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Abstract

In this paper we present an empirical analysis of the "credit-cost channel" (CCC) of monetary policy transmission. 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 the CCC makes both aggregate demand and aggregate supply dependent on monetary policy. As a consequence a) credit market conditions (e.g. risk spreads) are important sources and indicators of macroeconomic shocks, b) the real effects of monetary policy are larger and persistent. We have applied the Johansen-Juselius CVAR methodology to Italy and Germany in the "hard" EMS period and in the EMU period. The short-run and long-run effects of the CCC are detectable for both countries in both periods. We have also replicated the Johansen-Juselius technique for the simulation of rule-based stabilization policy for both Italy and Germany in the EMU period. As a result, we have found 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 stochastic equilibrium paths of both inflation and output.

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

JEL codes: E51, C32

1 Introduction

Analysis of the channels through which monetary policy operates and affects nominal as well as real macroeconomic variables has long been a matter of research. The issue is of general interest, but it is particularly important for monetary policy conduct in the European Monetary Union (EMU) because it is well known that such transmission channels may differ across countries with different economic structures and institutional legacies (e.g. Dornbusch et al. (1998), Angeloni et al. (2003)).

A good starting point is the distinction between "market-based" and "bank-based" economies popularized by scholars in comparative financial systems. Market-based economies are typically associated with the traditional demand-side effects of interest rates mainly affecting private sector's expenditure directly through efficient financial markets. Bank-based economies are instead characterized by private expenditure largely dependent on bank credit, with relatively underutilized financial markets. Notably, today's reference model adopted by the major central banks – the so called New Keynesian model with sticky prices (e.g. Woodford (2003)) – takes an efficient market-based structure for granted and gives no role to play to banks[1]. Since major continental economies such as those of Germany, France, Italy, and Spain have long been regarded as typical bank-based economies, it seems important to understand monetary transmission in economies of this type, and to monitor possible modifications away from bank-based structures resulting from monetary and financial integration. To date, empirical analyses of different transmission mechanisms in the EMU have delivered mixed and uncertain results (e.g. Angeloni et al. (2003), Angeloni and Ehrmann (2003)).

This state of the art may be due to limitations in the short-lived data set of the EMU as well as to some deficiencies in the theoretical framework adopted. Study of monetary policy in connection with the banking system has a long tradition (Trautwein (2000)), which was revived and completely refounded with modern methodological tools in the 1980s when the economics of imperfect capital markets was developed (e.g. Greenwald and Stiglitz (2003)). The specificity of bank-based economies was rooted in the asymmetric information between private borrowers and lenders that prevents development of direct financial relationships. Against this background, a major contribution to monetary policy has consisted in the so-called "credit channel", which refers to the means by which monetary policy affects aggregate demand via the credit supply of intermediary institutions (see e.g. Gertler and Gilchrist (1993), Bernanke and Gertler (1995) for surveys). A first round of empirical research on European countries tended to confirm the importance of the credit channel (e.g. Fiorentini and Tamborini (2001) for a survey). However, this class of models had the empirical shortcoming that they made it difficult to identify the causal direction between aggregate demand and outstanding credit.

Another related research path has pursued the idea that the credit channel of monetary

[1] The current financial crisis has seriously shaken the reliability of this foundation of the New Keynesian model: see e.g. Crockett (2003), Christiano et al. (2007b), Leijonhufvud (2007).

policy may affect not only aggregate demand but also aggregate supply. The financial constraints of firms may limit the demand for physical capital as well as for circulating capital. Hence, in a bank-dependent economy, also current production comes to depend on the terms and conditions of credit availability. This creates a so-called "credit-cost channel" (CCC) of monetary policy, meaning that the bank interest rate is also a production cost to firms. Greenwald and Stiglitz (1988, 1993), Christiano and Eichenbaum (1992), and Christiano et al. (1997) provide early examples of this CCC explicitly related to firms being financed by intermediaries in a context of imperfect capital markets[2]. Christiano et al. (2007a), Pfajfar and Santoro (2007), Tamborini (2009), and De Fiore and Tristani (2009), exemplify developments of this line of inquiry, showing that blending the credit and cost channels may provide a consistent framework for monetary policy analysis in bank-based economies and overcome the weaknesses of the two separate approaches.

The thrust of this class of CCC models is that monetary policy affects economic activity as policy-induced changes in bank interest rates, given rational expectations of future inflation, exert a *real* credit-cost effect on firms that shifts labour demand and output supply in parallel with labour supply and aggregate demand. Under suitable treatment of capital market imperfections, a specific role for credit risk (re)emerges both for business cycle analysis and for monetary policy design (e.g. Tamborini (2009), De Fiore and Tristani (2009)). Moreover, joint consideration of aggregate demand and supply effects in a general equilibrium set-up makes it possible to control for a variety of real-nominal effects of monetary policy. Put briefly, the key differences made by the CCC with respect to the traditional transmission mechanisms are that 1) credit market conditions (e.g. risk spreads) are important sources and indicators of macroeconomic shocks, 2) the real effects of monetary policy are larger and persistent.

In this paper we present an econometric analysis of the CCC hypothesis for two EMU test countries, Italy and Germany. They are not only two of the largest EMU economies, but are also representative of a wider range of bank-based economies (think of Eastern Europe's new entrants in the EMU). Whereas both countries have been subject to numerous investigations of the traditional demand effects of the credit channel, to our knowledge little attention has been devoted to the supply-side effects[3].

In this perspective, our econometric analysis is the empirical counterpart of the CCC

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[2] More recently, there has been growing interest in the cost channel of monetary policy *per se*, both in partial equilibrium (e.g. Bart and Ramey (2001)) and in general equilibrium (e.g. Ravenna and Walsh (2006), Chowdhury et al. (2006)). However, most of these recent models do not treat financial constraints and firms-banks relationships explicitly, and simply plug the policy rate into the production function or the Phillips curve. For the reasons explained below, this is not satisfactory.

[3] 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. A similar framework is adopted by Chowdhury et al. (2006) who estimate New Keynesian Phillips Curves augmented with short-term interest rates for a number of industrialized countries. They find evidence of a significant cost channel of interest rates for Italy and France, less so for Germany and Japan.

model proposed by Tamborini (2009). Unlike the other authors cited, who in various ways introduce the CCC into a New-Keynesian model with sticky prices and other "frictions", this model provides a theoretical assessment of the CCC in a competitive flex-price economy where the only "friction" is the bank-dependency of firms. This approach presents some comparative advantages. Firstly, Greenwald and Stiglitz (1988, 1993), and Christiano et al. (1997), have stressed that, in the presence of the CCC, real effects of monetary policy may also arise in a competitive flex-price economy. According to the latter authors, a broader set of stylized facts can be obtained than with sole demand-side effects with sticky prices. In particular, 1) monetary policy interventions negatively affect output, inflation *and real wages*;[4] 2) they have *larger effects on output* than prices even though the latter are flexible, 3) real effects *persist* in spite of full price adjustment. Tamborini (2009) has shown analytically under what testable conditions a CCC model of this type may generate such a pattern of macroeconomic relationships. The key lies precisely in the supply-side effects of the CCC. Secondly, the absence of "frictions" other than bank-dependency allows for clear-cut identification of the presence and strength of the CCC.

As regards the econometric technique, we have opted for the Johansen-Juselius cointegrated VAR approach (Johansen (1996), Juselius (2006)). We have chosen this approach because the thrust of the CCC model to be tested is the existence of long-run structural relationships between the policy interest rate and the other relevant variables: real wages, output and inflation. In addition, the same econometric apparatus can be exploited to conduct a formal test of the idea that the interest rate is an instrument of control for the central bank (Johansen and Juselius (2003)).

As regards the time period of analysis, in consideration of the institutional break represented by the EMU we have decided to operate on two separate sub-periods for each country, 1986:1-1998:12 and 1999:1-2007:9 for Italy, 1990:1-1998:12 and 1999:1-2007:9 for Germany. The first sub-period coincides with pre-EMU life in the so-called "hard" European Monetary System (EMS) (Germany's sub-period starts from 1990 in order to bypass the unification shock, which creates a number of outliers in the data). The second sub-period coincides with inception of the EMU and the coming of both countries under the control of the European Central Bank (ECB) in orderly operational conditions (that is, prior to the bank market turmoil which began in the last quarter of 2007). We expect comparison of econometric results across the different sub-periods and countries to be informative about whether

- the CCC hypothesis holds for both countries
- it held in the past but no longer applies after ten years of EMU
- there is any informative discrepancy between the observed data on monetary policy control by the ECB and the inferences that can be drawn from the estimated CCC model.

Our presentation divides into three parts. The first (section 2) sketches the CCC model

[4] A well-known shortcoming of the Old and New Keynesian sticky price hypothesis is that it implies that the real wages are instead *countercyclical* after a monetary shock (Christiano et al. (1997)).

in order to identify the structural relationships to be examined. The second one (section 3) presents the econometric results based on cointegration vectors, which do not reject the CCC hypothesis for both Italy and Germany and across sub-periods. The third part (section 4) is devoted to policy implications according to the Johansen-Juselius (2003) control theory: that is, a simulation of a rule-based stabilization policy where the control variable (the inter-bank rate) is aimed at the target variable of 2% inflation rate. This simulation is performed for both Italy and Germany, on the basis of the respective estimated CCC models for the EMU sub-period. As a result, we find confirmation that inflation-targeting via inter-bank rate control, grafted onto the estimated CCC model, would indeed stabilize inflation through structural shifts of the equilibrium paths of both output and inflation. The simulation results are then compared with actual data generated under the ECB policy. Summary and conclusions follow in section 5.

2 The structural relationships identifying the "credit-cost channel"

The key features of CCC models in general are that:

- production takes time,
- firms need external funds in advance to finance working capital (typically the wage bill)
- these funds come from bank credit.

Consequently, the connection with monetary policy is given by the labour demand function of the bank-dependent firm, and the bank-firm relationship is at the centre of the stage.

According to the competitive general equilibrium characterization proposed by Tamborini (2009)[5], when firms plan production at any time t for sale at time t, $Q(t)_{t+1}$, and need external funds to finance working capital (the wage bill for simplicity), they are uncertain about their future revenue from output sales. The unit revenue 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 (that is, firms operate with rational expectations). Under costly state verification, the efficient financial contract is a standard debt contract with a deposit and lending intermediary. These are commercial banks that collect deposits from households and extend loans to firms under default risk (the probability that total revenue may fall short of outstanding debt). Hedging is possible by borrowing reserves from the central bank at the policy rate k_t. As a result, and under suitable conditions, all banks charge the same competitive interest rate $r_t = k_t + \rho_t$, where ρ_t is the risk premium. In this set-up, firms that maximize expected profit plan output up to the point where the marginal product of labour equals its expected real cost Γ_t. This is given by the current real wage rate $\underline{W}_t$ increased by the (gross) expected real interest rate $\underline{R}_t$, $\Gamma_t \equiv \underline{W}_t\underline{R}_t$ (where $\underline{R}_t \equiv (1 + r_t)E_t(1 + \pi_{t+1})^{-1}$ and $1 + \pi_{t+1} \equiv P_{t+1}/P_t$). Given a neoclassical production function, changes in the determinants of the bank interest rate affect labour demand and

[5] Here we provide only a non-technical description of the model. Besides the original source, an analytical presentation of the model is available in a separate appendix.

planned output with negative sign.

Note also that co-movements in labour demand and credit supply correspond to equal variations in the wage bill and households' deposits, which are also a component of utility-maximizing consumption demand. This, besides ensuring the (risk-adjusted) cash flow equilibrium of banks, implies that aggregate demand and supply also move in parallel.

The thrust of the model is that variations of k_t can, under certain function-parametric conditions[6], generate a pattern of relationships which is consistent with the empirical regularities observed in major industrialized countries (Christiano et al. (1997)), i.e.

$$\mathrm{d}\underline{W}_t/\mathrm{d}k_t < 0,\ \mathrm{d}Q(t)_{t+1}/\mathrm{d}k_t < 0, \mathrm{d}\pi_{t+1}/\mathrm{d}k_t < 0,$$

$$|\mathrm{d}Q(t)_{t+1}/\mathrm{d}k_t| > |\mathrm{d}\pi_{t+1}/dk_t|$$

to the exclusion of ancillary hypotheses like monopolistic competition or price stickiness. Note that the usual negative relationships with output and inflation are associated with the negative one with the real wage rate, which indicates that the CCC in fact operates through the supply side with no frictions. Hence, the identification of the CCC hypothesis hinges on this unique pattern of signs of the variables k_t and ρ_t in the equations for π_{t+1}, $\underline{W}_t$, and $Q(t)_{t+1}$. Overall these results are analogous to those of the New Keynesian CCC model developed by De Fiore and Tristani (2009). However, the latter is crucially dependent on the non-zero parameter of price stickiness, which is instead excluded from this model[7].

Though this is not our main focus here, it is worth noting that the model predicts the same pattern of effects for changes in the credit risk premium ρ_t. This feature – which is also found by De Fiore and Tristani (2009) – is relevant to the current debate on the role of credit risk in the business cycle, and on the appropriate response of monetary policy, making a clear case for the use of the policy rate as a means to counteract autonomous shifts in the credit risk premium.

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. For the reasons already explained, the model has been estimated for both Italy and Germany over two separate sub-periods: 1986:1-1998:12 and 1999:1-2007:9 for Italy, 1990:1-1998:12 and 1999:1-2007:9 for Germany. Since our aim was to analyze the long-term, structural relationships in a fully specified system, for the reasons stated in the

[6] See Tamborini (2009, sec. 4). The relevant conditions result from five structural parameters, the labour input coefficient of the production function, the real balance effect and the intertemporal substitution effect in the consumption function, the real wage elasticity and intertemporal substitution effect in the labour supply function. The model also shows that different configurations of these parameters may instead generate different macroeconomic outcomes, such as the well-known "price puzzle" (monetary restrictions followed by higher prices) pointed out by Sims (1992).

[7] In models with monopolistic competition the effect of a change of interest rates on inflation is dampened or even reversed because firms transfer higher or lower interest rates onto higher or lower prices (see also Barth and Ramey (2001) and Chowdhury et al. (2006)). Price stickiness further enhances this effect. However, these concomitant "frictions" blur the identification of the CCC *per se*.

Introduction we have chosen the structural cointegration approach developed by Johansen (1996), Juselius (2006) and Johansen and Juselius (2003), for the estimation and identification of the long-run structural relationships, of the driving forces among the theoretically relevant variables, and for the evaluation of the policy variable as a control variable of the system[8]. Each point is developed in the subsequent paragraphs.

3.1 The data

Our econometric exercise has been designed in order to identify a common "core set" of variables of the CCC hypothesis, while introducing specific variables to account for specific country/period features.

To begin with the core CCC variables, a preliminary observation is that our theoretical focus is on the policy rate k_t. Thus, instead of including the rate on bank loans as an independent variable, we have directly considered its two components k_t and the credit-risk premium ρ_t. The latter mainly makes it possible to control for autonomous changes in credit conditions. Hence ρ_t, a problematic variable to measure, is not crucial for the significance of the model. Secondly, our theoretical model has a forward-looking structure. At each point in time, firms plan production on the basis of uncertain future revenue, so that their economic decision depends on the expected price level and therefore on expected inflation. This is an important characteristic of the CCC hypothesis, since it paves the way to bank dependence and default risk. Hence we have decided to replicate the forward-looking structure of the CCC hypothesis in the econometric model by introducing appropriate leads in the relevant variables.

Therefore, the core CCC variables of interest for both countries are:

- the (log of) 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 a proxy given by the short-term inter-bank rate
- the credit risk premium ρ_t, not observable, for which we have adopted, again as a proxy, a transformation of the spread between the bank lending rate and a reference rate[9]
- the (log of) output q_{t+s}, given by the industrial production index, where s indicates the time lead
- the inflation rate, $\pi_{t+s,}$ measured by the consumer price index10.

[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] In light of other empirical studies (e.g. Chiades and Gambacorta (2004)) we have opted for slightly different reference rates for the two countries, namely, the medium-term yield rate of government bonds for Italy and the money market rate for Germany (basically, the reference rate for primary bank loans).

[10] Since the CCC model assumes flexible prices and rational expectations, we have taken the actual inflation rate on the same time lead as output as a proxy for the theoretically expected inflation. A time lead of 12 months has been chosen empirically by means of sensitivity tests. Recall that the time lead *s* should capture the theoretical gestation time of output and the related time-horizon for expected inflation. Consequently, gestation time has the same empirical effect as the sticky-price hypothesis: that is, the *observed* changes in

(sources and charts of the variables, for each sub-period, are given in figures 1, 2, 3 and 4)

As to the pre-EMU sub-period, it should be borne in mind that Germany and Italy differed historically in their conduct of monetary policy. The role of the credit market in the monetary transmission mechanism was carefully monitored by the Italian monetary authorities, and it 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 rooted 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. More importantly, the pre-EMU sub-period that we have chosen for the two countries saw a substantial convergence between the policy frameworks in the two countries, first under the pressure of the exchange-rate constraints of the EMS, and subsequently during the transition to the single currency (Angeloni (1994), Visco (1995)). In the second half of the 1980s, the BoI abandoned the pervasive and recurrent administrative interventions that had 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 ECB (see ECB (2008)). Hence we have felt reasonably confident in our choice of applying the same structural model to both countries.

Yet some contour diversifications have been necessary. The most important one relates to the well-known "asymmetry" between Germany and the other countries participating in the EMS. Germany had *de facto* the power to choose an independent monetary stance (domestic nominal policy rate) while Italy, like all the other N–1 countries, faced constraints on domestic monetary policy and interest rate setting due to strong exchange-rate targeting and high capital mobility. Therefore, we have included the German inter-bank rate, k^*_t , in Italy's data set[11].

Although Germany's monetary policy was almost unconstrained in the EMS, it was nonetheless concerned with the external exchange value of the mark, in particular *vis-à-vis*

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output and prices occur with a time *lag* after the *observed* change in the policy rate. 12 months is in fact a common time lag associated with the effects of changes in policy rates.

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

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

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 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.

the dollar. Hence, we have also added to Germany's data set an exogenous foreign variable to control for world monetary conditions, namely the 3-month LIBOR in US dollars, Lib_t.

In the EMU sub-period we have treated the two countries' models symmetrically in their monetary part too. In order to give continuity to the data set, we have used for each country the same inter-bank rate series of the first sub-period as well as the previous domestic reference rate to compute the respective credit-risk premium. The underlying assumption is that these rates now respond to the ECB common monetary policy in both countries up to unspecified local factors. Like the BB, the ECB, too, should to some extent take account of the evolution of monetary rates worldwide. Hence we have also retained the LIBOR in US dollars in the set of variables of both countries.

3.2. Econometric results

Since the Johansen CVAR estimation procedure is by now standardized, here we concentrate on the results, whereas statistical details are confined to the footnotes.

To start with, for each country we have defined the same p-dimensional (p = 5) observed process $\mathbf{y}'_t = [\pi^i_{t+12}, w^i_t, q^i_{t+12}, k^i_t, \rho^i_t]$, ($i$ = ITA, GER). To avoid cumbersome notation, in what follows we shall drop the country superscript whenever the variable refers to the country under examination. Then we have assumed an unrestricted vector autoregressive (VAR) model written in error correction form (VECM). The model has been augmented, in both countries, to include the appropriate country/period exogenous variables $\mathbf{z}_t$ indicated above, and deterministic terms. The resulting general equation is the following:

$$(1) \qquad \Delta\mathbf{y}_t = \sum_{i=1}^{n-1} \mathbf{\Gamma}_{yi}\Delta\mathbf{y}_{t-i} + \sum_{i=0}^{n-1} \mathbf{\Gamma}_{zi}\Delta\mathbf{z}_{t-i} + \mathbf{\Pi}\mathbf{x}_{t-1} + \boldsymbol{\mu}_0 + \boldsymbol{\mu}_1 t + \mathbf{\Phi}\mathbf{D}_t + \boldsymbol{\varepsilon}_t \ ,$$

where $\mathbf{x}_t = [\mathbf{y}_t, \mathbf{z}_t]$, $\boldsymbol{\varepsilon}_t \sim IN_p(\mathbf{0}, \mathbf{\Omega})$ is a vector of disturbances, Δ is the first difference operator and $\mathbf{\Gamma}_{yi}$, $\mathbf{\Gamma}_{zi}$, $\mathbf{\Pi}$ and $\mathbf{\Phi}$ 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[12], a vector of intervention dummies $\mathbf{D}_t$. The singular matrix $\mathbf{\Pi}$, of reduced rank r, has the representation $\mathbf{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$, where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are matrices of full column rank r, with $\boldsymbol{\beta}' = \left[\boldsymbol{\beta}'_y, \boldsymbol{\beta}'_z\right]$. 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.

3.2.1 Italy 1986 - 1998

In the model specification search, with $\mathbf{z}_t = [k_t^{GER}]$, the choice of the dummy variables and of the number n = 3 of lags in (3.1) has been determined on the basis of misspecification tests[13].

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[12] 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.

[13] In order to obtain residuals close to Normality, we introduced five permanent intervention dummies and two transitory intervention dummies to account for the exit of the Italian Lira from the EMS in 1992 and for few

In order to determine the cointegration rank r, we have examined as much data information as possible, and not just the formal trace test procedure[14]. The latter has given $r = 2$ at a significance level of 5% and $r = 3$ at 10%, using the simulated critical values. On this basis, and in the light of additional statistical information[15] we have been led to accept $r = 3$.

As concerns long-run weak exogeneity, with $r = 3$ our tests have shown that shocks to k_t were pushing the system in this period, and that inflation, real wages and production adjusted to long-run disequilibrium errors. These results are consistent with our CCC model, which assumes the inter-bank rate to be the exogenous driving variable[16].

In order to obtain a fully identified long-run β structure, and given $r = 3$, we have sought for at least two restrictions on each relation in addition to normalization. The unrestricted cointegrating relations have been normalized with respect to the three variables that the theory, and preliminary evidence, indicate as "endogenous" [π_{t+12}, w_t, q_{t+12}] *vis-à-vis* the CCC "explanatory" variables [k_t, ρ_t, k_t^{GER}] and the trend. Identification has then been accomplished by using the information coming from the theoretical model[17]. Below

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other events. The permanent intervention dummies were defined for 1991:1, 1991:5, 1992:7, 1994:4 and 1995:3. The transitory intervention dummies were defined for shocks of opposite signs in 1992:9-1992:10, and in 1993:3-1993:4. The results of misspecification tests for the unrestricted VAR(3) model with dummies have been the following: 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 0.398; 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.

[14] It has been shown that the power of the test is often low for many economically interesting alternatives when the sample size is rather small. It is therefore advisable in such cases to use all the sources of information to determine the rank. These include the unrestricted characteristic roots of the companion matrix, and the graphs of the recursively calculated trace tests.

[15] Starting from the examination of the eigenvalues of the trace test we have obtained four values, corresponding to $r = 0, 1, 2, 3$, quite close to each other, and just one very small value. The modulus of the largest unrestricted characteristic root of the companion matrix is 0.872 for $r = 3$ and 0.909 for $r = 2$, indicating $r = 3$ as preferable. Moreover, inspection of the graphs of the unrestricted cointegrating relations has suggested stationarity of the third one. Even the graphs of the recursively calculated trace tests exhibit a clear linear growth for $r = 1, 2, 3$, suggesting that the first three cointegration relations have been quite stable over the sample period. For the other relations the trace test components grow more slowly and are not even close to the 5% critical test value.

[16] Assuming $r = 3$ and exploiting the information given by the estimated $\mathbf{\Pi}$ matrix, whose rows contain some tentative indications concerning the long-run relations and their adjustment dynamics, we have found no evidence of long-run relations in the inter-bank rate equation, thus confirming the weak exogeneity of k_t. The real wage equation and the production equation contain most of the significant coefficients, implying that they are more sensitive to long-run disequilibrium, and they prove to be equilibrium correcting.

[17] As shown in Chapter 10 of Juselius (2006), 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 β without having jointly to identify the short-run structure. Information provided by the theoretical model is that, of the three endogenous variables, one is forward-looking (inflation), one is contemporaneous (wage) and one has a gestation period (output). Therefore, we have imposed a "cascade" structure such that the first variable affects the others but not the other way round, the second affects the third but not the other way round, and the third is affected by the other two.

we report the final just-identified structural long-run relations (t-statistics in parentheses) together with the value of the *LR* test[18], and the adjustment dynamics of the system variables to equilibrium errors:

Table 1. Italy: the estimated identified long-run structure and the estimated adjustment coefficients over the sub-period 1986-1998 (bold coefficients denote significance at 10%)

	π_{t+12}	w_t	q_{t+12}	k_t	ρ_t	k_t^{GER}	$trend_t$
			$\chi^2_3 = 2.318,$	$p\text{-}value = 0.509$			
$\hat{\beta}_1$	**1.000**	-	-	**3.216**	**-1.995**	-1.236	**0.002**
				(3.068)	(-3.175)	(-1.509)	(4.405)
$\hat{\alpha}_1$	**-0.006**	-0.007	**-0.082**	0.000	**0.015**		
	(-2.887)	(-0.867)	(-3.639)		(1.852)		
$\hat{\beta}_2$	**2.806**	**1.000**	-	**0.658**	**0.658**	**-1.894**	-0.000
	(5.175)			(3.574)	(3.574)	(-8.895)	(-1.302)
$\hat{\alpha}_2$	-0.001	**-0.109**	**0.286**	0.000	**-0.077**		
	(-0.139)	(-3.799)	(3.597)		(-2.612)		
$\hat{\beta}_3$	-0.704	**0.655**	**1.000**	**0.417**	**0.417**	-	**-0.002**
	(-1.538)	(6.131)		(2.585)	(2.585)		(-8.135)
$\hat{\alpha}_3$	**0.019**	0.013	**-0.428**	0.000	0.005		
	(2.241)	(0.399)	(-4.812)		(0.145)		

These relationships, when isolating on the l.h.s. the normalized variable and moving the other variables on the r.h.s. of the relations, are broadly consistent with the theoretical model:

- 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^{GER} also has the expected sign (see also fn. 11) (apart from the equation for output where it is constrained to zero), but is significant only in the second relation
- the proxy for the credit-risk premium ρ_t proves consistent with the CCC 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 discussion in Tamborini (2009)).

From the estimated α coefficients we can infer the adjustment dynamics. It is interesting to note that inflation is equilibrium correcting in the first and the third relation; real wages are equilibrium correcting only in the second relation; output is adjusting in the

[18] The degrees of freedom of the *LR* test correspond to the weak exogeneity restrictions for the variable k_t.

equilibrium errors of the first and the third relation, but increasing in the equilibrium errors of the second relation; the credit-risk premium is equilibrium correcting in the first and in the second relation. Overall behaviour, however, is stable because of the stronger adjusting effects. The zero coefficients on the inter-bank rate are due to the weak exogeneity assumption. These results yield important information on the stability of the system that we shall exploit later.

3.2.2 Italy 1999 - 2007

On analysing the data set over the period 1999 – 2007, with $\mathbf{z}'_t = [Lib_t]$, we have found evidence of I(2) characteristics in particular for inflation, the inter-bank rate and the LIBOR. Following Juselius (2006), we have introduced significant breaks in the linear trend specified at 2001:9 and 2004:4[19]. The results of misspecification tests, calculated on residuals, have assessed the adequacy of the VAR specification with $n = 1$ lag in (3.1)[20].

The search procedure for the number of structural cointegrating relations, based on the trace test and the additional statistical information as in the preceding sub-period, has shown that $r = 3$ was still acceptable[21]. Below we report the final structural long-run relations (*t*-statistics in parentheses), together with the value of the *LR* test for the over-identifying restrictions, and the estimated adjustment dynamics.

In the first place, the inter-bank rate k_t always has the expected signs and significant coefficients on inflation, real wage rate and output. This is not the case of the credit-risk premium ρ_t, which is significant with the right sign for output, but with the wrong sign for inflation and real wage, and for the variable Lib_t, which is significant with the right sign for real wage, but insignificant for inflation.

The deterministic components in the first and in the third cointegrating relation are broken trends. In the second relation there is no evidence of any significant broken trend; therefore, an unrestricted constant characterizes this relation. From the estimated α coefficients we can infer the equilibrium correcting behaviour of inflation, real wages and output, while the inter-bank rate is increasing in the equilibrium error of the first relation and adjusting in the others, and risk is equilibrium correcting only in the first.

[19] The statistical significance of the breaks has resulted from the variable exclusion testing provided as an automated test procedure in RATS.

[20] In order to obtain residuals close to Normality we have introduced two permanent intervention dummies for 1999:4 and 1999:10. The misspecification tests for the unrestricted VAR(1) model with dummies have taken the following values. The *LM*(1) test for first order autocorrelation, asymptotically distributed as a χ^2_{25} variable, is equal to 24.700 with a *p-value* of 0.479. As regards Normality of residuals, the test, asymptotically distributed as a χ^2_{10} variable, is equal to 18.732, with a *p-value* of 0.044.

[21] The trace test has given $r = 3$ at a significance level of 1% and $r = 4$ at 10%, but with four values of the eigenvalues, corresponding to $r = 0, 1, 2, 3$, three quite high and one smaller. The modulus of the largest unrestricted characteristic root of the companion matrix is 0.854 for $r = 2$, 0.829 for $r = 3$, and 0.853 for $r = 4$, indicating $r = 3$ as preferable.

Table 2. Italy: the estimated identified long-run structure and the estimated adjustment coefficients over the sub-period 1999-2007 (bold coefficients denote significance at 10%)

	π_{t+12}	w_t	q_{t+12}	k_t	ρ_t	Lib_t	$D_t01{:}9$	$D_t04{:}4$	$trend_t$
				$\chi^2_2 = 0.696, \quad p\text{-}value = 0.706$					
$\hat{\beta}_1$	**1.000**	-	-	**0.881**	**-0.316**	-0.103	**0.001**	**-0.001**	**-0.0003**
				(6.008)	(-8.226)	(-1.286)	(5.994)	(-4.878)	(-2.134)
$\hat{\alpha}_1$	**-0.112**	-0.144	**-0.641**	**0.066**	**0.502**				
	(-3.919)	(-1.524)	(-3.729)	(4.591)	(6.021)				
$\hat{\beta}_2$	**-1.267**	**1.000**	-	**1.837**	**-0.657**	**0.722**	-	-	-
	(-1.907)			(4.988)	(-3.563)	(4.160)			
$\hat{\alpha}_2$	**0.046**	**-0.240**	-0.101	**-0.015**	0.012				
	(3.264)	(-5.141)	(-1.183)	(-2.087)	(0.293)				
$\hat{\beta}_3$	**-1.655**	**0.443**	**1.000**	**1.763**	**1.150**	-	-	**-0.003**	**0.001**
	(-1.770)	(1.707)		(2.578)	(5.447)			(-7.093)	(3.425)
$\hat{\alpha}_3$	0.013	**-0.133**	**-0.314**	**-0.027**	0.035				
	(0.977)	(-3.018)	(-3.924)	(-4.112)	(0.895)				

3.2.3 Germany 1990 - 1998

Plugging $\mathbf{z}'_t = [Lib_t]$ into (1), the search for structural cointegrating relations for Germany has followed the same procedure as for Italy, confirming $r = 3$ as cointegration rank[22]. The number $n = 2$ of lags in (1) has been determined on the basis of misspecification tests[23]. Below we report the final just-identified structural long-run relations (*t*-statistics in parentheses) together with the value of the *LR* test[24], and the estimated adjustment dynamics.

The overall picture appears even more consistent with the CCC hypothesis than in the case of Italy. Both the inter-bank rate k_t ,and the credit-risk premium ρ_t always have the expected, significant coefficients in the inflation, real wage rate and output relations. The coefficient associated with the variable LIBOR is significant and shows the expected sign in each relation in which it is not restricted[25].

[22] The trace test has given $r = 2$ at a significance level of 5%, but with four eigenvalues, corresponding to $r = 0, 1, 2, 3$, quite close to each other and one very small value. The modulus of the largest unrestricted characteristic root of the companion matrix is 0.736 for $r = 2$, 0.819 for $r = 3$, and 0.988 for $r = 4$, indicating $r = 3$ as still reasonable among the alternatives.

[23] We have introduced two permanent intervention dummies for 1991:10 and 1996:1 into the German data-set. The values of the misspecification tests for the unrestricted VAR(2) model with dummies are the following: the *LM*(1) test for first order autocorrelation is equal to 33.576 with a *p-value* of 0.117; the test for residual Normality is equal to 26.824, with a *p-value* of 0.003. Normality is rejected due to excess kurtosis in inflation. Because VAR estimates are more sensitive to deviations from Normality due to skewness than to excess kurtosis, we have considered the chosen model to be well specified.

[24] The degrees of freedom of the *LR* test correspond to the weak exogeneity restrictions for the variable k_t.

[25] Given that Germany had no explicit non-EMS exchange-rate target, Lib_t, unlike k_t^{GER} for Italy, should take the same sign as the domestic rate.

As regards the adjustment dynamics, the α coefficients show that inflation is not significantly equilibrium error correcting; real wages are equilibrium error correcting in the second and third relation but increasing in the first relation with a stronger adjusting overall effect; output is equilibrium error correcting in the first relation but increasing in the second relation; and risk is equilibrium correcting in the third relation. The prevailing equilibrium correcting behaviours make the system stable anyway, as implied by its characteristic roots.

Table 3. Germany: the estimated identified long-run structure and the estimated adjustment coefficients over the sub-period 1990-1998 (bold coefficients denote significance at 10%)

	π_{t+12}	w_t	q_{t+12}	k_t	ρ_t	Lib_t	$trend_t$
			$\chi^2_3 = 3.635,\ p\text{-}value = 0.304$				
$\hat{\beta}_1$	**1.000**	-	-	**0.644**	**0.765**	**0.364**	**0.001**
				(2.988)	(3.566)	(2.850)	(9.443)
$\hat{\alpha}_1$	-0.101	**6.347**	**-3.294**	0.000	-0.517		
	(-0.345)	(4.552)	(-2.506)		(-1.602)		
$\hat{\beta}_2$	**13.753**	**1.000**	-	**8.987**	**8. 987**	**6.424**	**0.005**
	(41.969)			(3.106)	(3.106)	(3.807)	(6.762)
$\hat{\alpha}_2$	-0. 008	**-0.515**	**0.239**	0.000	0.025		
	(-0.384)	(-5.162)	(2.538)		(1.101)		
$\hat{\beta}_3$	**-5.663**	0.329	**1.000**	**2.825**	**2.825**	-	**-0.004**
	(-8.138)	(1.259)		(2.693)	(2.693)		(-3.803)
$\hat{\alpha}_3$	-0.009	**-0.163**	-0.048	0.000	**-0.029**		
	(-0.850)	(-3.378)	(-1.061)		(-2.562)		

3.2.4 Germany 1999 - 2007

Also for Germany, we have found evidence of I(2) characteristics for the inter-bank rate and the LIBOR in the EMU subperiod. We have therefore specified a significant break in the linear trend at 2005:9. Moreover, we have introduced a mean-shift dummy variable starting at 2003/I[26], restricted to the cointegrating relations. The results of misspecification tests have assessed the adequacy of the VAR specification with n = 1 lag in (3.1)[27] and the choice of r = 3 cointegrating relations[28]. Below we report the final structural long-run

[26] In January 2003 the Bundesbank's earlier survey of lending and deposit rates was discontinued and replaced with the new harmonised MFI interest rate statistics. Since the two sets of statistics differ in their methodology, we have introduced a dummy for the change from the lending rate series SU0004 to lending rate series SUD123.

[27] In order to obtain residuals close to Normality we have introduced one permanent intervention dummy for 1999:9. The values of the misspecification tests for the unrestricted VAR(1) model with dummies are following: the *LM*(1) test for first order autocorrelationis equal to 27.466 with a *p-value* of 0.333; the test for residuals' Normality is equal to 10.175, with a *p-value* of 0.425.

[28] The trace test has indicated r = 4 at a significance level of 1%, but three values of the eigenvalues, corresponding to r = 0, 1, 2, one quite high and two smaller. The modulus of the largest unrestricted characteristic root of the companion matrixis 0.906 for r = 2, 0.911 for r = 3, and 0.935 for r = 4, indicating r

relations (*t*-statistics in parentheses), together with the value of the *LR* test for the over-identifying restrictions, and the estimated adjustment dynamics.

Table 4. Germany: the estimated identified long-run structure and the estimated adjustment coefficients over the sub-period 1999-2007 (bold coefficients denote significance at 10%)

	π_{t+12}	w_t	q_{t+12}	k_t	ρ_t	Lib_t	$D_s03{:}1$	$D_t05{:}9$	$trend_t$
				$\chi^2_2 = 1.273, \quad p\text{-}value = 0.529$					
$\hat{\beta}_1$	**1.000**	-	-	**0.552**	**0.612**	**0.143**	**0.009**	-	-
				(2.876)	(3.053)	(2.076)	(1.695)		
$\hat{\alpha}_1$	**-0.154**	-0.325	**-0.889**	**0.141**	**-0.145**				
	(-2.706)	(-1.227)	(-4.504)	(5.738)	(-2.507)				
$\hat{\beta}_2$	**-2.983**	**1.000**	-	0.145	**-1.046**	**1.329**	-	**0.0043**	**0.0004**
	(-5.052)			(0.397)	(-3.240)	(7.163)		(4.424)	(2.246)
$\hat{\alpha}_2$	**0.077**	**-0.263**	**0.501**	**0.022**	-0.010				
	(2.680)	(-1.980)	(5.053)	(1.821)	(-0.354)				
$\hat{\beta}_3$	**-3.044**	**1.151**	**1.000**	**3.248**	-0.063	-	**0.041**	-	**-0.0015**
	(-6.042)	(13.662)		(6.129)	(-0.138)		(3.000)		(-10.12)
$\hat{\alpha}_3$	-0.034	**-0.365**	**-0.645**	-0.004	0.010				
	(-1.368)	(-3.121)	(-7.393)	(-0.340)	(0.395)				

The results are in line with previous findings, in that the inter-bank rate k_t has the expected coefficients in every relation. However it is not significant in the real wage relation, which weakens the evidence on the supply-side channel of monetary policy for Germany in the EMU. Lib_t has the expected significant coefficients in each relation in which it is not restricted. The credit-risk premium ρ_t has the expected, significant coefficient in the inflation relation, significant but with the wrong sign in the real wage relation, and not significant in the output relation.

The mean-shift dummy significantly affects the first relation and the third relation. The deterministic component in the second relation is a broken trend, and in the third relation it is a linear trend.

To be noted from the estimated α coefficients is the overall equilibrium correcting behaviour, with some non equilibrium correction of the interbank rate.

3.2.5 Preliminary comment

In our view, the evidence based on cointegration analysis in the previous section is consistent with the CCC hypothesis set out in section 3 according to which a set of long-run equilibrium inverse relationships is detectable between the inter-bank rate, on the one hand, and inflation, output and real wages, on the other. In textbook graphical terms, shifts in the values of k_t displace *both* the "AD curve" *and* the long-run "AS curve" in the output-

= 3 as a reasonable choice among the alternatives.

inflation space, leaving a net effect with negative sign on both inflation and output. Though comparable with the empirical results produced by Chowdhury et al. (2006) and De Fiore and Tristani (2009), our CCC test may be regarded as stronger in that it does not hinge (necessarily) on the ancillary hypotheses of nominal or real rigidities. We can also conclude that the CCC is detectable in both Italy and Germany, and that the advent of the euro has not created a major structural shift (with weaker evidence for the role of real wages in Germany in the EMU sub-period).

4. Is the inter-bank rate a control variable in the system, and does "one size fit all"?

Information drawn from the foregoing econometric analysis is relevant for the transmission of monetary policy in the euro area, and this section is devoted to this issue. The close relationship between the control of policy rates and the inter-bank rates within the "corridor" is the cornerstone of monetary policy in the euro area. Reliance on this relationship, however, raises two well-known questions: is the inter-bank rate a true control variable, and can one single control rule fit all different controlled systems?

We have performed a rigorous statistical analysis of these debated issues on the basis of the CCC transmission mechanism by drawing on Johansen and Juselius (2003) extension of the CVAR methodology to policy control analysis, obviously over the EMU sub-period. Their approach 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. Second, 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. Third, given controllability, *a control rule* specifies interventions on the instrument conditional on the observed state of the target variable relative to the target. The first step hinges on the results of our cointegration analysis. The second and third steps will be expounded in this section.

4.1. The "pushing forces" of the system

Johansen and Juselius (2003) policy control analysis starts from the well-known distinction in the CVAR methodology between the "pulling forces" and the "pushing forces" of the system. The "pulling forces" have been identified in the cointegration analysis. Information on the "pushing forces" can be gained from the inverted CVAR (1) yielding the vector moving average (VMA) representation for $\Delta\mathbf{y}_t$. Rewriting it in terms of the levels of the variables by recursive substitution and focusing the attention only on the common stochastic and deterministic trends characterizing it, we get the following VMA representation for $\mathbf{y}_t$:

$$(2) \qquad \mathbf{y}_t = \mathbf{C}\sum_{i=1}^{t}\boldsymbol{\varepsilon}_i + \mathbf{C}\boldsymbol{\mu}_0 t + \ldots$$

where the ($p\times p$) matrix $\mathbf{C}$, given the existence of r = 3 cointegrating vectors, has reduced rank ($p-r$) = 2. Note that it can be written as $\mathbf{C} = \tilde{\boldsymbol{\beta}}_{\mathrm{y}\perp}\boldsymbol{\alpha}'_{\perp}$, a decomposition similar to $\boldsymbol{\Pi} = \boldsymbol{\alpha\beta}'$, characterizing the CVAR (1)[29].

The matrix **C** plays an important role: its rank corresponds to the number of common stochastic trends, and its elements convey information about the long-run impact of cumulated shocks to the system variables. In other words, the matrix **C** allows us to determine which empirical shocks have *permanent* effects on the system variables. An eligible policy instrument should therefore be identifiable as a "pushing force" of the system by way of matrix **C**. Given the assumption of exogeneity of $\mathbf{z}_t$ (Ericsson et al. (1998), Definition 3) and its non controllability, we have focused our attention on the effects of unanticipated shocks to the system variables $\mathbf{y}_t$. In this way, it has been possible to evaluate the effects on inflation and output of unexpected changes in the policy action. Here we only report the results, whereas the procedure and statistics are expounded in a separate appendix.The results as to identification of common stochastic trends are quite similar and stable across the two countries and the two sub-periods. Hence we can summarize them as follows.

- The first common stochastic trend in all estimated models is given either by the cumulated empirical shocks to k_t (in the sub-period 1986 - 1998, when this variable is weakly exogenous), or by the cumulated empirical shocks to k_t, with a significant contribution from shocks to ρ_t (in the EMU sub-period): it can be labelled as a nominal stochastic trend.
- This stochastic trend has significantly and negatively affected inflation, real wages and production in the long run, consistently with our model, in both countries in both periods.
- In the first sub-period, the second common stochastic trend is primarily associated with the credit-risk premium ρ_t in both countries, with a borderline contribution of real wages and inflation in Italy, and of output in Germany. In the second sub-period, primary contribution comes from inflation and real wages in both countries.

It can be seen from the corresponding rows of the estimated matrix $\hat{\mathbf{C}}$, reported in Table 5 for Italy and in Table 6 for Germany, that the target variables can in fact be controlled by the inter-bank rate in both countries.

Table 5. Italy: the long-run impact on inflation and output of unanticipated shocks to the system over the sub-period 1999-2007 (*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}	**0.307**	**0.076**	0.028	**-0.332**	**0.169**

[29] In this decomposition, $\tilde{\boldsymbol{\beta}}_{\mathrm{y}\perp}$ is given by $\boldsymbol{\beta}_{\mathrm{y}\perp}(\boldsymbol{\alpha}'_{\perp}\boldsymbol{\Gamma}_{\mathrm{y}}\boldsymbol{\beta}_{\mathrm{y}\perp})^{-1}$, where $\boldsymbol{\beta}_{\mathrm{y}\perp}$ and $\boldsymbol{\alpha}_{\perp}$ are ($p\times(p\text{-}r)$) matrices orthogonal to $\boldsymbol{\beta}_{\mathrm{y}}$ and $\boldsymbol{\alpha}$, respectively, and $\boldsymbol{\Gamma}_{\mathrm{y}} = -\mathbf{I}_p + \sum_{i=1}^{n-1}\boldsymbol{\Gamma}_{\mathrm{y}i}$.

	(2.292)	(2.427)	(0.970)	(-2.739)	(3.991)
q_{t+12}	**-2.290**	-0.359	**0.471**	**-4.089**	0.522
	(-2.494)	(-1.674)	(2.358)	(-4.917)	(1.743)

Table 6. Germany: the long-run impact on inflation and output of unanticipated shocks to the system over the sub-period 1999-2007 (*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}	0.103	-0.007	-0.007	**-0.529**	**-0.561**
	(0.538)	(-0.264)	(-0.251)	(-5.552)	(-4.843)
q_{t+12}	-2.416	-0.351	0.313	**-3.931**	**-2.371**
	(-1.462)	(-1.538)	(1.242)	(-4.790)	(-2.376)

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 with a 12-month time lag. Notably, the size of coefficients is consistent with the usual pattern of larger adjustments in quantities than in prices. Johansen and Juselius (2003) obtain similar results in the case of the United States, though with the anomaly of a *positive* sign. They conjecture that this anomaly may be due to the absence of the supply side in their model; our result suggests that their conjecture may be right. Finally, in the case of Germany, though not of Italy, shocks to the credit risk premium have effects similar to those of the inter-bank rate, in line with the theoretical model. This indicates that credit risk shocks displace credit supply, affect bank interest rates and are transmitted to the real economy by way of the CCC. More in-depth analyses and accurate measures of credit risk with disaggregate data in the euro area, such as Altunbas et al. (2009), yield broadly the same picture.

4.2. Inflation control by means of the inter-bank rate

At this point, CVAR econometric analyses usually exploit the additional information contained in the estimated α coefficient to gauge how the system adjusts to shocks dynamically – the so-called impulse responses of the system. Here instead we present the result of the Johansen-Juselius third step of policy-control analysis, which essentially makes use of the same statistical information as an input of a policy control rule aimed at a target variable. Specifically, we have written a control rule using the inter-bank rate as an instrument to directly control inflation[30]. Then we have performed a simulation of the effects of this control rule on the inflation rate and the industrial production index by way of the estimated CCC model for both Italy and Germany in the EMU sub-period fed by the model's residuals.

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

[30] Hence we have simulated a pure inflation targeting regime, rather than a common Taylor rule where output is also a 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 **y** is therefore the following:

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

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

$$\mathbf{y}_t^{ctr} = \mathbf{y}_t^{new} + \mathbf{v}_t \, .$$

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.

For each country we report graphs of

- 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% (the baseline is zero, so positive values denote contractionary impulses and negative values denote expansionary impulses)
- the observed and the new inflation rate
- the observed and the new output
- the resulting inflation-output scatter (the statistical equivalent of the AS curve) both observed and new

Table 7 provides the appropriate summary statistics allowing for more precise quantitative assessment.

Table 7. Summary statistics of the observed and simulated time series of output and inflation in the EMU sub-period

	Italy			
	Output-observed	Output-simulated	Inflation-observed	Inflation-simulated
Mean	4.58	4.61	2.29%	1.97%
St. Deviation	0.02	0.06	0.38	0.24
St. Dev. from 2%			0.49	0.24
	Germany			
	Output-observed	Output-simulated	Inflation-observed	Inflation-simulated
Mean	4.65	4.68	1.62%	1.86%
St. Deviation	0.06	0.06	0.44	0.50
St. Dev. from 2%			0.58	0.52

The simulation results serve two purposes. The first, and most important, one is assessment of the hypothesis that the inter-bank rate is a means to control inflation with the ensuing long-run consequences on output and inflation. The second is that the simulated paths of inflation and output may also provide a benchmark against which their actual paths – a result of the single ECB policy on the respective inter-bank rates – can be assessed. If our estimated CCC economies are good statistical representations of the true ones, differences between estimated and observed variables can be traced back to three factors: 1) a *country-specific factor* (the ECB does not make use of country-specific state information), 2) a *model specification factor* (the ECB may not consider the CCC effects at the Union-area level either), 3) a *policy specification factor* (the ECB does not follow the same rule as our control rule).

Firstly, Figures 5 and 6 show that the interventions on the observed inter-bank rate dictated by our control rule of the CCC model are neither large nor systematic in both countries. This evidence can be read as mutual consistency between the hypothesis that the observed inter-bank rates are indeed driven by at least a component of 2% inflation-targeting and that our control rule fits observed inter-bank rates reasonably well (see also Johansen and Juselius (2003) on the US case). However, there are also a few differences between Italy and Germany.

Figure 5 points up two spells of time in which the control of the CCC model determines notable interventions on Italy's inter-bank rate path. First, from early 1999 to end 2001 we see a sequence from negative to positive interventions. This indicates a *reversal of the policy sequence* (first a lower then a higher rate) with respect to the observed one (first a higher then a lower rate). Second, from mid-2003 onwards we see a sequence of negative interventions. This, from 2004, indicates a *reversal of the policy stance* (from neutral to expansionary) with respect to the observed one (from neutral to contractionary). As to Germany, Figure 6 presents a similar pattern to Italy’s between 1999 and 2001. We can then observe a more systematic alignment with the observed inter-bank rate, apart from a few positive and negative spikes – probably due to local shocks.

Secondly, Figure 7 and Figure 8 exemplify our foregoing econometric result that the inter-bank rate in the CCC model is indeed a means to make the inflation process stationary around the 2% target in both countries. Statistics in Table 7 indicate that the simulated inflation control is quite effective in terms of standard deviation from target in both countries, remarkably so for Italy. In the case of Italy, a phase of significant downward deviations from the observed inflation path is also observable, which is broadly consistent with the simulated inter-bank rate *vis-à-vis* the actual one (from 1999 to 2003 the policy reversal generated by the control rule is consistent with inflation lower than observed). As in the case of the inter-bank rate, the path of the simulated inflation process in Germany is closer to the observed one, with the upward spikes in the simulated inter-bank rate in 2002-2003 corresponding to upward spikes in the simulated inflation rate.

Thirdly, Figure 9 and Figure 10 illustrate the effect on output of controlling the inter-bank rate according to the CCC hypothesis. Again, whereas Germany displays a path of the simulated output closer to the observed one – though systematically higher – Italy presents quite a different picture. The (moderately) different paths of the controlled inter-bank and inflation rates produced by the CCC model with respect to the real data seem to produce major systematic differences in the output path. A clear intertemporal relocation emerges with lower output in the early years of the EMU against higher values from 2002 onwards.

Interestingly, the simulated path of output for Italy resembles the one observed for Germany. A suggested interpretation may be that Italy's puzzling poor industrial performance after 2002 may not be entirely unrelated with the ECB monetary policy not fitting both countries[31].

Finally, Figures 11 and 12 provide visual evidence of our further key econometric finding that inter-bank rate shocks exert long-run effects on the stochastic paths of inflation and output represented in the usual inflation-output space. Interpolation lines have been added for the reader's convenience. The control of inflation is obtained by structurally shifting the inflation-output equilibria. The new equilibrium realizations are located around a flat locus corresponding to 2% inflation. This is not, however, a consequence of sticky prices but of pure targeting of the 2% inflation rate and the CCC joint effects on aggregate demand and supply.

Comparison of the simulated inflation-output realizations with the ones observed prompts further considerations. The simulated inflation-output loci in both countries appear flatter while the dispersion of output is widened. For the reasons discussed above, this phenomenon is more pronounced for Italy, where the left-hand side of the diagram reflects the 1999-01 downward shift of the output process, and the right-hand side its 2002-07 upward shift, in the simulated economy. From the statistics in Table 7, we can see that the observed and simulated time series of Germany’s output and inflation are virtually undistinguishable, whereas in the case of Italy output variability is indeed higher and inflation variability is lower in the simulated time series. An interpretation may be that, according to a widespread opinion, the ECB follows flexible inflation targeting, which also includes output in the control rule. It is indeed an expected consequence of this policy strategy that output variability is reduced at the expense of inflation variability (Taylor (1998)).

5 Conclusions

In this paper we have reported an empirical analysis of the CCC hypothesis of monetary policy transmission. This hypothesis combines bank credit supply, as a means whereby monetary policy affects economic activity ("credit channel"), with interest rates on loans as a cost to firms ("cost channel"). The thrust of the model is that firms' reliance on bank loans makes both aggregate demand and aggregate supply dependent on monetary policy to the extent that this affects bank interest rates. These joint effects yield a pattern of relationships consistent with the empirical regularities associated with monetary policy interventions, with no recourse to additional non-competitive hypotheses. Moreover, the presumption arises that the CCC may also have permanent, rather than transitory, effects on

[31] Note that the 2006-07 inflationary phase took off much earlier in Germany than in Italy both in the real data and in the simulated ones. The timing of the rise in the observed inter-bank rates follows that of Germany's inflation. Accordingly, the simulated control rule yields no interventions on the inter-bank rate for Germany but downward impulses for Italy.

real variables.

The statistical methodology adopted has enabled us to apply a single integrated framework to the identification of both 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. Our econometric analysis supports the CCC hypothesis for Italy and Germany, two typical bank-based economies, and shows that the CCC is detectable in the pre-EMU as well as in the EMU set-up. By way of the CCC transmission mechanism, the inter-bank rate is a "pushing force" of the system, with the expected negative correlation, of the long-run stochastic equilibrium paths of the real wage rate, output and inflation around which transitory dynamics takes place.

Secondly, by exploiting the properties of Johansen-Juselius 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 in the EMU set-up. By simulating a pure 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.

Finally, upon comparing our simulated data with those observed under the ECB control, we can conclude that the hypotheses that the ECB does control inter-bank interest rates in view of a 2% inflation target, and that our estimated control model fits the observed inter-bank rates, are mutually supportive. On the other hand, a few significant discrepancies between simulated and observed data of inflation and industrial production also support the hypothesis that the ECB follows flexible, rather than pure, inflation targeting. This is welcome because it mitigates the variability of output that may be induced by a central bank following pure inflation targeting and/or ignoring the CCC long-run effects on output. Yet our exercise also provides some evidence on the "one size does not fit all" problem that may have penalized Italy in the second half of the euro decade.

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.

References

Altunbas Y., Gambacorta L., Marques-Ibanez D. (2009), "Bank Risk and Monetary Policy", Working Paper Series n. 1075, European Central Bank.

Angeloni I. (1994), "Strumenti e obiettivi operativi della Banca d'Italia: verso un modello europeo", *Note Economiche*, 14, 128-161.

Angeloni I., Ehrmann M. (2003), "Monetary Policy Trnsmission in the Euro Area: Any Changes After EMU?", Working Paper Series n. 240, European Central Bank.

Angeloni I., Kashyap A., Mojon B. (2003), *Monetary Policy Transmission in the Euro Area*, Cambridge, Canbridge University Press.

Banca d'Italia (1997), "I canali di trasmissione della politica monetaria nel modello econometrico trimestrale della Banca d'Italia", Temi di Discussione, n. 316, Banca d'Italia.

Barth M. J., Ramey V. (2001), "The Cost Channel of Monetary Transmission", in Bernanke B., Rogoff K. (eds.), *NBER Macroeconomics Annual 2001*, Cambridge Mass., MIT Press.

Bernanke B., Gertler M. (1995), "Inside the Black Box: the Credit Channel of Monetary Policy Transmission", *Journal of Economic Perspectives*, 9, 27-48.

Chiades P., Gambacorta L. (2004). "The Bernanke and Blinder Model in an Open Economy: the Case of Italy", *The German Economic Review*, 5, pp. 1-34

Chowdhury I, Hoffmann M., Schabert A. (2006), "Inflation Dynamics and the Cost Channel of Monetary Transmission", *European Economic Review*, 50, pp. 995-1016.

Christiano L.J., Eichenbaum M. (1992), "Liquidity Effects and the Monetary Transmission Mechanism", *American Economic Review*, 80, 703-725.

Christiano L.J, Eichenbaum M., Evans C.L. (1997), "Sticky Price and Limited Participation Models of Money: A Comparison", *European Economic Review*, 41, 1201-1249.

Christiano L.J., Motto R., Rostagno M. (2007a), "Financial Factors in Business Cycles", mimeo.

Christiano L. J., Motto R., Rostagno M. (2007b), "Two Reasons Why Money and Credit May Be Useful in Monetary Policy", Paper presented at the 4th ECB Central Banking Conference, Frankfurt, November 9-10, 2006.

Crockett A. (2003), "International Standard Setting in Financial Supervision", Institute of Economic Affairs Lecture, Cass Business School, mimeo.

De Fiore F., Tristani O. (2009), "Optimal Monetary Policy in a Model of the Credit Channel", Working Paper Series, European Central Bank, n. 1043.

De Grauwe P. (1992), *The Economics of Monetary Integration*, Oxford, Oxford University Press.

Dornbush R., Favero C.A., Giavazzi F., (1998) "Immediate Challenges for the European Central Bank", *Economic Policy*, 26, pp.15-64.

ECB (2008), *10th Anniversary of the ECB, 1998-2008*, Frankfurt, European Central Bank.

Ericsson N.R., Hendry D.F., Mizon G.E. (1998), "Exogeneity, Cointegration, and Economic Policy Analysis", *Journal of Business and Economic Statistics*, 16, 370-387.

Fiorentini R., Tamborini R. (2001) , "The Monetary Transmission Mechanism in Italy: The Credit Channel and a Missing Ring", *Giornale degli Economisti*, 61, 1-42.

Fiorentini R., Tamborini R. (2002) "Monetary Policy, Credit and Aggregate Supply", *Economic Notes*, 31, 1-48.

Gaiotti E., Secchi A. (2004), "Are Monetary Policy Actions a Supply Shock? An Exploration Using Italian Firm-Level Data", Banca d'Italia, mimeo.

Gertler M., Gilchrist S. (1993), "The Role of Credit Market Imperfections in the Monetary Transmission Mechanism: Arguments and Evidence", *Scandinavian Journal of Economics*, 93, 43-64.

Greenwald B.C., Stiglitz J.E. (1988), "Imperfect Information, Finance Constraints and Business Fluctuations", in Kohn, M., Tsiang S.C. (eds.), *Finance Constraints, Expectations, and Macroeconomics,* Oxford, Claredon Press, 103-140.

Greenwald B.C., Stiglitz J.E. (1993), "Financial Market Imperfections and Business Cycles", *Quarterly Journal of Economics*, 108, 77-113.

Greenwald, B.C., Stiglitz, J.E. (2003), *Towards a New Paradigm in Monetary Economics*, Cambridge, Cambridge University Press.

Johansen S. (1996), "*Likelihood Based Inference in Cointegrated Vector Auto-Regressive Models*", Oxford, Oxford University Press.

Johansen S., Juselius K. (2003), "*Controlling Inflation in a Cointegrated Vector Autoregressive Model with an Application to US Data*", Department of Applied Mathematics and Statistics and Economics Department, University of Copenhagen.

Juselius K. (2006), *The Cointegrated VAR Model. Methodology and Applications,* Oxford, Oxford University Press.

Leijonhufvud A. (2007), "Monetary and Financial Stability", CEPR Policy Insight, n. 14

Pfajfar D., Santoro E. (2007), *Credit Market Distortions, Asset Prices and Monetary Policy*, University of Cambridge (UK), mimeo.

Ravenna F., Walsh C.E. (2006), "Optimal Monetary Policy With the Cost Channel", *Journal of Monetary Economics*, 53, pp. 199-216.

Sims C. (1992), "Interpreting the Macroeconomic Time Series Facts: The Effects of Monetary Policy", *European Economic Review*, 36, pp. 975-1000.

Tamborini R. (2009), "The “Credit–Cost Channel” of Monetary Policy. A Theoretical Assessment". *Economics E-Journal*, 3, 2009, http://www.economics-ejournal.org/economics/journalarticles/2009-13

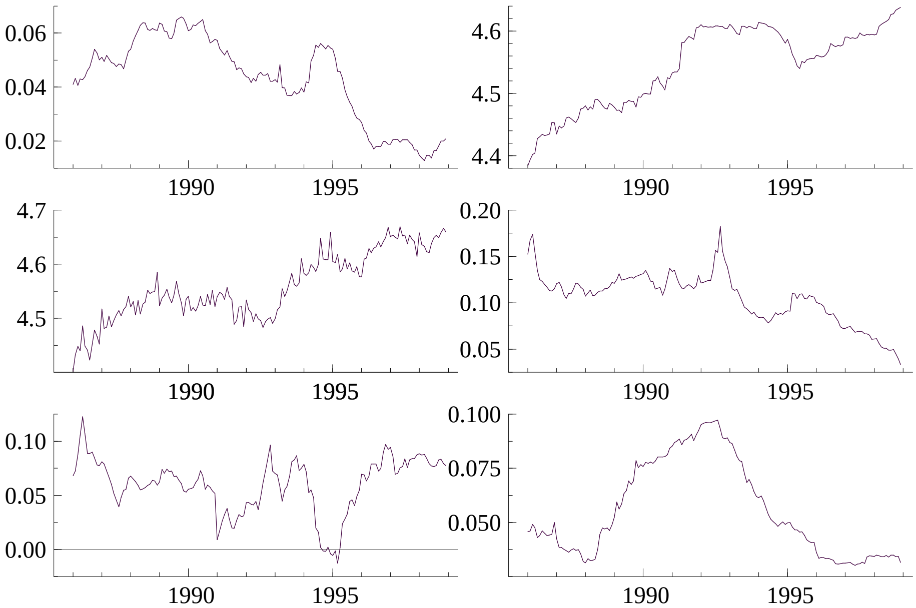
Taylor J.B. (1998), *Inflation, Unemployment and Monetary Policy*, Cambridge (Mass), MIT Press.

Trautwein H.M. (2000), "The Credit View, Old and New", *Journal of Economic Surveys*, 14, pp.155-136.

Visco, I. (1995), "Inflation, Inflation Targeting and Monetary Policy: Notes for Discussion on the Italian Experience" in Leiderman, L. and L.E.O. Svensson, eds., *Inflation Targets*, London, CEPR.

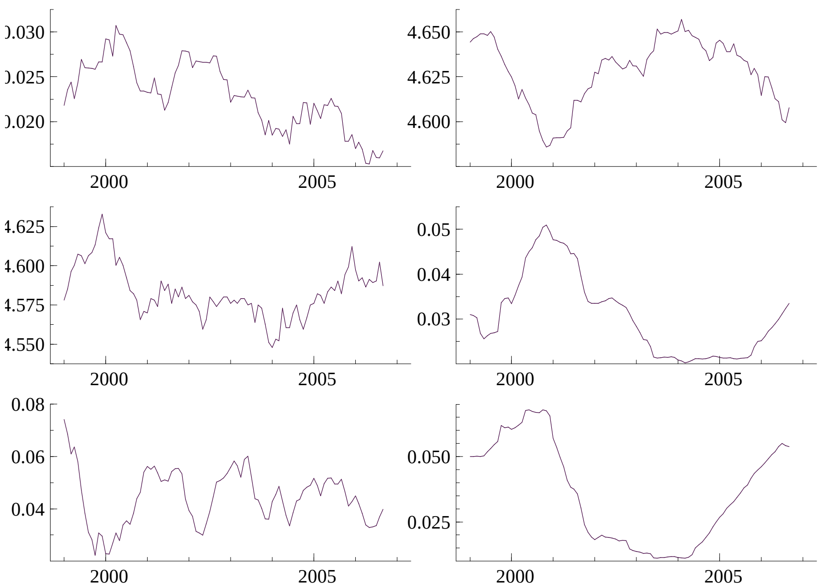
Woodford M. (2003), *Interest and Prices”*, Princeton, Princeton University Press.

Figure 1 Italy, sub-period 1986-1998, 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]



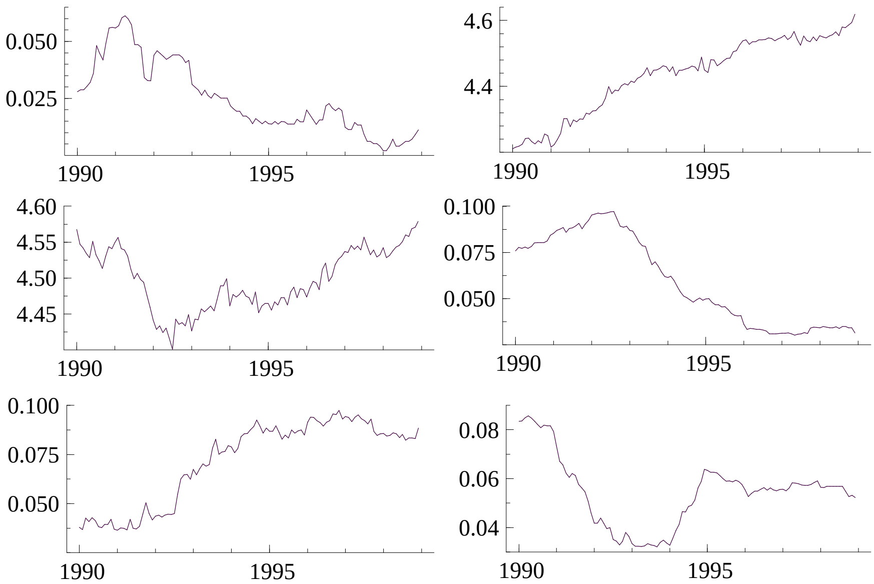
Sources: [1]IMF, *International Financial Statistics*; [2]Bank of Italy, *Monetary Statistics*

Figure 2. Italy, sub-period 1999-2007, 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[1]; credit risk premium;LIBOR US Dollar rate[2]



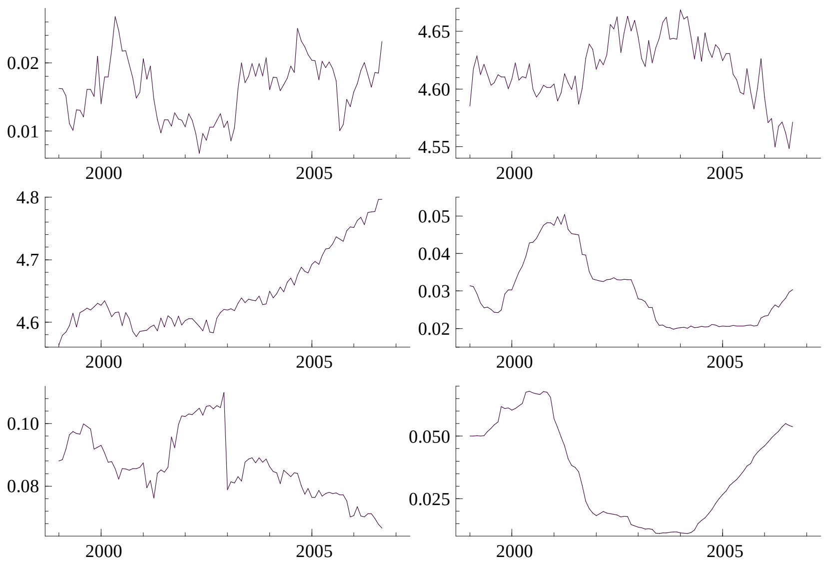
Sources: [1]IMF, International Financial Statistics; [2]Eurostat

Figure 3. Germany, sub-period 1990-1998, 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 4. Germany, sub-period 1999-2007, 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[1]; credit risk premium[2]; LIBOR US Dollar[3].



Sources: [1]IFS, *International Financial Statistics*; [2]Bundesbank; [3]Eurostat**.**

Figure 5. Italy: representation of the inter-bank rate (solid line) and the derived intervention (dotted line) to make inflation stationary around 2%.

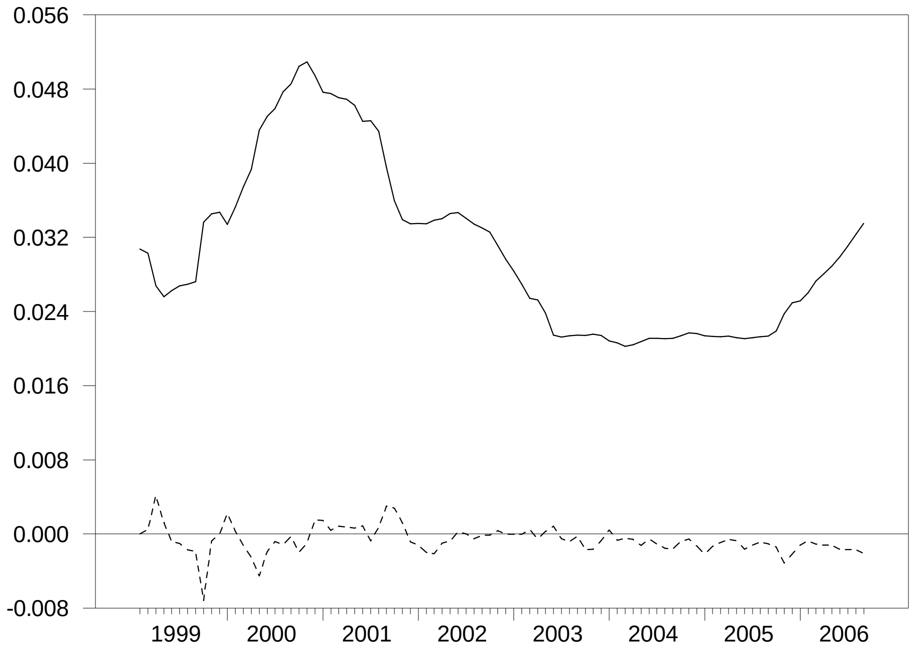


Figure 6. Germany: representation of the inter-bank rate (solid line**) and the derived intervention** (dotted line) **to make inflation stationary around 2%.**

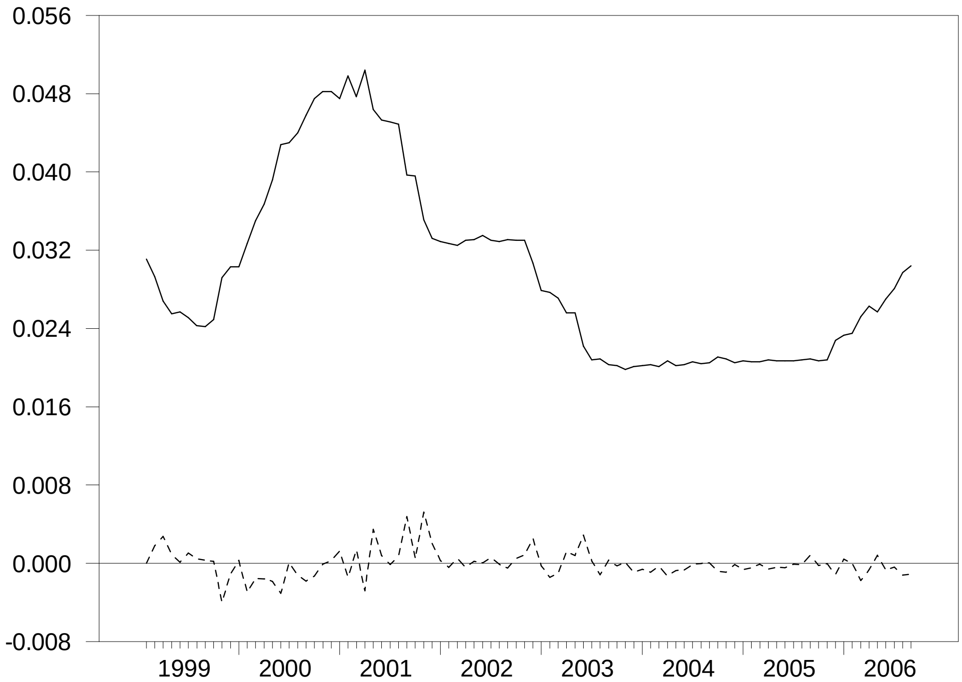


Figure 7. Italy: observed (solid line) **and "new" inflation** (dotted line)

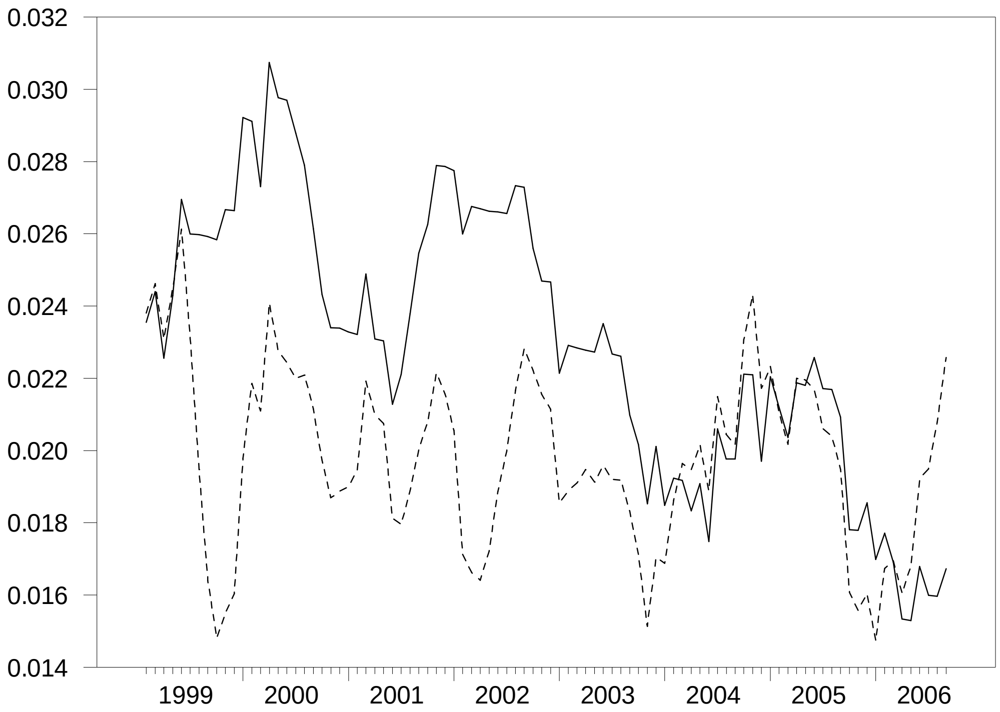


Figure 8. Germany: observed (solid line) **and "new" inflation** (dotted line)

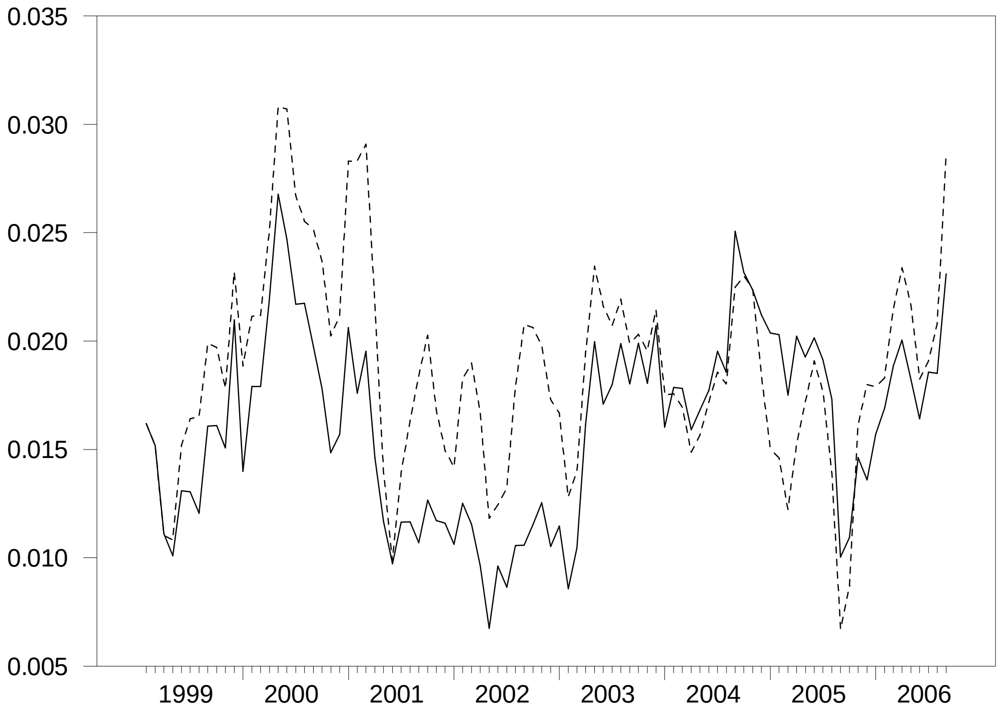


Figure 9. Italy: observed (solid line) **and "new" output** (dotted line)

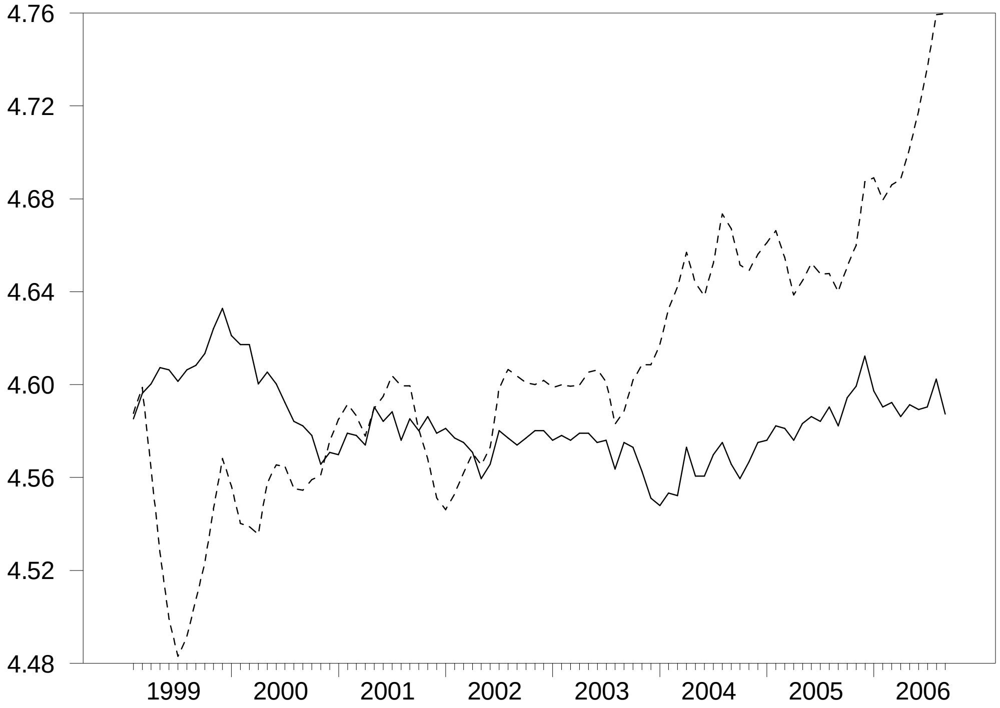


Figure 10. Germany: observed (solid line) **and "new" output** (dotted line)

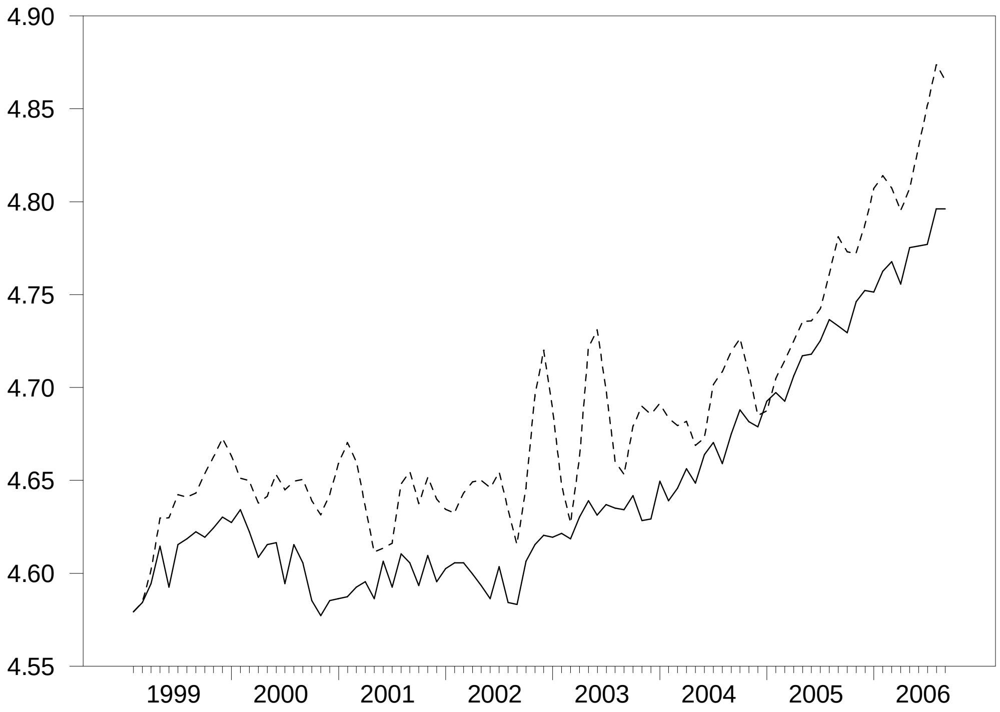


Figure 11. Italy: observed (squares) **and "new"** (dots) **AS curves**

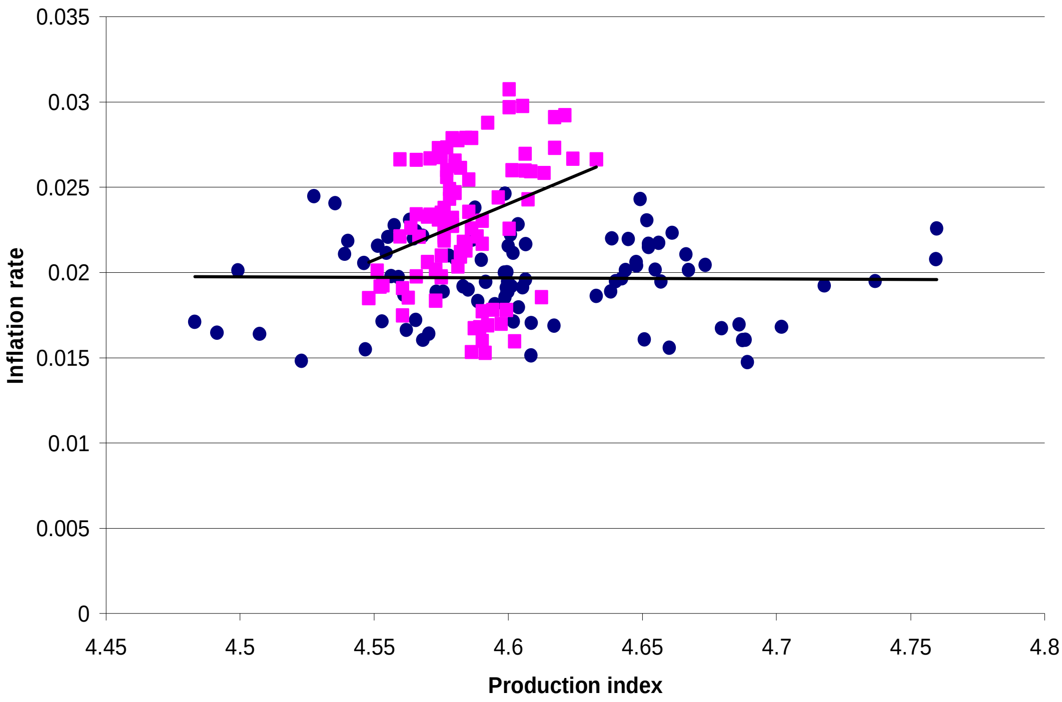
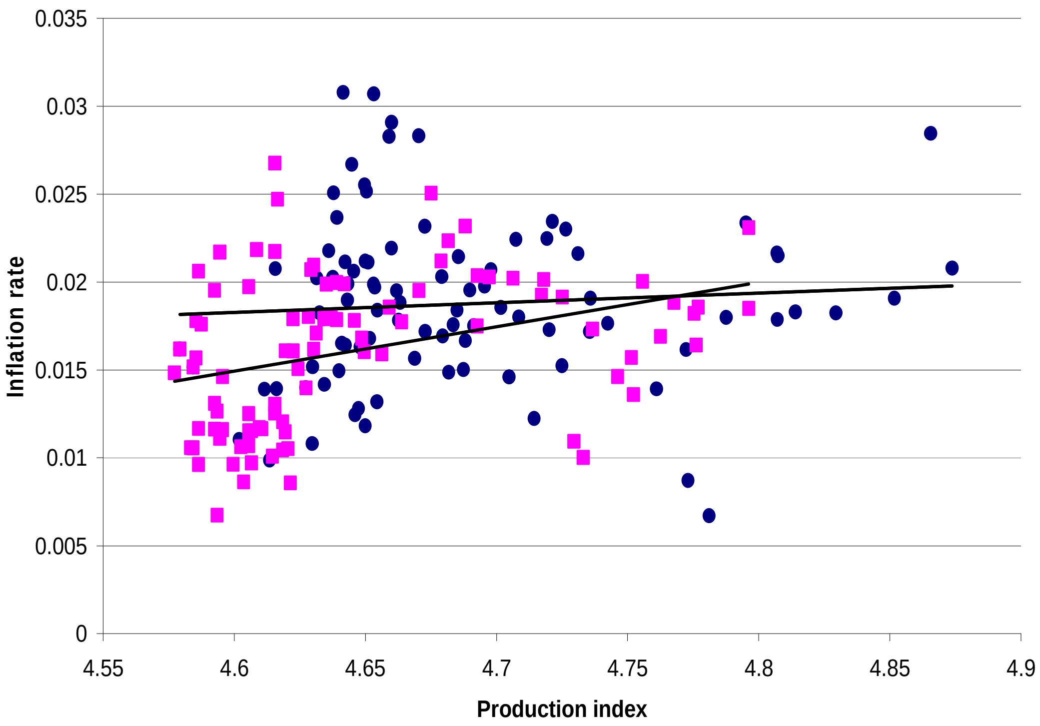


Figure 12 Germany: observed (squares) **and "new"** (dots) **AS curves**



Appendix

Focusing our attention just on the stochastic and deterministic trends $\mathbf{C}\sum_{i=1}^{t}\boldsymbol{\varepsilon}_i + \mathbf{C}\boldsymbol{\mu}_0 t$, where $\mathbf{C} = \tilde{\boldsymbol{\beta}}_{y\perp}\boldsymbol{\alpha}'_{\perp}$, and interpreting $\hat{\boldsymbol{\alpha}}'_{\perp}\sum_{i=1}^{t}\hat{\boldsymbol{\varepsilon}}_i$ as an estimate of the common stochastic trends $\sum_{i=1}^{t}\mathbf{u}_i$, and $\hat{\tilde{\boldsymbol{\beta}}}_{y\perp}$ as an estimate of their loading matrix, we get for the VMA representation of $\mathbf{y}_t$, subject to the retrictions on $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ previously imposed, the following results for Italy over the sub-period 1986-1998:

$$\begin{bmatrix} \pi_{t+12} \\ w_t \\ q_{t+12} \\ k_t \\ \rho_t \end{bmatrix} = \begin{bmatrix} \mathbf{-0.426} & 0.150 \\ \mathbf{-0.853} & \mathbf{-0.657} \\ \mathbf{-1.040} & 0.386 \\ \mathbf{1.274} & 0.108 \\ \mathbf{1.841} & 0.250 \end{bmatrix} \begin{bmatrix} \hat{\boldsymbol{\alpha}}'_{\perp,1}\sum_{i=1}^{t}\hat{\boldsymbol{\varepsilon}}_i \\ \hat{\boldsymbol{\alpha}}'_{\perp,2}\sum_{i=1}^{t}\hat{\boldsymbol{\varepsilon}}_i \end{bmatrix} + \begin{bmatrix} -0.000 \\ 0.001 \\ 0.002 \\ -0.001 \\ -0.000 \end{bmatrix} t + \ldots$$

Given the weak exogeneity of k_t, the first common stochastic trend is defined as $\sum_{i=1}^{t}\hat{\varepsilon}_{ki}$, the cumulated empirical shocks to the inter-bank rate[32], and it can be labelled as a nominal stochastic trend. We can see that this stochastic trend has significantly and negatively affected inflation, real wages and production, consistently with our model.

In order to identify the second common stochastic trend we impose a 0 restriction on the coefficient of $\hat{\varepsilon}_{kt}$ and a normalization restriction on the coefficient of $\hat{\varepsilon}_{\rho t}$ in the second row of the $\hat{\boldsymbol{\alpha}}'_{\perp}$ matrix. Thus we have:

$$\hat{\boldsymbol{\alpha}}'_{\perp,2}\hat{\boldsymbol{\varepsilon}}_t = \underset{(1.888)}{2.012\hat{\varepsilon}_{\pi t}} \underset{(-1.804)}{-0.5043\hat{\varepsilon}_{wt}} \underset{(0.895)}{+0.083\hat{\varepsilon}_{yt}} + 1.000\hat{\varepsilon}_{\rho t}$$

which shows that it is associated with risk, inflation and real wages, the last two being just borderline significant.

As a consequence of this identification procedure the $\hat{\tilde{\boldsymbol{\beta}}}_{y\perp}$ matrix corresponds exactly to the fourth and fifth column of the **C** matrix. Therefore, we can again argue that the long-run impact of cumulated shocks to the inter-bank rate is negative for inflation, real wages and production and positive for risk, while the long-run impact of cumulated shocks to risk have a negative long-run impact just on real wages.

The estimates of the linear trend effects show positive growth rates for output and real wages, , while the interbank-rate shows a negative growth rate.

The VMA representation subject to the retrictions on $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ previously imposed, has given the following results for Italy over the sub-period 1999-2007:

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[32] Imposing the assumption of weak exogeneity of the inter-bank rate, the corresponding $\boldsymbol{\alpha}_{\perp,1}$ is just a unit vector,

$$\begin{bmatrix} \pi_{t+12} \\ w_t \\ q_{t+12} \\ k_t \\ \rho_t \end{bmatrix} = \begin{bmatrix} \mathbf{-0.332} & \mathbf{0.307} \\ \mathbf{-1.116} & \mathbf{1.027} \\ \mathbf{-4.089} & \mathbf{-2.290} \\ \mathbf{1.055} & 0.246 \\ \mathbf{1.892} & \mathbf{1.660} \end{bmatrix} \begin{bmatrix} \hat{\boldsymbol{\alpha}}'_{\perp,1} \sum_{i=1}^{t} \hat{\boldsymbol{\varepsilon}}_i \\ \hat{\boldsymbol{\alpha}}'_{\perp,2} \sum_{i=1}^{t} \hat{\boldsymbol{\varepsilon}}_i \end{bmatrix} + \begin{bmatrix} -0.000 \\ -0.001 \\ -0.002 \\ 0.001 \\ 0.000 \end{bmatrix} t + \ldots$$

Following the exact identification procedure of normalizing on the coefficient of $\hat{\varepsilon}_{kt}$ in the first common stochastic trend and on the coefficient of $\hat{\varepsilon}_{\pi t}$ in the second, we have that the first common stochastic trend is defined as the cumulated empirical shocks to the inter-bank rate and to risk, while the second is defined as the cumulated empirical shocks to inflation with significant contribution from risk and production:

$$\hat{\boldsymbol{\alpha}}'_{\perp,1}\hat{\boldsymbol{\varepsilon}}_t = \underset{(-0.359)}{-0.032\hat{\varepsilon}_{wt}} \underset{(-1.561)}{-0.104\hat{\varepsilon}_{yt}} + 1.000\hat{\varepsilon}_{kt} \underset{(-3.144)}{-0.272\hat{\varepsilon}_{\rho t}}$$

$$\hat{\boldsymbol{\alpha}}'_{\perp,2}\hat{\boldsymbol{\varepsilon}}_t = 1.000\hat{\varepsilon}_{\pi t} \underset{(1.935)}{+0.213\hat{\varepsilon}_{wt}} \underset{(-0.243)}{-0.020\hat{\varepsilon}_{yt}} \underset{(2.379)}{+0.258\hat{\varepsilon}_{\rho t}}$$

We can note that the first stochastic trend has significantly and negatively affected inflation, real wages and production, as expected, while the second has significantly and negatively affected production, but negatively inflation and real wages. Comments related to the matrix **C** in this sub-period have already been made in the paper.

The estimates of the linear trend effects show negative growth rates for inflation, output and real wages, while the interbank-rate shows a positive growth rate.

For Germany the identification of the common stochastic and deterministic trends through the VMA representation with $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ restricted, has given the following results over the sub-period 1986-1998:

$$\begin{bmatrix} \pi_{t+12} \\ w_t \\ q_{t+12} \\ k_t \\ \rho_t \end{bmatrix} = \begin{bmatrix} \mathbf{-0.336} & \mathbf{-0.224} \\ \mathbf{-1.576} & 0.309 \\ \mathbf{-3.337} & \mathbf{-2.245} \\ \mathbf{1.593} & 0.098 \\ \mathbf{-0.903} & 0.210 \end{bmatrix} \begin{bmatrix} \hat{\boldsymbol{\alpha}}'_{\perp,1} \sum_{i=1}^{t} \hat{\boldsymbol{\varepsilon}}_i \\ \hat{\boldsymbol{\alpha}}'_{\perp,2} \sum_{i=1}^{t} \hat{\boldsymbol{\varepsilon}}_i \end{bmatrix} + \begin{bmatrix} -0.000 \\ 0.003 \\ 0.001 \\ -0.000 \\ 0.000 \end{bmatrix} t + \ldots$$

where the first common stochastic trend is defined as $\sum_{i=1}^{t} \hat{\varepsilon}_{ki}$, due to the weak exogeneity of k_t, and the second common stochastic trend is given by the cumulated following empirical shocks:

$$\hat{\boldsymbol{\alpha}}'_{\perp,2}\hat{\boldsymbol{\varepsilon}}_t = \underset{(-1.405)}{-0.542\hat{\varepsilon}_{\pi t}} \underset{(-0.934)}{-0.067\hat{\varepsilon}_{wt}} \underset{(-1.704)}{-0.270\hat{\varepsilon}_{yt}} + 1.000\hat{\varepsilon}_{\rho t}$$

These shocks are made up mainly by risk, with a borderline significant contribution from output.

We can note that the first stochastic trend has significantly and negatively affected inflation, real wages and production, as expected, while the second has significantly and negatively affected inflation and production. From the $\hat{\tilde{\boldsymbol{\beta}}}_{y\perp}$ corresponding in this identification scheme to the fourth and fifth column of the **C** matrix, we can again argue that the long-run impact of cumulated shocks to the inter-bank rate is negative for inflation, real wages and production, and the long-run impact of cumulated empirical shocks to risk is negative for inflation and production.

The estimates of the linear trend effects show positive growth rates for real wages, while output grows at a slower rate.

The VMA representation with $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ restricted, has given the following results for Germany over the sub-period 1999-2007:

$$\begin{bmatrix} \pi_{t+12} \\ w_t \\ q_{t+12} \\ k_t \\ \rho_t \end{bmatrix} = \begin{bmatrix} \mathbf{-0.528} & 0.103 \\ \mathbf{-2.347} & -1.232 \\ \mathbf{-3.931} & -2.416 \\ \mathbf{1.536} & 1.252 \\ -0.522 & -1.298 \end{bmatrix} \begin{bmatrix} \hat{\boldsymbol{\alpha}}'_{\perp,1} \sum_{i=1}^{t} \hat{\boldsymbol{\varepsilon}}_i \\ \hat{\boldsymbol{\alpha}}'_{\perp,2} \sum_{i=1}^{t} \hat{\boldsymbol{\varepsilon}}_i \end{bmatrix} + \begin{bmatrix} -0.000 \\ -0.001 \\ 0.001 \\ 0.000 \\ 0.000 \end{bmatrix} t + \ldots$$

Following the exact identification procedure of normalizing on the coefficient of $\hat{\varepsilon}_{kt}$ in the first common stochastic trend and on the coefficient of $\hat{\varepsilon}_{\pi t}$ we have that the first common stochastic trend is defined as the cumulated empirical shocks to the inter-bank rate and to risk, while the second is defined as the cumulated empirical shocks to inflation with significant contribution from real wages:

$$\hat{\boldsymbol{\alpha}}'_{\perp,1}\hat{\boldsymbol{\varepsilon}}_t = \underset{(1.409)}{0.032}\hat{\varepsilon}_{wt} - \underset{(-0.282)}{0.009}\hat{\varepsilon}_{yt} + 1.000\hat{\varepsilon}_{kt} + \underset{(4.003)}{0.950}\hat{\varepsilon}_{\rho t}$$

$$\hat{\boldsymbol{\alpha}}'_{\perp,2}\hat{\boldsymbol{\varepsilon}}_t = 1.000\hat{\varepsilon}_{\pi t} + \underset{(2.036)}{0.094}\hat{\varepsilon}_{wt} - \underset{(-1.827)}{0.116}\hat{\varepsilon}_{yt} - \underset{(-1.155)}{0.565}\hat{\varepsilon}_{\rho t}$$

The first stochastic trend has significantly and negatively affected inflation, real wages and production, as expected, while the second has no significant effects. Comments related to the matrix **C** have already been made in the paper.

The estimates of the linear trend effects show negative growth rates for inflation and real wages, while the others show positive growth rates.

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