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The Macroeconomic Consequences of EMU: International Evidence from a DSGE Model

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Abstract

In this paper, we estimate a New Keynesian DSGE model developed by Ireland (2003) on French, German and Spanish data with the aim to explore the macroeconomic consequences of EMU. In order to validate the results from the DSGE model, we amend this analysis by stability tests of monetary policy reaction functions for these countries. We find that (a) the DSGE structure is well suited for the characterization of key macroeconomic features of the three economies; (b) significant efficiency gains were realized in terms of lower adjustment cost of prices and the capital stock; (c) the behavior of monetary policy did not change in Germany, unlike in France and Spain. Specifically, the impact of inflation on interest rates increased considerably in the two latter countries.

Keywords: DSGE, Monetary Policy, EMU
JEL Classification: E31, E32, E52

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1 Introduction

> "Don't you think it is too early to tell?"
> Chou En-Lai in response to Henry Kissinger's question about the consequences of the French revolution.

In terms of macroeconomic outcomes, the record after 10 years of European Monetary Union (EMU) looks quite favorable (European Commission 2008), despite some more sceptical voices that usually argue along the lines of the time-honored argument that "one size doesn't fit all" (see e.g. Moons and Van Poeck 2008). It is, however, much more difficult to pin down the channels through which monetary unification in general and EMU in particular indeed works. Attempts to do so, focused on a variety of issues.

One strand of the literature looks at the implications of a common currency for other economic institutions like regulation or wage setting; see e.g. von Hagen (1999), Cukierman and Lippi (2001), Jerger (2002) and Fratzscher and Stracca (2009). A second one looks at the (change of) different transmission channels of monetary policy, usually by employing some variety of a (structural) VAR model (e.g. Ehrmann 2003, Angeloni and Ehrmann 2006 and Jarocinski 2008). Thirdly, the availability of micro data, especially for loans and prices, led to a large literature that usually identifies statistically and economically significant convergence across countries due to monetary union (Beck and Weber 2005, Ongena and Popov 2009). A fourth and relatively recent literature uses dynamic stochastic general equilibrium (DSGE) models to characterize the Euro area or the economies in this region within some well-defined theoretical framework (e.g. Milani 2009, Reis 2009).

In this paper, we contribute to the last strand of the literature by separately estimating the same DSGE model for three European economies (France, Germany, Spain). The model allows to look at changes in the "deep" structural parameters as well as policy parameters over time. We focus on two issues, namely price rigidities and monetary policy behavior. Whereas price rigidities may or may not be affected by monetary union – despite some good arguments in favor of such an effect –, it is clear that changes in and convergence of monetary policy behavior simply must be present and therefore should be detected by the model. If one accepts this point, our exercise might also be interpreted as a plausibility test of DSGE models. Although within these models, the link between economic theory and data is as close as it may get, there is still a lot of scepticism about DSGEs concerning their use in policy analysis (see Chari et al. 2009). Therefore, it is interesting in its own right whether the theoretical structure of the model in Ireland (2003), originally applied to US data, leads to plausible results for different countries and samples. Among other things, we look at the question whether the data are consistent with the hypothesis that monetary policy converged to the pre-EMU behavior of the German Bundesbank.

This paper is one of only a few (see Langedijk and Roeger (2007) for a similar effort) that applies a given DSGE model to different countries. In the light of the rather critical assessments of DSGEs in the the recent literature, the ability of such a structure to generate reasonable estimates across different countries is at least noteworthy and may justify some confidence in this class of models.

A major problem that hurts the analysis of EMU in even moderately sophisticated structural models is that the available 10 years of EMU data do not suffice for a reliable estimation of the parameters that could then be compared to a pre-EMU sample. This problem is aggravated by the fact that EMU didn't come as a surprise in 1999, although the odds of becoming a member of EMU differed considerably across countries before 1999. Nevertheless, the massive interest rate convergence in the 1990s suggests major policy changes during the run-up to EMU (see Begg et al. 1997 for an analysis of the "end-game" involved). In this sense, it is simply impossible to clearly distinguish between EMU and pre-EMU samples. Therefore, the evaluation of the macroeconomic consequences of EMU may suffer from the same evaluation problem as the effects of the French revolution, famously stated in the introductory quote above by Chou En-Lai.

In the present paper we deal with this problem in two ways. First, we estimate the model on a pre-EMU sample (1980q1-1998q4) and then compare the results to the full sample (1980q1-2008q3). This ensures sufficiently large samples for reliable estimates and at the same time allows to detect differences between the two samples. Second, we perform formal stability tests on monetary policy reaction functions outside the DSGE model. This allows us to pin down the time of significant changes in policy behavior.

The DSGE framework employed in this paper is due to Ireland (2003). This model is especially designed to look at the relevance of price rigidities and the behavior of monetary policy.

The rest of the paper is structured as follows. Section 2 presents the model. Data issues are discussed in section 3, whereas the estimation results are presented and interpreted in section 4. The stability test on the coefficients of the monetary policy rule is conducted in section 5. Section 6 concludes the paper.

2 The Model

The model we use for Germany, France and Spain is developed and applied to US data in Ireland (2003). It is a closed-economy New Keynesian setting featuring a representative household, a representative finished goods-producing firm, a continuum of intermediate goods-producing firms indexed by $i \in [0, 1]$ and a monetary policy authority. During each period $t = 0, 1, 2, ...$, the intermediate goods producing firms produce a distinct, perishable intermediate good, also indexed by $i \in [0, 1]$. As usual, the solution requires these firms to be treated symmetrically. We now proceed to characterize the decisions taken by households and firms before looking at the behavior of the monetary authority and sketching the solution of the model.

Before describing the model, it is necessary to comment on the fact that we apply a closed-economy model to these very open economies. The most important reason is the obvious fact that we get around the notorious difficulties of modeling exchange rates and their implications for bilateral trade flows. In the present context, we are not particularly interested in those, since the exchange rate consequences of EMU on member states are pretty clear. Furthermore, openness makes it very difficult to characterize the process of capital formation that is a central part of the present model. See also the discussion by DiCecio and Nelson (2007) who apply a closed-economy model to the UK.

2.1 Households

The representative household enters period t holding M_{t-1}, B_{t-1} and K_{t-1} units of money, one-period bonds and physical capital rented to the intermediate goods sector, respectively. In addition to this endowment, the household receives a lump sum transfer T_t from the monetary authority at the beginning of period t. The household receives $W_t h_t$ units of labor income, with W_t denoting the nominal wage rate and h_t working hours; $K_t Q_t$ in capital income, where Q_t represents the rental rate for capital and K_t household's capital supply; and a nominal dividend D_t from the intermediate goods producing firm. Each source of income is measured in units of money.

The household uses its funds to purchase new bonds at the nominal cost B_t/r_t, where r_t denotes the gross nominal interest rate between time periods, or output from the final goods sector at price P_t. This good can be used for consumption C_t or investment I_t.

In the latter case, quadratic capital adjustment cost given by

$$\frac{\phi_K}{2}\left(\frac{K_{t+1}}{gK_t}-1\right)^2 K_t \tag{1}$$

accrue to the firm. g denotes the steady state growth rate of the capital stock. $\phi_K \geq 0$ governs the size of these adjustment costs. The capital accumulation process is given by $K_{t+1} = (1-\delta)K_t + x_t I_t$, with $0 < \delta < 1$ denoting the rate of depreciation and x_t representing a shock to the efficiency of investment. This shock is specified as

$$\ln(x_t) = \rho_x \ln(x_{t-1}) + \varepsilon_{xt}, \tag{2}$$

with $0 < \rho_x < 1$ and $\varepsilon_{xt} \sim N(0, \sigma_x^2)$ as introduced by Greenwood, Hercowitz and Huffman (1988).

The budget constraint of the representative household is given by

$$\frac{M_{t-1} + T_t + B_{t-1} + W_t h_t + Q_t K_t + D_t}{P_t} \geq C_t + I_t + \frac{\phi_K}{2}\left(\frac{K_{t+1}}{gK_t}-1\right)^2 K_t + \frac{B_t/r_t + M_t}{P_t}.$$

Facing this constraint, the household maximizes the stream of expected utility given by

$$E\sum_{t=0}^{\infty}\beta^t\{a_t[\gamma/(\gamma-1)]\ln[C_t^{(\gamma-1)/\gamma} + e_t^{1/\gamma}(M_t/P_t)^{(\gamma-1)/\gamma}] + \eta\ln(1-h_t)\}, \tag{3}$$

where $0 < \beta < 1$ is a discount factor. $\eta > 0$ measures the relative weight of leisure. $-\gamma$ can be easily shown to be the interest rate elasticity of money demand. (3) contains two preference shocks, which are both assumed to follow an autoregressive process. More specifically,

$$\ln(a_t) = \rho_a \ln(a_{t-1}) + \varepsilon_{at}, \tag{4}$$

where $0 < \rho_a < 1$ and $\varepsilon_{at} \sim N(0, \sigma_a^2)$ denotes an IS shock (McCallum and Nelson 1999), whereas

$$\ln(e_t) = (1-\rho_e)\ln(e) + \rho_e \ln(e_{t-1}) + \varepsilon_{et} \tag{5}$$

with $0 < \rho_e < 1$,$e > 0$ and $\varepsilon_{et} \sim N(0, \sigma_e^2)$ represents a money demand shock.

2.2 Firms

The final good Y_t is produced by firms acting in a perfectly competitive market by combining the intermediate goods $Y_t(i)$ according to

$$\left[\int_0^1 Y_t(i)^{(\theta-1)/\theta} di\right]^{\theta/(\theta-1)} \geq Y_t,$$

where $\theta > 1$ represents the elasticity of substitution between intermediate goods $Y_t(i)$. With $P_t(i)$ denoting the price of intermediate good i, profit maximization leads to the following demand function for intermediate goods

$$Y_t(i) = \left[\frac{P_t(i)}{P_t}\right]^{-\theta} Y_t, \tag{6}$$

where $P_t = \left[\int_0^1 P_t(i)^{1-\theta} di\right]^{1/(1-\theta)}$.

Each intermediate good i is produced by a single monopolistically competitive firm according to the constant returns to scale technology

$$K_t(i)^\alpha [g^t z_t h_t(i)]^{1-\alpha} \geq Y_t(i),$$

where g denotes the gross rate of labor-augmenting technological progress and $1 > \alpha > 0$ represents the elasticity of capital with respect to output. The technology shock z_t follows the autoregressive process

$$\ln(z_t) = (1-\rho_z)\ln(z) + \rho_z \ln(z_{t-1}) + \varepsilon_{zt} \tag{7}$$

with $1 > \rho_z > 0$, $z > 0$ and $\varepsilon_{zt} \sim N(0, \sigma_z^2)$. As it is clear from (6), each firm i exerts some market power, but is assumed to act as a price taker in the factor markets. Furthermore, the adjustment of its nominal price $P_t(i)$ is assumed to be costly, where the cost function is concave in the size of the price adjustment. More specifically, following Rotemberg (1982), these costs are specified as

$$\frac{\phi_P}{2}\left[\frac{P_t(i)}{\pi P_{t-1}(i)} - 1\right]^2 Y_t, \tag{8}$$

where $\phi_P \geq 0$ governs the size of price adjustment costs and π denotes the gross steady-state rate of inflation targeted by the monetary authority (described below). Due to the concavity of (8), the firm's problem becomes dynamic. It chooses $h_t(i), K_t(i), Y_t(i)$ and $P_t(i)$ to maximize its total market value $E\sum_{t=0}^{\infty}\beta^t\lambda_t[D_t(i)/P_(t)]$, where λ_t measures the period t marginal utility to the representative household provided by an additional dollar of profits that are distributed to the household as dividends. These dividends are defined in real terms by

$$\frac{D_t(i)}{P_t} = \left[\frac{P_t(i)}{P_t}\right] Y_t(i) - \frac{W_t h_t(i) + Q_t K_t(i)}{P_t} - \frac{\phi_P}{2}\left[\frac{P_t(i)}{\pi P_{t-1}(i)} - 1\right]^2 Y_t.$$

2.3 Monetary policy

Monetary policy is represented by a generalized Taylor (1993) rule of the form

$$\omega_r \ln(r_t/r) = \omega_\mu \ln(\mu_t/\mu) + \omega_\pi \ln(\pi_t/\pi) + \omega_y \ln(y_t/y) + \ln(\upsilon_t). \tag{9}$$

This specification encompasses monetary policies that are conducted by steering interest rates r_t, gross money growth $\mu_t = M_t/M_{t-1}$ or any (linear) combination thereof. These monetary policy instruments may respond to deviations of gross inflation $\pi_t = P_t/P_{t-1}$ and detrended output $y_t = Y_t/g^t$ from their steady-state values. Clearly, $\omega_\mu = 0$ and $\omega_\pi/\omega_r > 1$ generate the usual Taylor rule.

The monetary policy shock υ_t follows the autoregressive process

$$\ln(\upsilon_t) = \rho_v \ln(\upsilon_{t-1}) + \varepsilon_{\upsilon t}, \tag{10}$$

where $0 < \rho_v < 1$ and $\varepsilon_{\upsilon t} \sim N(0, \sigma_\upsilon^2)$. Following Ireland (2003), we normalize the standard deviation of $\varepsilon_{\upsilon t}$ by setting $\sigma_\upsilon = 0.01$.

It is important to note that this characterization of the monetary authority does not even ask the question of optimal monetary policy. (9) just describes monetary policy which, however, is enough for our question at hand.

2.4 Solution

The model is characterized by a set of nonlinear difference equations, namely the first-order conditions for the three agents' problems, the laws of motion for the five exogenous shocks (2), (4), (5), (7) and (10) and the monetary policy rule (9). Two additional steps are required to close the model. First, in order to get from sectoral to aggregate variables, symmetric behavior within the intermediate sector is assumed, implying

$P_t(i) = P_t$, $Y_t(i) = Y_t$, $h_t(i) = h_t$, $K_t(i) = K_t$ and $D_t(i) = D_t$ for all $i \in [0, 1]$. Second, the market clearing conditions for both the money market $M_t = M_{t-1} + T_t$ and the bond market $B_t = B_{t-1} = 0$ must hold for all $t = 0, 1, 2....$

Since the model is nonlinear, there is no exact closed-form solution. An approximate one is obtained by calculating the stationary representation of the model, computing the steady state, log-linearizing the system around the steady state and then applying the method of Blanchard and Kahn (1980) to solve linear difference models under rational expectations. The solution takes on the form of a state space representation with a state equation $\mathbf{s_t} = \mathbf{A}\mathbf{s_{t-1}} + \mathbf{B}\varepsilon_\mathbf{t}$ and an observation equation $\mathbf{f_t} = \mathbf{C}\mathbf{s_t}$, where $\mathbf{s_t}$ contains the model's state variables including the current capital stock, lagged real balances and the five exogenous shocks. $\varepsilon_\mathbf{t}$ consists of the mutually as well as serially uncorrelated innovations ε_{at}, ε_{et}, ε_{xt}, ε_{zt}, ε_{vt} and $\mathbf{f_t}$ comprises the model's flow variables including current values of consumption, investment, inflation and the nominal interest rate. The matrices $\mathbf{A}$, $\mathbf{B}$, and $\mathbf{C}$ contain (functions of) the "deep" as well as the policy rule parameters of the model. These parameters are estimated using maximum likelihood. As outlined in Hamilton (1994) or Canova (2007), the likelihood function of a state space model can be expressed in terms of one-step-ahead forecast errors of the observables, conditional on the initial observations, and of their recursive variance, both of which are obtained using the Kalman filter.

3 Data

To estimate the structural parameters of the model we use French, German and Spanish quarterly (seasonally adjusted) data for consumption, investment, money balances, inflation and the interest rate from 1980q1 to 1998q4 and 1980q1 to 2008q3, respectively. Consumption and investment are measured by real personal consumption and real gross fixed capital formation in per capita terms. Real money balances are constructed by dividing the monetary aggregate M3 (again per capita) by the consumer price index that is also used for our measure of inflation. The interest rate is measured by the three month money market rate. The data sources are detailed in the appendix.

Following Fagan, Henry and Mestre (2005), we deal with the break in the series for Germany due to re-unification by re-scaling the West German series for consumption, investment and money prior to re-unification by the ratio of the values for West Germany and Germany at re-unification.

While being aware of the potential problem of spuriousness, as discussed in DeJong and Dave (2007), we follow Ireland (2003) – and much of the DSGE literature – in using linearly detrended time series for (logs of) consumption, investment and M3. As an additional caveat one might note the result by Delle Chiaie (2009) who showed that Bayesian estimates of DSGE models are rather sensitive to different ways of detrending the data.

Despite its relative simplicity, the model contains a considerable number of parameters that are difficult to estimate precisely on only five time series. Hence, a number of parameters not central to the aim of our investigation is fixed prior to estimation. More specifically, η is set to 1.5 which implies that the representative household's labor supply in the steady state amounts to one-third of its time. In addition, the depreciation rate δ is set to 0.025, corresponding to an annual depreciation rate of about 10 percent and θ is fixed at 6, implying a steady state markup of prices over marginal cost of 20 percent. Lastly, following Sahuc and Smets (2008), we set the elasticity of capital with respect to output to 0.29, and equate the steady state money growth rate with the average rate of inflation in the data.

4 Results

In this section we present the maximum likelihood estimates for the parameters. For each country, tables 1 and 2 report the estimates for the 1980q1 to 1998q4 period and for the total sample (1980q1 to 2008q3), respectively. The modulus of the maximized value of the log likelihood function is indicated by $|L|$.

In order to interpret the results, we have to compare the estimated coefficients

- across countries $j \in \{\text{France, Germany, Spain}\}$ for a given sample $k \in \{1, 2\}$, where $k = 1$ indicates the short sample until 1998 and $k = 2$ the full sample; and

- across samples k for a given country j.

The first task is relatively straightforward since the samples on which the coefficients are estimated are disjoint. Denoting the point estimate of some parameter a for country j in sample k and the associated standard deviation by a_{jk} and $\sigma_{a_{jk}}$, respectively, we use the Andrews and Fair (1988) Wald test $W = \frac{(a_{jk}-a_{-jk})^2}{\sigma^2_{a_{jk}}+\sigma^2_{a_{-jk}}}$, where $j \neq -j$. W follows a $\chi^2(1)$ distribution under the null of $a_{jk} = a_{-jk}$.

	France		Germany		Spain	
Parameter	Estimate	Std Error	Estimate	Std Error	Estimate	Std Error
β	0.9999	0.0017	0.9957	0.0008	0.9983	0.0013
γ	0.0000	0.0053	0.0556	0.0134	0.0127	0.0056
ϕ_P	18.0112	5.3625	19.9115	8.3694	54.0482	22.4337
ϕ_K	25.9372	1.9953	21.7369	3.2108	18.7675	2.9162
ω_r	2.0503	0.2241	2.2517	0.2900	1.3221	0.1640
ω_μ	0.5733	0.1498	0.6468	0.1336	0.9282	0.1547
ω_π	1.8480	0.2326	3.0386	0.3601	0.5674	0.2324
ω_y	−0.2119	0.0434	0.0235	0.0179	−0.0846	0.0209
e	43.8794	1.3958	2.6598	0.1497	3.5073	0.1808
z	5328.3446	344.1424	2789.9589	875.4551	2207.5298	62.7261
ρ_a	0.9998	0.0002	0.9005	0.0254	0.9933	0.0086
ρ_e	0.9999	0.0001	0.8460	0.0484	0.9857	0.0086
ρ_x	0.9984	0.0015	0.9979	0.0023	0.9965	0.0039
ρ_z	0.9839	0.0069	0.9987	0.0019	0.9252	0.0188
ρ_υ	0.5336	0.0448	0.2856	0.0989	0.3618	0.0659
σ_a	0.0736	0.0073	0.0159	0.0018	0.0232	0.0049
σ_e	0.0116	0.0009	0.0152	0.0015	0.0094	0.0008
σ_x	0.0922	0.0158	0.0652	0.0155	0.0323	0.0101
σ_z	0.0240	0.0060	0.0215	0.0025	0.0314	0.0065
$\lvert L \rvert$	1351.1586		1309.3003		1288.8190	

Table 1: Maximum Likelihood Estimates: 1980-1998

The comparison between the two samples is less straightforward due to the fact that the pre-EMU sample ($k = 1$) is a (sizeable) part of the full sample ($k = 2$). Moreover, this overlap makes it rather difficult to detect changes of the parameters due to EMU. We handle this problem by treating the estimate from the pre-EMU sample as a fixed value and using a simple t-test in order to test whether the estimate from the full sample is significantly different. Formally, we employ the test statistic $\frac{|a_{j2}-a_{j1}|}{\sigma_{a_{j2}}}$.

A full set of the test statistics is available from the authors upon request.

Turning to the results, we first note that the estimates for the discount factor β are below unity, but exceed 0.99 for all of the three economies in both samples. This is well in line with economic intuition as well as the empirical literature – and gives some confidence in the suitability of the model specification. Moreover, neither country differences nor differences across the two samples are significant at a 1% level.

The money demand equation that follows from (3) implies an interest elasticity for real money holdings of $-\gamma$. Hence, we estimate significant, albeit small values of this elasticity with the correct sign for Germany and Spain on both samples. For France, the

	France		Germany		Spain	
Parameter	Estimate	Std Error	Estimate	Std Error	Estimate	Std Error
β	0.9999	0.0011	0.9970	0.0006	0.9976	0.0007
γ	0.0000	0.0048	0.0356	0.0105	0.0195	0.0064
ϕ_P	16.4569	3.7043	7.6166	3.4101	23.9138	7.1255
ϕ_K	24.1030	0.8750	21.3831	2.7162	13.4216	1.7388
ω_r	2.1694	0.2209	1.8002	0.1930	1.2694	0.1545
ω_μ	0.4661	0.1108	0.3819	0.0797	0.5268	0.1328
ω_π	2.4653	0.2094	2.6668	0.2171	0.9186	0.2710
ω_y	-0.3305	0.0575	0.0284	0.0138	-0.0713	0.0216
e	50.8731	2.4069	2.8224	0.1348	3.7047	0.1763
z	4907.6935	250.9087	3465.2643	389.6705	2071.0932	37.8501
ρ_a	0.9996	0.0009	0.8989	0.0190	0.9964	0.0042
ρ_e	0.9999	0.0001	0.9280	0.0294	0.9790	0.0112
ρ_x	0.9999	0.0001	0.9971	0.0021	0.9997	0.0004
ρ_z	0.9874	0.0036	0.9994	0.0009	0.9249	0.0190
ρ_v	0.5555	0.0307	0.2905	0.0681	0.2950	0.0754
σ_a	0.0709	0.0056	0.0148	0.0014	0.0216	0.0028
σ_e	0.0113	0.0008	0.0145	0.0011	0.0112	0.0008
σ_x	0.0929	0.0148	0.0552	0.0109	0.0227	0.0047
σ_z	0.0239	0.0049	0.0176	0.0016	0.0180	0.0023
$\|L\|$	2132.5877		1995.1746		1964.1076	

Table 2: Maximum Likelihood Estimates: 1980-2008

elasticity is insignificant and almost exactly zero. These results are in line with a large empirical literature detecting small and sometimes insignificant interest rate elasticities of money demand. For the pre-EMU sample, the estimates for both France and Spain are significantly different from the estimate for Germany. In the full sample, only the difference between France and Germany is significant at the 1% level.

Next, we turn to the estimates for the rigidity parameters. For all countries and both samples, the estimates are significant for both the adjustment cost parameter for capital ϕ_K defined in (1) and prices ϕ_P defined in (8).[1] Moreover, in the long sample, both coefficients are smaller for all countries. The decline in ϕ_P is significant at the 1% level in both Germany and France, whereas the decline in ϕ_K is statistically insignificant in the case of Germany; the p-values for France and Spain are 0.0383 and 0.0026, respectively.

To check the plausibility of the price adjustment parameters, we apply the ap-

[1] The p-values for all estimates of ϕ_K are less than 1%, for the estimates of ϕ_K, the highest p-value is 0.027 in the case of Germany.

proach of Keen and Wang (2007) to translate the estimates of ϕ_P into an average duration of quoted prices. For France and Germany we get an average duration until re-optimization of 7 to 8 months. The findings are supported for France by the results of Baudry et al. (2004) using French CPI micro-data. Spain shows a higher degree of price stickiness implying an average of 11 to 12 months between price adjustments. This is in line with micro evidence as reported in de Walque, Smets and Wouters (2006).

Turning to the monetary policy reaction function, we first observe that for all three countries and for both samples the reaction coefficients show a significant response of the short term nominal interest rate to deviations of money growth and inflation from their steady state values. Relative to France and Spain, ω_π is significantly higher in Germany. This reflects the well-documented higher pre-occupation with inflation in this country. Concerning the response of interest rates to money growth, these results are consistent with the findings of Andrés, Lopéz-Salido and Vallés (2006) for the euro area. The fact that money growth is more important for interest rate decisions in Spain might reflect the various monetary policy regimes in Spain from 1980 and 1998, described by Arghyrou and Gadea (2008). It is important to note that for each of the three countries the estimates of ω_μ/ω_r and ω_π/ω_r sum up to a value greater than unity. This ensures that the monetary policy rule is consistent with a unique rational expectations equilibrium (see Clarida, Gali and Gertler 2000).

The response of the interest rate to the output gap is significant with negative sign in both France and Spain, which might be interpreted as evidence for the presence of an endogenous money channel. The estimates of ω_y are positive, albeit very small for Germany. In the pre-EMU sample the estimate for Germany is insignificant; for the full sample, the p-value is just below the 5% benchmark (0.043). This finding again is in line with the predominance of the goal of price stability in Germany. Whereas the reduction of the coefficient for France between $k = 1$ and $k = 2$ is marginally significant ($p = 0.0414$), changes across samples are not significant for Germany and Spain.

The estimates of e and z are not interesting from an economic policy point of view; the parameters simply allow the steady state values of real balances and output in the model to match the average values of these variables in the data (Ireland 2003).

The estimates of ρ_a, ρ_e, ρ_x, ρ_z and ρ_v indicate a very high persistence of the first four shocks, whereas the monetary policy shock is less – albeit significantly – persistent. For each of the three countries, the estimated standard deviations of the innovations are dominated by the ones of the investment shock. This result is in line with the findings of Justiano, Primiceri and Tambalotti (2008) for the US economy, identifying the marginal efficiency of investment shock as the most important driver of business cycle fluctuations.

Summing up these results, there are three major points. First, despite the fact that a common structure of a specific DSGE model is imposed on three different economies, the results are plausible and in line with evidence from outside this class of models. This to some extent rebuts the scepticism voiced against the use of DSGE models. Second, the estimates for the rigidity parameters Φ_K and Φ_P declined significantly over time. Despite the methodological problem stemming from the lack of a proper "before-after"-comparison as discussed above, this can be interpreted as evidence for a more efficient economic environment due to the EMU. Third, the results clearly show the convergence of monetary policy behavior, as expected. It is worthwhile to note that whereas this is a trivial statement of the obvious for the policy rates after 1999, it is not trivial with respect to a specification of the monetary policy reaction function that is specified in terms of the three month market rate. Moreover, our modeling strategy in this section does not account for the possibility of policy convergence before 1999. Hence, the approach of estimating a structural model with two overlapping samples is clearly not appropriate to detect let alone clearly date changes in policy reaction functions. Therefore, we look at this reaction function more closely in the next section.

5 Stability of the monetary policy reaction function

More specifically, we try to identify dates of regime breaks in monetary policy behavior by using the Quandt-Andrews breakpoint test (also known as Andrews (1993) sup-Wald test). This procedure tests for one or more unknown structural breakpoints in the sample given a specified equation and a pre-specified admissible range within which the breakpoint is located. The basic idea is that simple Chow breakpoint tests are performed at every point within this range and then summarized into a single test statistic. More precisely, the values of Max LR-F, Ave LR-F and Exp LR-F, reported in table 3, are the maximum, the average and the log of the average of the exp, respectively, of the Likelihood Ratio F-statistics from each Chow test; see Andrews and Ploberger (1994) and Andrews et al. (1996) for details. The Max LR-F test identifies the most likely time of the break.

For each country, we estimate a reaction function of the form

$$\hat{r}_t = \rho_\mu \hat{\mu}_t + \rho_\pi \hat{\pi}_t + \rho_y \hat{y}_t + \varepsilon_{rt},$$

Country	Break identified	Max LR-F	Exp LR-F	Ave LR-F
Admissible range: 1984q4 - 2004q1				
France	2002q1	12.5312	4.7377	8.0249
		(0.0332)	(0.0106)	(0.0045)
Germany	1985q2	9.0710	2.0005	1.5627
		(0.1366)	(0.1799)	(0.5343)
Spain	1999q1	9.9564	3.6299	6.4767
		(0.0963)	(0.0321)	(0.0127)
Admissible range: 1989q1 - 1999q4				
France	1999q2	12.3640	4.7305	8.8550
		(0.0208)	(0.0098)	(0.0044)
Germany	1989q1	2.6186	0.4052	0.6854
		(0.7834)	(0.8361)	(0.8476)
Spain	1999q1	9.9564	3.9474	7.6779
		(0.0569)	(0.0218)	(0.0094)

Table 3: Quandt-Andrews tests for structural breaks

using OLS where the hat denotes linearly detrended and demeaned variables in logs as defined in section 3. The analysis is undertaken for the long sample 1980q1 to 2008q3, testing for structural breaks in the response coefficients ρ_μ and ρ_π. We look at two different admissible ranges, namely a "long range" (1984q4 to 2004 q1) and a "short range" (1989q1 to 1999q4). The former covers most of the sample as a plausibility check, whereas the latter is focused on the period of the run-up to and inception of the common monetary policy.

The results are reported in table 3. Below the test statistics, the p-values for rejecting the null hypothesis are given in parentheses. These probabilities have been calculated using Hansen's (1997) method.

In the case of Germany all three of the summary statistics measures fail to reject the null hypothesis of no structural breaks within for both ranges. The results for France, however, indicate a significant structural break in 2002q1 within the test sample 1984q4 to 2004q1, and 1999q2 for the 1989q1 to 1999q4 range. For Spain both ranges clearly (and significantly) identify the inception of EMU, i.e. 1999q1 as the date of the structural break. Taken together, these results strongly suggest that there was no discernible difference between the policies of the German Bundesbank (up to 1998) and the ECB. As members of EMU, France and Spain simply adjusted to this behavior. This more direct – and quite plausible – evidence corroborates the implications of the DSGE models concerning the change of monetary policy behavior.

6 Conclusions

Despite some scepticism voiced in the literature, DSGE models became increasingly popular also for the description and evaluation of monetary policy. Being firmly rooted in microeconomic foundations, this class of models is able to identify structural characteristics of economies – such as adjustment costs of different sorts – that are not easily recovered from a very parsimonious set of macroeconomic time series. Among other things, the use of DSGE models enables cross-country comparisons of such characteristics without having to resort to micro data.

In this paper we applied the New Keynesian DSGE model due to Ireland (2003) to France, Germany and Spain and formally tested the stability of monetary policy reaction functions for these countries. A general result worth mentioning is that the DSGE model could be successfully applied to these rather different countries, where "success" is defined by the plausibility of the estimation results and their consistency – e.g. concerning price stickiness – with evidence from outside the model. In order to identify the macroeconomic consequences of the EMU, one would ideally estimate a model structure on data before and after the inception of EMU. However, 10 years of data are not sufficient for reliable estimates of a DSGE model. Hence, we took a double approach of estimating the DSGE models on a pre-EMU sample and an (overlapping) full sample, and of formally testing for the stability of the monetary policy reaction functions.

Clearly, the comparison of estimation results from overlapping samples falls short of a proper "before-after" comparison. Nevertheless, our results point to efficiency gains over time in terms of lower adjustment costs, both for the capital stock and prices. Furthermore, the monetary policy reaction functions estimated within the models point to a convergence in France and Spain towards the behavior of the Bundesbank. The latter aspect is also present in formal tests for structural breaks for single equation estimations of monetary policy reaction functions. This again lends some confidence to the results obtained within the DSGE models.

Appendix: Data sources

- France:

  Real personal consumption: EUROSTAT

Gross fixed capital formation: EUROSTAT

Money balances (M3): Banque de France

Consumer price index: OECD

Interest rate (Pibor): OECD

Population: National Institute for Statistics and Economic Studies (INSEE)

- Germany:

  Real personal consumption: Federal Statistics Office

  Gross fixed capital formation: Federal Statistics Office

  Money balances (M3): Deutsche Bundesbank

  Consumer price index: OECD

  Interest rate (Fibor): OECD

  Population: Federal Statistics Office

- Spain:

  Real personal consumption: EUROSTAT

  Gross fixed capital formation: EUROSTAT

  Money balances (M3): Banco de España

  Consumer price index: OECD

  Interest rate (three-months money market rate): OECD

  Population: EUROSTAT

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