{2.845}\rho_t + \mathbf{0.004}t + \hat{u}_{q_{t+12}} \\
\qquad\quad (-8.170) \qquad (-1.160) \ (-2.791) \ (-2.791) \ (3.821) \\
\qquad\qquad\qquad \chi^2_3 = 3.361, \quad p\text{-}value = 0.304
\end{array}
$$

The inter-bank rate $k_{t,}$and the risk ρ_t always have the expected, significant coefficients on inflation, real wage rate and output. The coefficient on variable LIBOR is significant and shows the expected sign in each relation.

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. In other words, shifts in the values of k_t should displace the long-run "AS curve" in the output-inflation space.

We have then performed a more rigorous statistical analysis of this transmission mechanism by drawing on Johansen and Juselius (2003). Their methodology hinges on three elements. First, a variable is *controllable* if it can be made stationary around a desired target value by using an instrument variable. Secondly, a necessary condition for a variable to be an *instrument* is that there be a significant long-run impact of a shock to the instrument on the target variable. Thirdly, given controllability, *a control rule* specifies interventions on the instrument conditional on the observed

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[17] See Statistical Appendix A1

state of the target variable relative to the target[18].

To implement this procedure it has been necessary to move from the VAR representation to the VMA representation of the process $\mathbf{y}_t$ and, in particular, to the $(p \times p)$ 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. Its elements convey information about the long-run impact of cumulated shocks to the system variables. Given the reduced rank r of matrix $\mathbf{\Pi}$, the matrix $\mathbf{C}$ is also of a reduced rank (p-r), which corresponds to the number of driving forces or common stochastic trends. In other words, the matrix $\mathbf{C}$ informs on how the endogenous variables react to the nonstationary forces driving the system. Our aim was not to identify these forces, but to understand how the variables react to the cumulated shocks whose combinations give rise to the non stationary driving forces. In this way, it has been possible to evaluate the effects on inflation and output of unexpected changes in the policy action.

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 3 for Italy and in Table 4 for Germany, that the target variables can in fact be controlled by the inter-bank rate in both countries.

Table 3. Italy: the long-run impact on inflation and output of unanticipated shocks to the system (*t-values* in parentheses, bold coefficients denote significance at 5%)

	$\varepsilon_{\pi_{t+12}}$	ε_{w_t}	$\varepsilon_{q_{t+12}}$	ε_{k_t}	ε_{ρ_t}
π_{t+12}	0.301	-0.075	0.012	**-0.443**	0.150
	(1.108)	(-1.235)	(0.433)	(-3.046)	(1.689)
q_{t+12}	0.777	-0.194	0.032	**-1.607**	0.386
	(1.158)	(-1.291)	(0.452)	(-2.976)	(1.765)

[18] See Statistical Appendix A2.

Table 4. Germany: the long-run impact on inflation and output of unanticipated shocks to the system (*t-values* in parentheses, bold coefficients denote significance at 5%)

	$\varepsilon_{\pi_{t+12}}$	ε_{w_t}	$\varepsilon_{q_{t+12}}$	ε_{k_t}	ε_{ρ_t}
π_{t+12}	0.120	0.015	**0.060**	**-0.328**	**-0.220**
	(1.859)	(1.086)	(4.789)	(-2.097)	(-2.519)
q_{t+12}	1.212	0.150	**0.608**	**-3.329**	**-2.232**
	(1.859)	(1.086)	(4.788)	(-2.103)	(-2.519)

The interesting information obtained is that innovations in the inter-bank rate have a negative, significant at 5%, *long-run* impact on inflation and on industrial production in both countries. Johansen and Juselius detect the same result in the case of the United States, though with the anomaly of a *positive* sign. They conjecture that this anomaly may be due to lack of the supply side in their model; our result suggests that their conjecture may be right.

It is now possible to see how a derived control rule for the instrument variable, applied at all points in time, would make a nonstationary target variable become stationary around a desired mean value. Following Johansen and Juselius, we have performed a simulation analysis of the use of the inter-bank rate as an instrument to directly control inflation, and of the consequences of this control rule on the other relevant variable, industrial production[19].

In order to derive the control rule, we have first assumed that monetary policy sets the value of the controlled instrument (*ctr*) as a reaction to the observed value of the target variable with respect to its target value. Then the market reacts, generating a new observed value (*new*). Monetary policy intervenes again on the controlled instrument and then the market reacts again. The ordering of the observed values for the process $\mathbf{y}$ is therefore the following:

$$...\mathbf{y}_{t-1} \to \mathbf{y}_{t-1}^{ctr} \to \mathbf{y}_{t}^{new} \to \mathbf{y}_{t}^{ctr} \to \mathbf{y}_{t+1}^{new} \to \mathbf{y}_{t+1}^{ctr} \to \mathbf{y}_{t+2}^{new} \to \mathbf{y}_{t+2}^{ctr} \to \cdots ,$$

At any time t the control rule applied by the monetary authority has the following form:

$$\mathbf{y}_{t}^{ctr} = \mathbf{y}_{t}^{new} + \mathbf{v}_{t} .$$

19 Hence we have simulated a pure inflation targeting regime, rather than a common Taylor rule where output is also a target.

Given our estimated VECM model, the intervention $\mathbf{v}_t$ is a complicated matrix function that depends on (Johansen and Juselius (2003, p.10)):

- the actual discrepancy between the observed and desired value of the target variable;
- the observed deviation of the process from the steady state value on the attractor set and its short-run adjustment dynamics.

Figure 3 shows, for Italy, the interventions $(k_t^{ctr} - k_t^{new})$ on the interbank rate needed to make the inflation rate stationary around a target mean of 2%, and Figure 4 and 5 report, respectively, the observed and the new inflation rate and the observed and the new output, using the derived control rule at all time points. Our main comments are the following.

First, in Figure 3 it is apparent that had the BoI aimed at stabilizing inflation around 2% throughout the sample period, the interventions needed, once the control process had begun, would have been very small indeed and would have made the interbank rate a bit higher than observed in the early 1980s (high inflation) but a bit lower than observed in the late 1990s (low inflation).

Secondly, Figure 4 shows that, as implied by the estimated parameters of matrix **C**, inflation would have fallen under the control of the inter-bank rate, in the sense that the "controlled" pattern exhibited by Figure 4 would in fact have made the "new" inflation path stationary around the 2% target. In particular note that the speed of convrgence towards the 2% target would have been faster.

Thirdly, the estimated parameters of matrix **C**, however, also imply that inflation-targeting would have not been neutral on the industrial production stochastic path. This is in fact highlighted by Figure 5, where it can be seen that as long as the "controlled" inter-bank rate exceeds the actual one (1980s), industrial production is shifted onto a path lower than the observed one. The reverse occurs when the "controlled" inter-bank rate is lower than the actual one (1990s).

Summary evidence of the previous two findings is provided by Figure 6, which plots the scatter diagrams of observed *vis-à-vis* "new" inflation and output realizations. Controllability of inflation by means of the inter-bank rate is obtained by structurally shifting the (long-run) "AS curve". As implied by targeting the 2% inflation rate, the "new AS" is flattened. The lower side of the diagram reflects the loss of output incurred by aiming at 2% inflation in the 1980s. Given the 2% inflation target, and given that the model embodies rational expectations, the dispersion of realizations of output and inflation around the "new AS" is only due to optimal responses to non-policy random shocks. Taylor's (1998) concept of output-inflation *variability* trade-off is then relevant and can be neatly captured by sample standard errors. These in fact reveal that, in line with theory, pure inflation targeting would reduce inflation variability (from the observed 0.016 to the "new" 0.008) but it would also raise output variability (from 0.06 to 0.1, respectively).

Figure 7, Figure 8 and Figure 9 show for Germany, respectively, the interventions $(k_t^{ctr} - k_t^{new})$ on the inter-bank rate needed to make the inflation rate stationary around a target mean of 2%, the observed and the "new" inflation rate and the observed and the "new" output, using the derived control rule at all time points. As concern Figures 7 and 8, the comments are very similar to the ones for Italy. Yet the disinflationary gain of a consistent 2% rule in the early part of period is less pronounced than in Italy (probably, the BB policy was closer to the simulated rule than that of the BoI). By contrast, like Italy, Figure 9 shows that inflation targeting by means of control of the inter-bank rate would have greatly reduced output during the first years whereas it would have allowed for output gains in the terminal part of the period. It should be recalled that we are trying to see how the control process would have worked over a rather short time period, the post-reunification years. The time compression may thus force larger swings in the convergence process.

Figure 10 shows the scatter diagram of observed *vis-à-vis* "new" inflation and output realizations for Germany. Again, the result is similar to Italy, namely a clear structural shift of the long-run "AS curve" which is "horizontalized" around the 2% inflation target.

5. Conclusions

In this paper we have put forward an empirical extension of the CCC model of monetary policy presented in a previous work (PT (2005)). This model combines bank credit supply, as a means whereby monetary policy affects economic activity ("credit channel"), and interest rates on loans as a cost to firms ("cost channel"). The thrust of the model is that firms' reliance on bank loans makes *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. This yields a pattern of relationships consistent with the set of empirical regularities that are today regarded as the *explanandum* of monetary macroeconomics, with no recourse to additional non-competitive hypotheses. Moreover, the presumption arises that the CCC may also have permanent, rather than transitory, effects on real variables.

The emprical extensions of the model presented in this paper consisted of two parts. First, we have re-estimated the model for Italy (1986:1 to 1998:12) with a new measure of the credit risk premium, and estimated it also for Germany (1990:1 to 1998:12). The statistical methodology adopted has enabled us to apply a single integrated framework to 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. Statistics support the hypothesis that, in both countries, by way of the CCC transmission mechanism, the inter-bank rate - which turns out to be a weakly exogenous variable for the system of variables in both countries - 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.

Second, by exploiting the properties of Johansen-Juselius's theory of control, we have also provided a statistical test and measure that supports the hypothesis that the inter-bank rate qualifies as a control variable for output and inflation. By simulating a control rule of inflation, we have also shown that control is gained because innovations in the inter-bank rate exert a significant long-run impact on *both* the inflation *and* output stochastic paths. Graphically, this transmission mechanism shifts the long-run AS curve.

We believe that our main conclusions may be of general interest, at least for countries where firms significantly depend on bank credit. Italy and Germany are also major economies in the euro area, where inflation-targeting by means of inter-bank rates control is one official pillar of monetary policy, and where better understanding of country-specific transmission mechanisms is a priority for the monetary authority.

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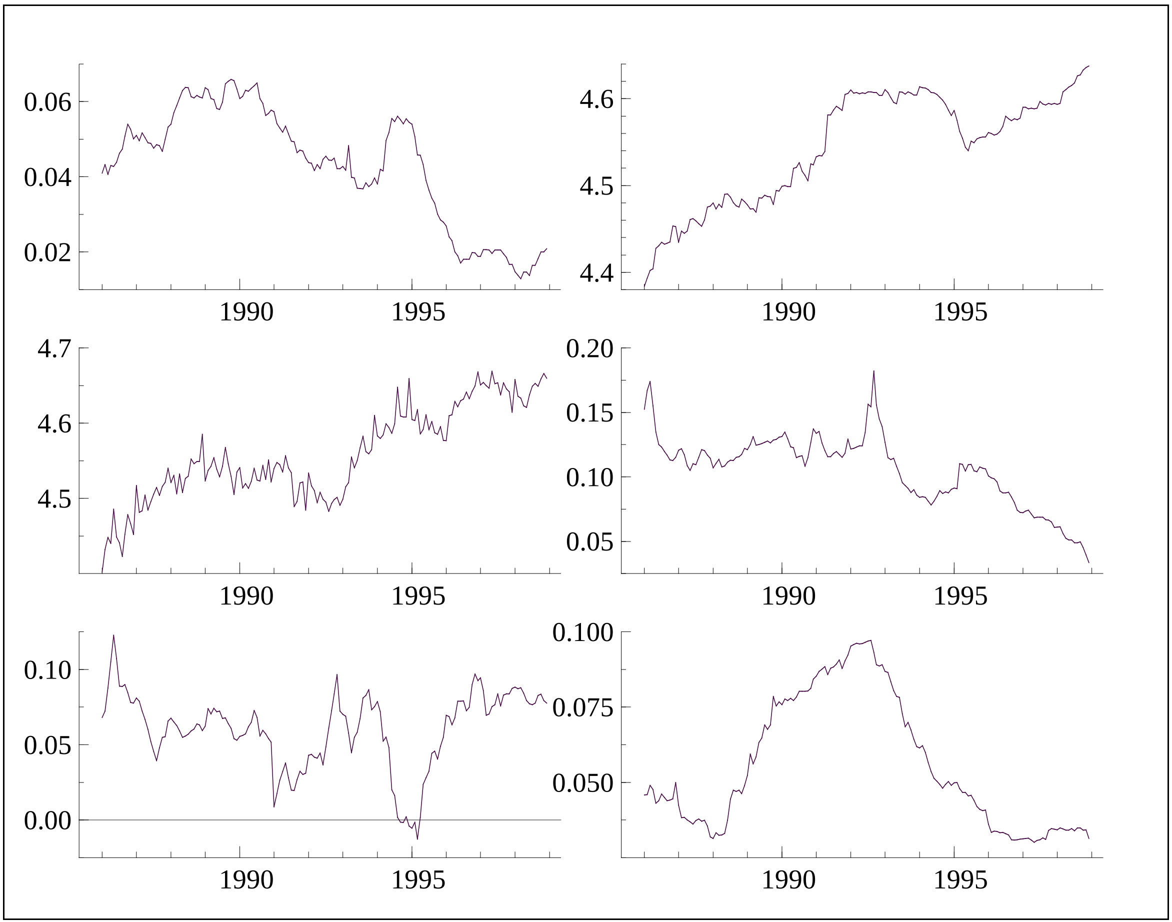
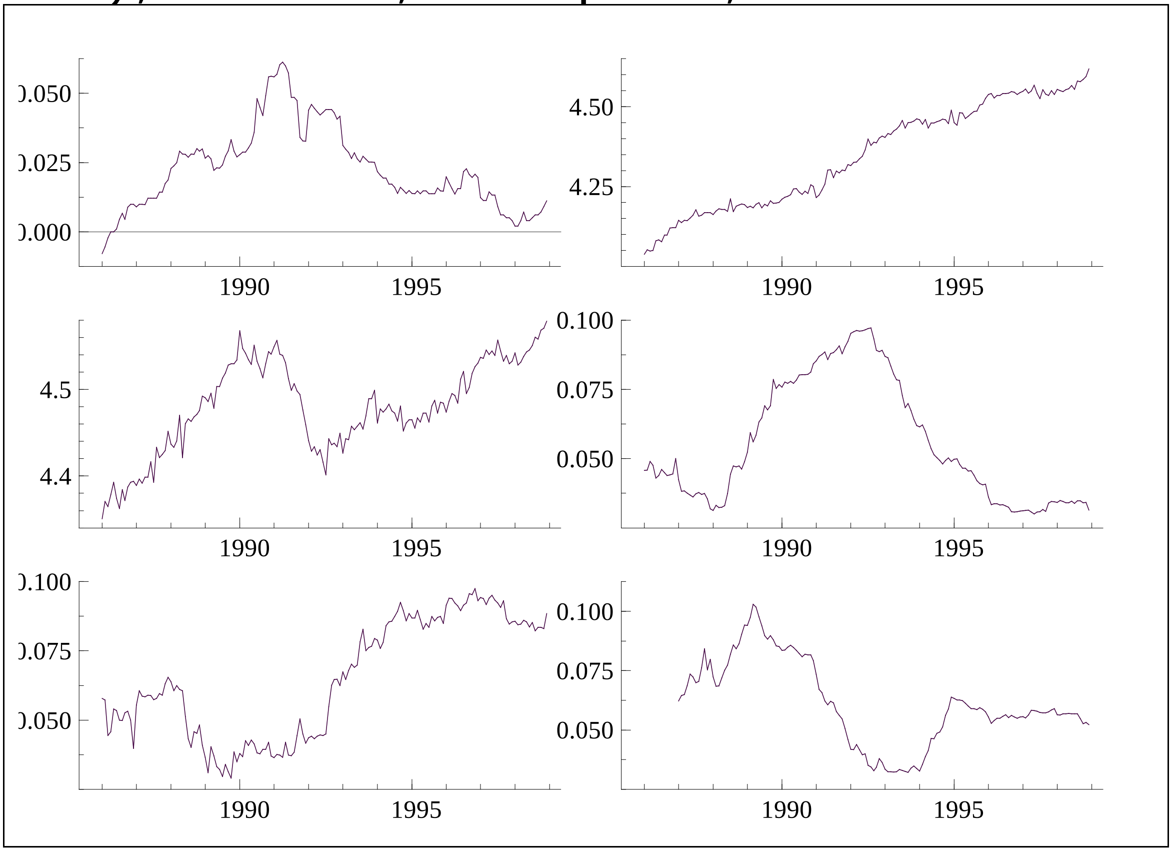
Figure 1. Italy, plots of variables (left to right): inflation rate[1]; index of industrial real wages[1]; index of industrial production (12 months ahead)[1]; inter-bank rate[2]; credit risk premium; German inter-bank rate[1]; Average rate on bank loans[2]



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

Figure 2. Germany, plots of variables (left to right): inflation rate[1]; index of industrial real wages[2,1]; index of industrial production (12 months ahead)[2]; inter-bank rate[2]; credit risk premium[3]; LIBOR US Dollar[4].



Sources: [1]OECD, *Statistical Compendium*; [2]IFS, *International Financial Statistics*; [3]Bundesbank; [4]Economagic.com.

Figure 3. Italy: representation of the inter-bank rate (solid line) and the derived intervention (dotted line) to make inflation stationary around 2%.

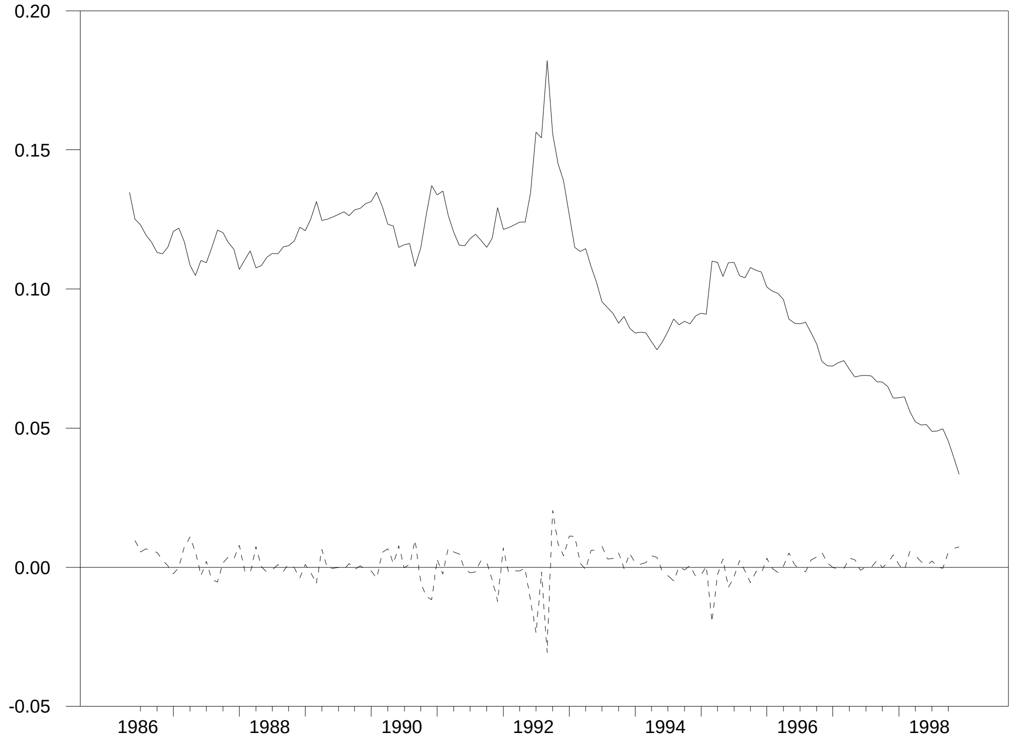


Figure 4. Italy: observed (solid line) **and "new" inflation** (dotted line)

Figure 5. Italy: observed (solid line) **and "new" output** (dotted line)

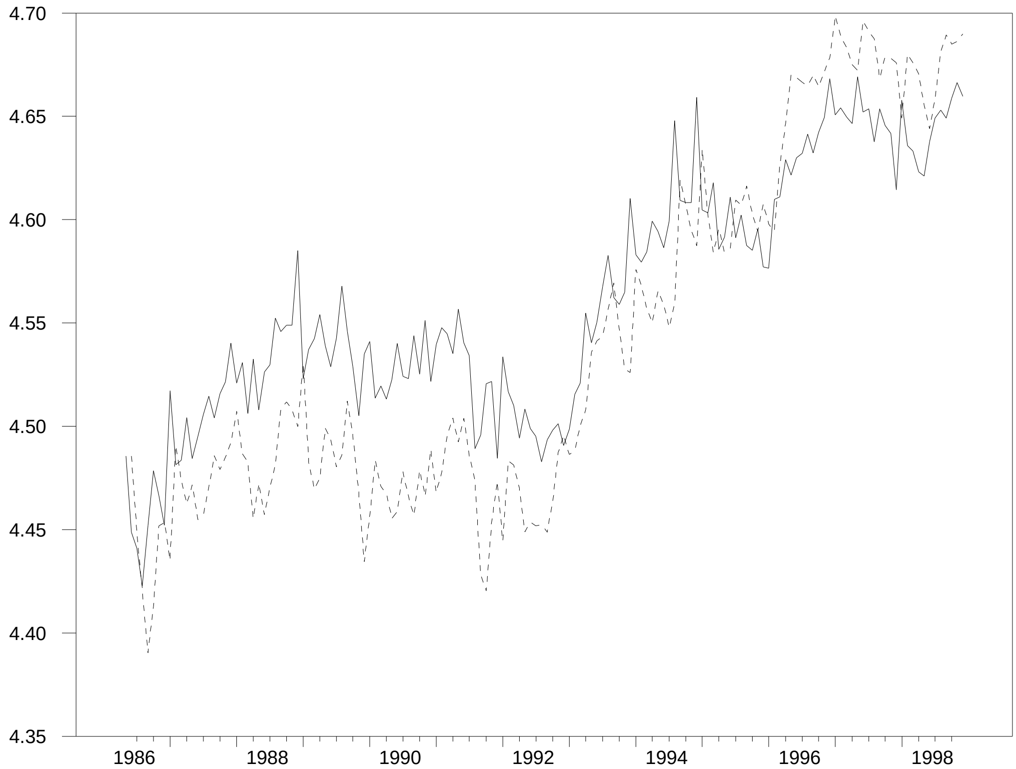


Figure 6. Italy: observed (dots) **and "new"** (stars) **AS curves**

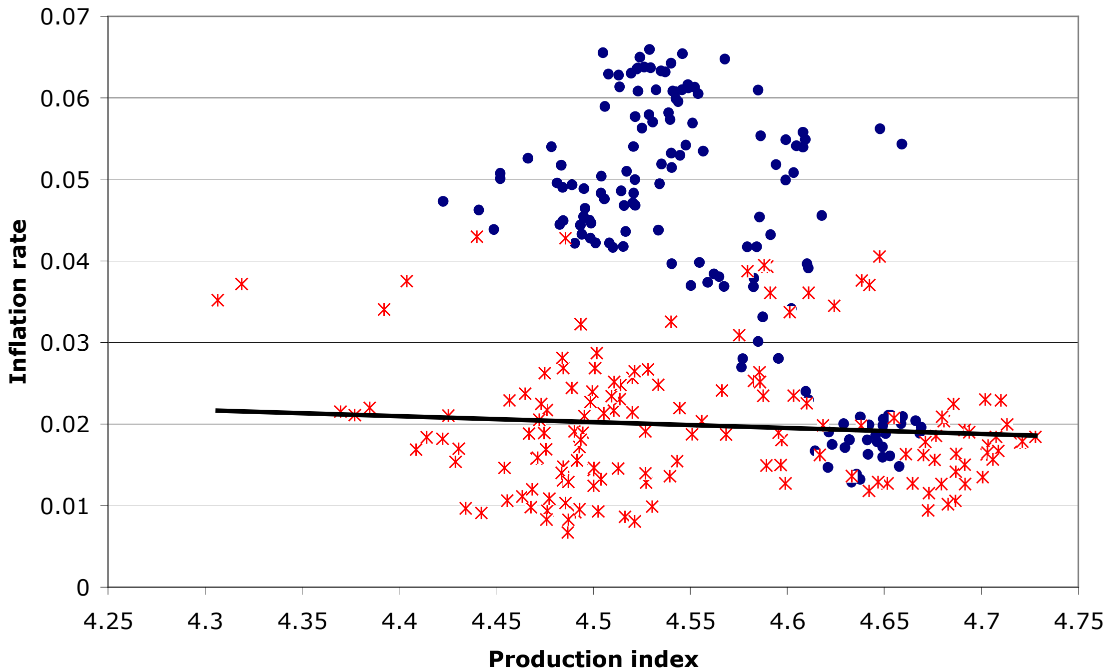


Figure 7. Germany: representation of the inter-bank rate (solid line**) and the derived intervention** (dotted line) **to make inflation stationary around 2%.**

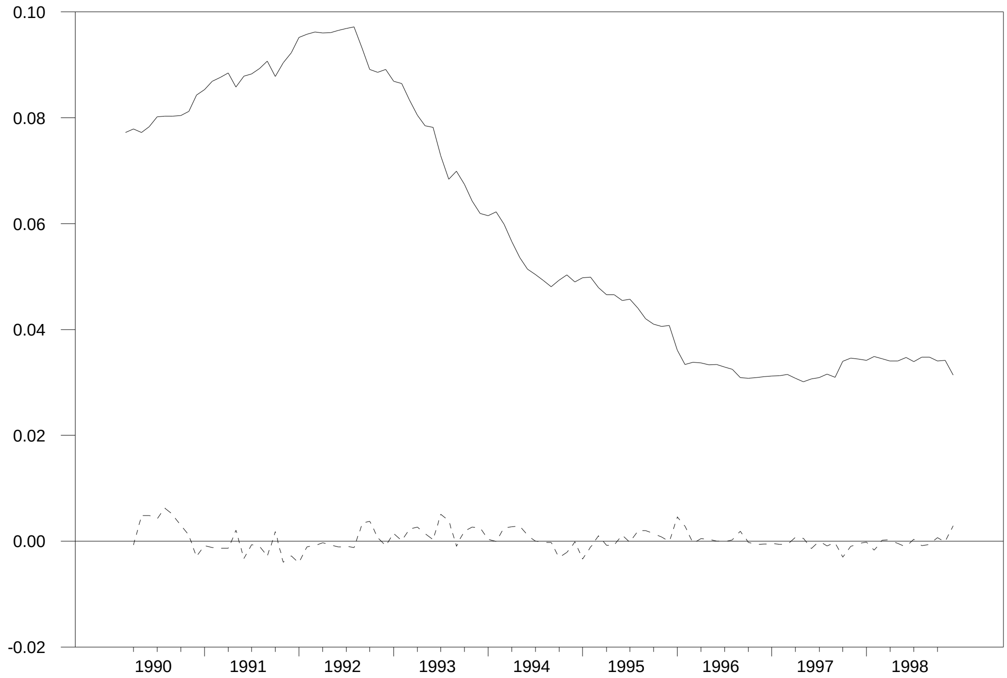


Figure 8. Germany: observed (solid line) **and "new" inflation** (dotted line)

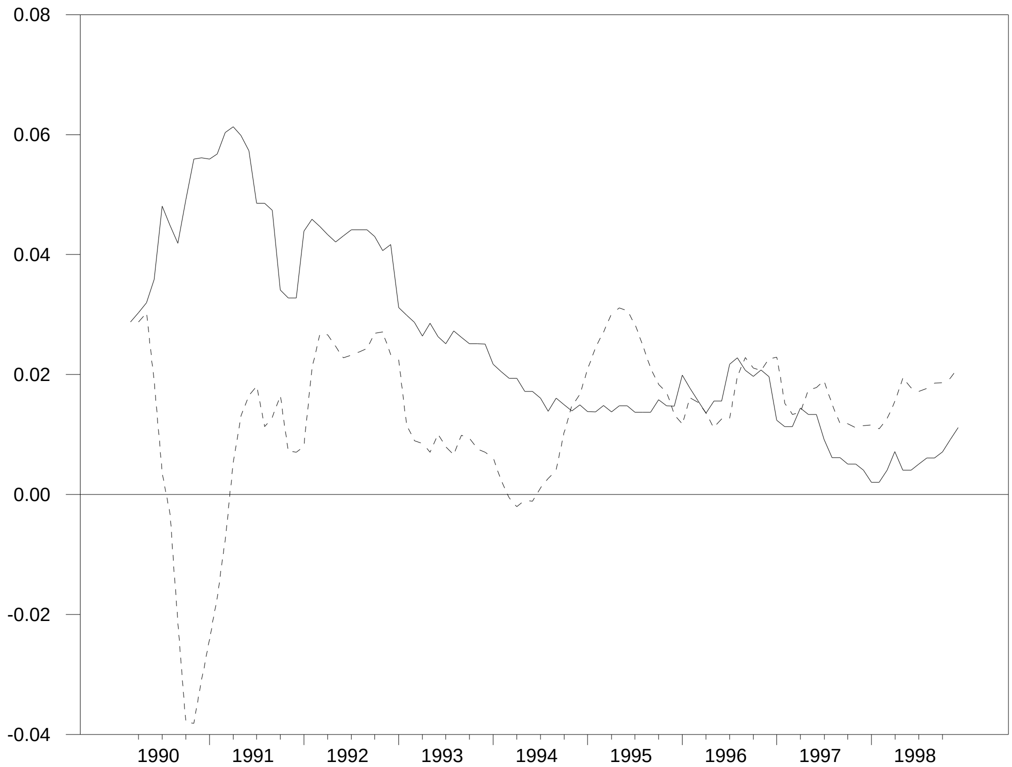


Figure 9. Germany: observed (solid line) **and "new" output** (dotted line)

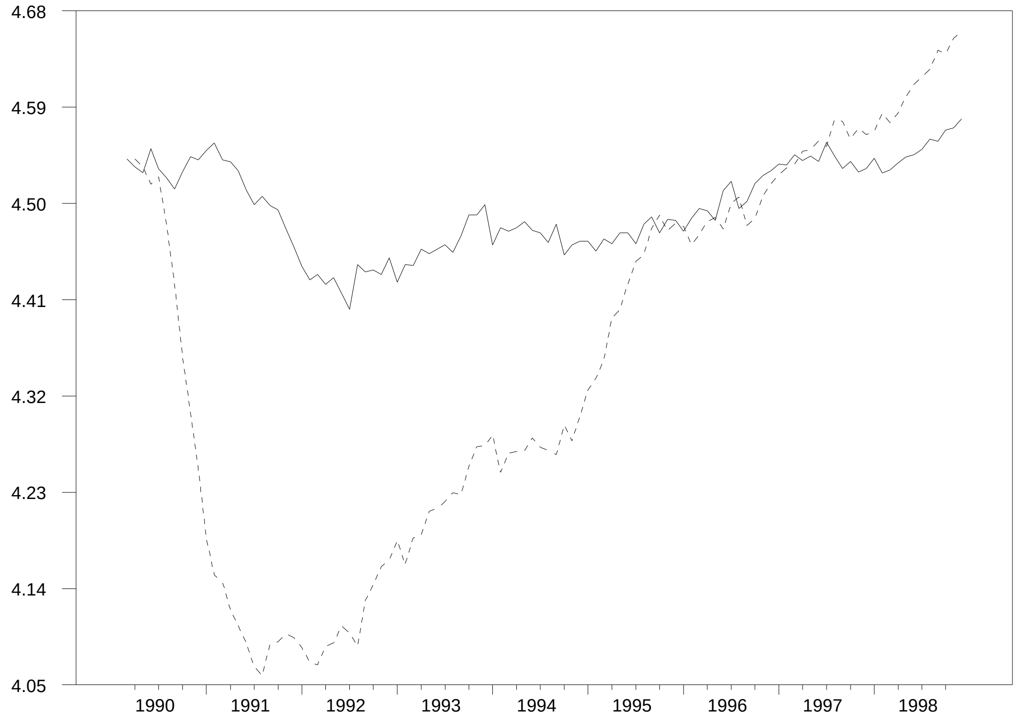
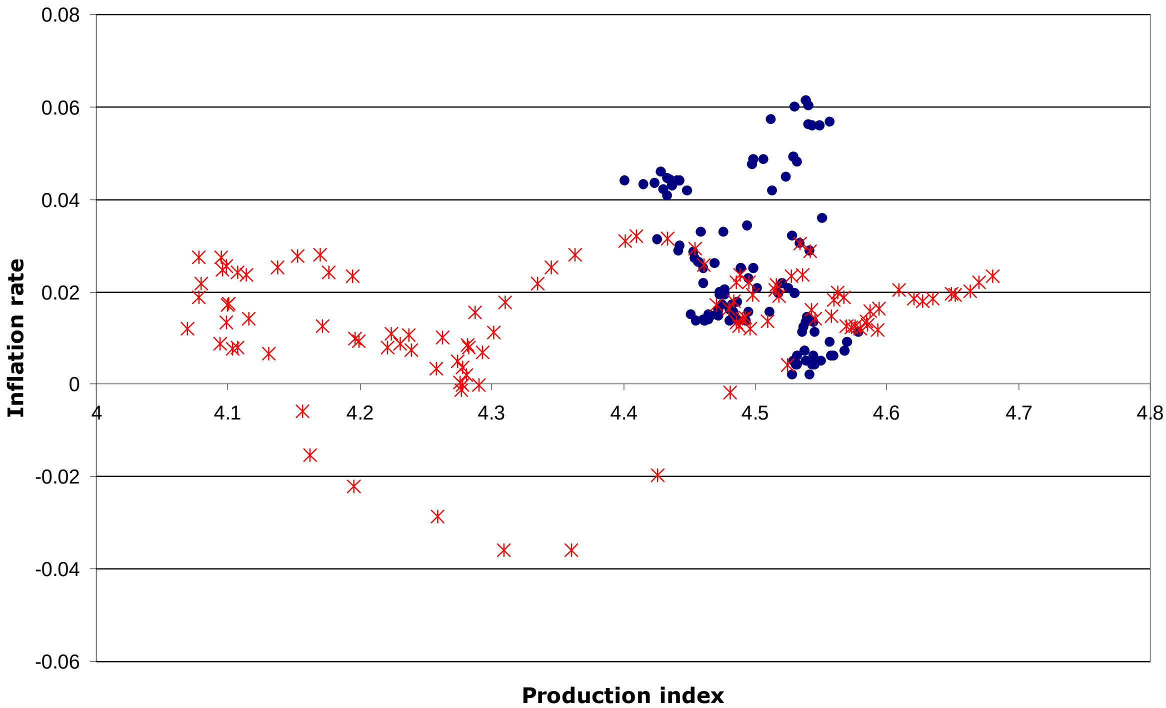


Figure 10. Germany: observed (dots) **and "new"** (stars) **AS curves**



Statistical Appendix

A1. The unrestricted cointegrated model

Given the p-dimensional (p = 5) observed process $\mathbf{y}'_t = [w_t, k_t, \rho_t, q_{t+12}, \pi_{t+12}]$ and the unrestricted vector autoregressive model written in error correction form (VECM), the LR trace test suggested an eigenvector decomposition of the long-run matrix $\mathbf{\Pi}$ into r = 3 stationary directions, the cointegration vectors, and $(p - r)$ = 2 nonstationary directions. With r = 3 the modulus of the largest stationary root in the model is 0.87 for Italy and 0.81 for Germany.

The singular matrix $\mathbf{\Pi}$, of rank r, has the representation $\mathbf{\Pi} = \boldsymbol{\alpha\beta}'$, where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are matrices of full rank r. The columns of $\boldsymbol{\beta}$ correspond to the r cointegrating relations, which represent the *long-run relationships* that can be detected among the variables $\mathbf{x}_t$ ("attractor set"), whereas the elements in the columns of $\boldsymbol{\alpha}$ are the adjustment coefficients of endogenous variables *towards* the long-run relationships. Associated with the $(p - r)$ = 2 nonstationary relations is a matrix, $\boldsymbol{\alpha}_\perp$, orthogonal to $\boldsymbol{\alpha}$, whose elements measure the extent to which cumulated stochastic shocks push the variables *along* their long-run relationships.

Table A.1 for Italy and Table A.2 for Germany report the eigenvector decomposition of $\mathbf{\Pi}$ into the r cointegrating relations, together with their adjustment coefficients. Using the information given by the covariances of the data and the finding of 3 cointegrating relations, the unrestricted relations were normalized with respect to the 3 variables that the theory indicates as "endogenous" (π_{t+12}, w_t, q_{t+12}) *vis-à-vis* the CCC "explanatory" variables (k_t, ρ_t, k^*_t) and the trend for Italy and (k_t, ρ_t, Lib_t) and the trend for Germany.

Table A.1. Italy: the stationary components of $\mathbf{y}_t$ (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	0.172	-0.218	-0.023	0.197	-0.543	0.000
$\hat{\boldsymbol{\beta}}_2'$	-1.191	1.000	1.743	-4.474	1.925	2.281	-0.006
$\hat{\boldsymbol{\beta}}_3'$	-0.299	0.917	1.000	1.708	0.507	-0.944	-0.002

	The adjustment coefficient matrix $\hat{\boldsymbol{\alpha}}$ (transposed)				
$\hat{\boldsymbol{\alpha}}_1'$	-0.011	**-0.326**	**0.987**	0.011	**-0.231**
$\hat{\boldsymbol{\alpha}}_2'$	**0.006**	0.006	-0.001	**0.008**	**-0.016**
$\hat{\boldsymbol{\alpha}}_3'$	0.009	**-0.051**	**-0.169**	**-0.033**	-0.005

Table A.2. Germany: the stationary components of $\mathbf{y}_t$ (bold α coefficients denote significance at 5%)

	π_{t+12}	w_t	q_{t+12}	k_t	ρ_t	Lib^*_t	$trend_t$
	The cointegrating matrix $\hat{\boldsymbol{\beta}}$ (transposed)						
$\hat{\boldsymbol{\beta}}_1'$	1.000	-0.371	-0.062	0.191	0.899	-0.097	0.001
$\hat{\boldsymbol{\beta}}_2'$	0.647	1.000	0.350	2.817	1.614	1.992	-0.002
$\hat{\boldsymbol{\beta}}_3'$	-1.687	-1.053	1.000	5.173	7.830	-0.985	0.002
	The adjustment coefficient matrix $\hat{\boldsymbol{\alpha}}$ (transposed)						
$\hat{\boldsymbol{\alpha}}_1'$	**-0.112**	**0.416**	0.037	0.026	**-0.077**		
$\hat{\boldsymbol{\alpha}}_2'$	**-0.061**	**-0.338**	0.135	**-0.017**	0.009		
$\hat{\boldsymbol{\alpha}}_3'$	0.000	0.021	**-0.097**	-0.004	-0.017		

As regards the adjustment coefficients matrices $\hat{\boldsymbol{\alpha}}$, since a zero row of $\boldsymbol{\alpha}$ is the condition for the corresponding variable to be weakly exogenous w.r.t. the cointegration relations, interbank rate for Italy as well as interbank rate and risk for Germany can be safely taken as exogenous, as required by the theoretical model[20].

[20] The relevant hypotheses 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: $\mathbf{R}' = [0,0,0,1,0]$ if we want to test the weak exogeneity hypothesis of k_t, $\mathbf{R}' = [0,0,0,0,1]$ if we want to test the weak exogeneity hypothesis of ρ_t. The *LR* test statistic, distributed as a χ^2_3, is equal to 2.316, with a *p-value*=0.510 for k_t in the Italian data set and is equal 3.600 wit a *p-value*=0.308 for k_t in the German data set; the same statistic is equal to 2.604, with a *p-value*=0.457 for ρ_t in the German data set. According to Garratt et al. (2003), weakly exogenous variables can be considered to be as "long-run forcing" variables, and, on this assumption, the cointegrating properties of the model can be analysed without having to specify their relative equations.

A2. The control problem

As in Johansen and Juselius (2003), we define as target variables the nonstationary variables $\mathbf{b'x}_t$ that we would like to control so that they become stationary with mean **b*.** To this end, we use a control rule and the instruments $\mathbf{a'x}_t$, where **a** and **b** are (*pxm*) matrices, with *m* corresponding to the number of target variables and of instruments.

The necessary condition for controllability is that $\mathbf{b'Ca} \neq \mathbf{0}$, where $\mathbf{C} = \boldsymbol{\beta}_{\perp}(\boldsymbol{\alpha}'_{\perp}\boldsymbol{\Gamma}\boldsymbol{\beta}_{\perp})^{-1}\boldsymbol{\alpha}'_{\perp}$, with $\boldsymbol{\alpha}_{\perp}$ and $\boldsymbol{\beta}_{\perp}$ ($p\times(p-r)$) matrices orthogonal to $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$, respectively, and $\boldsymbol{\Gamma} = \mathbf{I} - \sum_{i=1}^{n-1}\boldsymbol{\Gamma}_i$. Under this condition it is possible to define a recursive control rule (Johansen and Juselius, 2003, p.19), which takes the following form for our model :

$$\mathbf{y}_t^{ctr} = \mathbf{y}_t^{new} - \mathbf{a}(\mathbf{b'Ca})^{-1}[\mathbf{b'y}_t^{new} - \mathbf{b}^* + \mathbf{b'}(\mathbf{C\Gamma} - \mathbf{I}_p)\boldsymbol{\beta}(\boldsymbol{\beta\beta})^{-1}(\boldsymbol{\beta}'\mathbf{y}_t^{new} + \boldsymbol{\beta}'_z\mathbf{z}_t + \boldsymbol{\beta}'_t t)$$
$$- \mathbf{b'C}\sum\nolimits_{i=1}^{n-1}\boldsymbol{\Gamma}_i(\mathbf{y}_{t-1}^{ctr} - \mathbf{y}_t^{new})].$$

The next value for the process will therefore be the following:

$$\mathbf{y}_{t+1}^{new} = \mathbf{y}_t^{ctr} + \boldsymbol{\alpha}(\boldsymbol{\beta}'\mathbf{y}_t^{ctr} + \boldsymbol{\beta}'_z\mathbf{z}_t + \boldsymbol{\beta}'_t t) + \sum_{i=0}^{n-1}\boldsymbol{\Gamma}_{zi}\Delta\mathbf{z}_{t+1-i} + \sum_{i=0}^{n-2}\boldsymbol{\Gamma}_i\Delta\mathbf{y}_{t-i}^{ctr} + \boldsymbol{\mu}_0 + \boldsymbol{\Phi}\mathbf{D}_{t+1} + \boldsymbol{\varepsilon}_{t+1}.$$

We have used the estimated parameters and the residuals of the VECM model to generate $\mathbf{y}_t^{ctr}$, $\mathbf{y}_t^{new}$ and the intervention $\mathbf{v}_t$ of Section 3.4.

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