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Testing Heterogeneity within the Euro Area Using a Structural Multi-Country Model*

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Abstract

In this paper, we investigate the sources of heterogeneity within the euro area. For this purpose, we build an optimization-based multi-country model (MCM) that allows different sources of heterogeneity across countries, both in terms of behavior of economic agents and in terms of asymmetry of shocks. Using Bayesian techniques, we estimate the MCM and several constrained versions of this model. We then test different sources of heterogeneity within the euro area. We show that the main source is the asymmetry of shocks affecting the different economies. In contrast, the heterogeneity of behaviors does not seem to be of empirical relevance for the euro area.

Keywords: Euro area, heterogeneity, Bayesian econometrics, multi-country model

JEL Classification: C51, C52, F41

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1 Introduction

As stipulated in the Maastricht Treaty (Art. 105), the primary objective of the European Central Bank (ECB) is formulating and implementing monetary policy that guarantees price stability within the European Monetary Union. To this end, although it may use a battery of economic indicators, including country-specific ones, decisions are taken on the basis of aggregate developments, while national idiosyncrasies are left to the care of national governments. One clear disadvantage of such area-wide monetary policy is that it cannot stabilize output or inflation in response to a country-specific shock. This issue is exacerbated by the fact that national fiscal policies are strongly restricted in the context of the Stability and Growth Pact, so that individual countries cannot necessarily cope with country-specific shocks. Another important question in this context is whether national economies are likely to react in a similar way to area-wide monetary policy shocks.

While some studies clearly suggest that business cycles have converged to a large extent over the past decades (see the contributions in Angeloni, Kashyap, and Mojon, 2003), several recent studies focus on the differences between euro-area countries across several dimensions and obtain rather mixed evidence. A first source of heterogeneity, that may be named *structural heterogeneity*, corresponds to differences in the preferences, technology, and constraints of private agents across countries or, more generally, in the propagation mechanism of shocks within the economy (Benigno and López-Salido, 2002, Fabiani and Morgan, 2003, Angeloni and Ehrmann, 2004, and Campa and González Mìnguez, 2004). One particular aspect is the on-course process of convergence, as countries catch up to euro-area average. Other structural sources of heterogeneity across countries may be different industrial or sectorial concentrations, or lack of labour or capital mobility. In their investigation of the national Phillips curves, Benigno and Lopez-Salido (2002) and Fabiani and Morgan (2003) provide empirical evidence that large differences do exist in parameter estimates, even across "core" countries. Demertzis and Hugues Hallett (1998) also show that unemployment disparities are mainly due to structural differences in market fundamentals, rather than to the asymmetry of shocks.

A second component of heterogeneity is the asymmetry in the conduct of country-specific policies and may be named *policy heterogeneity* (Demertzis and Hugues Hallett, 1998, or Cecchetti, 1999). It includes monetary policy (until 1999), fiscal policy and regulation. On the eve of the start of the EMU, monetary policies within core countries were more coordinated and, in some sense, had already converged. In addition, as the internal market continues to deepen, there may be fewer regulatory differences between countries. Although fiscal policies are not coordinated, they are constrained to a some extent by the Stability and Growth Pact.

A last source of heterogeneity relies on the often called asymmetry of shocks across countries, or *stochastic heterogeneity* (Bayoumi and Eichengreen, 1993, Artis, 1999, and Verhoef, 2003).[1] One may expect that while the European integration progressively takes place, shocks

[1] In this context, the asymmetry of shocks generally means that country-specific shocks are weakly correlated,

become more and more symmetric within the area. The empirical evidence provided for core countries is that the correlation across countries of demand and supply shocks is generally rather large (Bayoumi and Eichengreen, 1993). However, there is no clear evidence that the correlation increased over the recent period, suggesting that there was no convergence of demand and supply shocks between core countries (Verhoef, 2003).

A difficulty with this previous evidence is the comparability of the results obtained with different approaches. On one hand, the tests of structural or policy heterogeneity have generally been performed using reduced-form, single-equation regressions.[2] On the other hand, the tests of stochastic heterogeneity have been typically based on structural vector autoregression (SVAR) models in order to identify structural shocks. Consequently, one cannot easily recover the relative importance of the different sources of heterogeneity.

The objective of this paper is to investigate the various sources of heterogeneity across euro-area countries within an *optimization-based framework*. For this purpose, we construct a multi-country model (MCM) that provides a simplified but fully micro-founded representation of the major economies of the area. The deep parameters pertaining to the private agents' behavior, the monetary policy rules, and the characteristics of the structural shocks are therefore explicitly specified. We then estimate simultaneously all the equations of the MCM, while allowing parameters for each country to differ from the others.

Our approach comprises several challenges both on theoretical and empirical grounds. From a theoretical point of view, we derive a simple but complete MCM which resorts to the "New Open Economy Macroeconomics" literature (Obstfeld and Rogoff, 2000). By incorporating significant frictions in the form of nominal rigidities, Dynamic Stochastic General Equilibrium (DSGE) models have been shown to provide a sufficiently rich dynamics to fit the actual data fairly well (Christiano, Eichenbaum, and Evans, 2003, or Smets and Wouters, 2003, SW thereafter). However, in our open-economy context, additional mechanisms must be introduced: (i) cross-country differences in the structural parameters are allowed, (ii) perfect risk sharing and a home bias in preferences are incorporated in the model to deal with exchange-rate indeterminacy, and (iii) cross-country correlations between shocks are introduce to capture co-movements in the joint dynamics of national conditions. From an empirical point of view, pure Full Information Maximum Likelihood (FIML) estimation turned out to be very sensitive to the specification of medium- or large-scale macroeconomic models and in many cases resulted in unrealistic parameter estimates. Consequently, we resort to Bayesian econometrics, which introduces priors on unknown parameters in an FIML framework. This avenue has been followed for instance by Schorfheide (2003), SW and Onatski and Williams (2004).

Following this strategy, we first model the joint dynamics of the major economies in the

so that countries cannot be viewed as driven by the same business cycle.

[2] An exception is the paper of Benigno and López-Salido (2002) that is based on the estimation of structural Phillips curves.

euro area (Germany, France and Italy) assuming full heterogeneity (complete MCM). Then, we investigate the various sources of heterogeneity described above. Since the underlying models are in general non-nested, we adopt the approach followed for instance by Schorfheide (2000) and Chang, Gomes, and Schorfheide (2002). It consists in defining a more general VAR model that is then used to compare the performances of the competing hypotheses. Our investigation reveals that the differences in the behavior of private agents or monetary authorities are not empirically relevant sources of heterogeneity across core countries of the euro area. In fact, the main source of heterogeneity is the asymmetry of shocks across countries. Interestingly, in a recent study, Smets and Wouters (2004) have found, in separate closed-economy models, a very similar result for the US and the euro area. Indeed, they obtain that the differences in business cycle behavior are due to differences in the type of shocks, rather than to differences in the structure of the economy or in the way the central bank responds to economic developments. To some extent, our investigation goes one step further, since it highlights that, even when international transmission mechanisms are taken into account, such results also hold for core countries of the euro area. Notice that a major source of such asymmetry of shocks may be the fiscal policy, that is known to differ dramatically from one country to the other. Of course, other structural mechanisms are missing in the model, that may explain the stochastic heterogeneity we find.

The remainder of the paper is organized as follows. In section 2, we describe the theoretical model designed to incorporate the different sources of heterogeneity described above. In section 3, we present the data and briefly describe the Bayesian approach we adopt to estimate our model. Then, we discuss the estimates of the MCM and some models imposing a given set of constraints on the MCM. In section 4, we test the different sources of heterogeneity between the countries under consideration. Section 5 summarizes our main findings and concludes.

2 Structure of the multi-country model

The euro area is modelled as the aggregate of several economies. For each country, we formulate an open-economy sticky-price model, which is inspired by recent theoretical models derived from the "New Open Economy Macroeconomics", and which has a sufficiently rich dynamics to fit actual data fairly well.[3] The model is enriched in several dimensions, to offer a comprehensive framework that encompasses and generalizes other previous contributions. Most elements of this model are individually already present in the closed or open economy macroeconomic literature, but they have not been brought together in a single framework as is done here. In terms of dynamics, first key modifications are the explicit incorporation of habit

[3] See, among others, Obstfeld and Rogoff (2000), Corsetti and Pesenti (2000), Devereux and Engel (2000), Clarida, Galí, and Gertler (2002), Smets and Wouters (2002), Benigno and Benigno (2003), Benigno (2004), Galí and Monacelli (2004).

formation in the households' preferences and partial indexation in a price-setting framework *à la* Calvo (1983). These assumptions provide us with micro-founded "hybrid" versions of the IS and Phillips curves. Second, we do not assume that preferences and technologies are the same across countries, since we are interested in studying the sources of heterogeneity within the euro area. In addition, domestic and foreign shocks are allowed to be imperfectly correlated. Third, to cope with the indeterminacy in the exchange rate, we resort to the perfect risk sharing assumption. Although this assumption is admittedly heroic in empirical work, it avoids assuming non-rational expectations of exchange rate that has been shown to be an alternative way of dealing with indeterminacy.[4] Finally, households are assumed to have a taste bias towards home-produced goods. Since preferences differ across countries, the price of consumption bundles will differ when expressed in a common currency. The real exchange rate thus deviates from purchasing power parity (PPP).[5] This assumption is crucial, because it allows the perfect-risk-sharing equation to determine uniquely the dynamics of the terms of trade.

In order to lighten the notations, we assume that there are two countries in the euro area, denoted H(ome) and F(oreign). Since commercial links are much stronger between countries within the area than with countries outside the area, we neglect trade with the rest of the world. The population of the euro area is a continuum of agents on the interval $[0, 1]$. The population of country H belongs to $[0, n)$, while the foreign population belongs to $[n, 1]$. Therefore, n is the relative measure of the home country size into the area. An agent in the home country is indexed $h \in [0, n)$, while a foreign agent is indexed $f \in [n, 1]$. Variables in the home country are denoted X_t while foreign variables are denoted X_t^*. The home economy produces a continuum of differentiated goods indexed on the interval $[0, n)$. Foreign goods (or, equivalently, goods produced in the rest of the area) are indexed on the interval $[n, 1]$. All goods are tradeable.

2.1 Households

The home economy is populated by infinitively-living households, consuming Dixit-Stiglitz aggregates of domestic and imported goods. A home household h owns a firm producing goods h and receives dividends from it. We assume that households in a given country have the same preferences and endowments. Although there may be idiosyncratic shocks among households, we assume that households have access to complete markets for state-contingent claims, so that there is no heterogeneity among agents in a given country. Consequently, all households in the same country behave in the same manner and then we consider the optimization problem of a representative household. The representative household in country H maximizes the following expected sequence of present and future utility flows that depends

[4] See, e.g., Lubik and Schorfheide (2003).

[5] An earlier contribution that introduced home bias in preferences is due to Warnock (2000).

positively on consumption (C_t) and negatively on labor (hours worked, L_t):[6]

$$\mathcal{U}_t = \mathbb{E}_t \sum_{k=0}^{\infty} \beta^k \varepsilon_{p,t+k} \left[\frac{1}{1-\sigma} \left(C_{t+k} - \gamma \mathcal{H}_{t+k} \right)^{1-\sigma} - \frac{1}{1+\varphi} \left(L_{t+k} \right)^{1+\varphi} \right] \tag{1}$$

where $\mathbb{E}_t$ denotes the expectation operator conditional on the information set at time t, β is the intertemporal discount factor, with $0 < \beta < 1$, σ is the inverse of the intertemporal elasticity of substitution of consumption, and φ is the inverse of the elasticity of labor disutility with respect to hours worked. $\varepsilon_{p,t}$ denotes a country-specific preference shock that affects the intertemporal substitution of all households in the same manner in the home economy.[7] Preferences display "external" habit formation as in Abel (1990). The habit stock is supposed to equal the level of aggregate consumption in the previous period ($\mathcal{H}_t = C_{t-1}$), and γ represents the habit persistence parameter, measuring the effect of past consumption on current utility ($0 \leq \gamma < 1$). Including habit formation in a macroeconomic model results in a better fit of the data and captures the "hump-shaped" gradual responses of spending (see Fuhrer, 2000).

The aggregated consumption index for home households and the corresponding consumption index for foreign households are defined by[8]

$$C_t = \frac{\left(C_{H,t} \right)^{\omega} \left(C_{F,t} \right)^{1-\omega}}{\omega^{\omega} \left(1-\omega \right)^{1-\omega}} \quad \text{and} \quad C_t^* = \frac{\left(C_{H,t}^* \right)^{\omega^*} \left(C_{F,t}^* \right)^{1-\omega^*}}{\left(\omega^* \right)^{\omega^*} \left(1-\omega^* \right)^{1-\omega^*}} \tag{2}$$

where ω and ω^* denote the share of home goods in the consumption of home and foreign households respectively. $C_{H,t}$ (resp. $C_{F,t}$) is the sub-index of consumption of imperfectly substitutable, home (resp. foreign) goods, which is in turn given by the following CES aggregators:

$$C_{H,t} = \left[\left(\frac{1}{n} \right)^{1/\theta} \int_0^n C_t \left(h \right)^{\frac{\theta-1}{\theta}} \mathrm{d}h \right]^{\frac{\theta}{\theta-1}} \quad \text{and} \quad C_{F,t} = \left[\left(\frac{1}{1-n} \right)^{1/\theta} \int_n^1 C_t \left(f \right)^{\frac{\theta-1}{\theta}} \mathrm{d}f \right]^{\frac{\theta}{\theta-1}} \tag{3}$$

where $C_t(h)$ (resp. $C_t(f)$) is consumption of the generic good h (resp. f) produced in country H (resp. F). Parameter θ denotes the elasticity of substitution across goods produced within a given country. The corresponding consumption price indexes (CPI) are given by:

$$P_t = \left(P_{H,t} \right)^{\omega} \left(P_{F,t} \right)^{1-\omega} \quad \text{and} \quad P_t^* = \left(P_{H,t}^* \right)^{\omega^*} \left(P_{F,t}^* \right)^{1-\omega^*}.$$

Here, $P_{H,t}$ (resp. $P_{F,t}$) is the price sub-index for home- (resp. foreign-) produced goods expressed in the home currency, defined as

$$P_{H,t} = \left[\frac{1}{n} \int_0^n P_{H,t} \left(h \right)^{1-\theta} \mathrm{d}h \right]^{\frac{1}{1-\theta}} \quad \text{and} \quad P_{F,t} = \left[\frac{1}{1-n} \int_n^1 P_{F,t} \left(f \right)^{1-\theta} \mathrm{d}f \right]^{\frac{1}{1-\theta}},$$

[6] We do not introduce money in the utility function. Given separability of the utility function, the resulting money demand equation would not affect the rest of the model.

[7] We assume that $\varepsilon_{p,t}$ follows an AR(1) process: $\varepsilon_{p,t} = \left(1-\rho_p\right) \bar{\varepsilon}_p + \rho_p \varepsilon_{p,t-1} + \eta_{p,t}$.

[8] As shown by Corsetti and Pesenti (2000), the Cobb-Douglas consumption index is a necessary condition for the trade to be invariably balanced.

where $P_{H,t}(h)$ (resp. $P_{F,t}(f)$) is the price in units of country H of a generic good h (resp. f) produced in country H (resp. F).

We also assume that prices are set in the producer currency and that the law of one price holds. We then have $P_{H,t}(h) = P^*_{H,t}(h) S_t$ and $P_{F,t}(f) = P^*_{F,t}(f) S_t$, where S_t is the nominal exchange rate expressed as units of domestic currency needed for one unit of foreign currency.[9] Since we assume the same elasticity of substitution among goods in a given country, we also have $P_{H,t} = P^*_{H,t} S_t$, and $P_{F,t} = P^*_{F,t} S_t$. Yet, from the definition of the CPI, we obtain that

$$P_t = P^*_t S_t \left(\frac{P_{H,t}}{P_{F,t}} \right)^{\omega - \omega^*}.$$

Therefore, if we assume that there exists a home bias in preferences ($\omega \neq \omega^*$), PPP does not necessarily hold, i.e. $P_t \neq P^*_t S_t$. We expect $\omega > \omega^*$, so that home households put a higher weight on home goods than foreign households.

As indicated above, we assume complete markets for state-contingent claims. Consequently, households can transfer wealth to the next period by holding B_{t+1} unit of the one-period nominal bond denominated in the domestic currency.[10] We thus obtain the following home household's budget constraint:

$$P_t C_t + \frac{B_{t+1}}{1+i_t} = W_t L_t + B_t + \Pi_t - TR_t \tag{4}$$

where W_t is the nominal wage income, Π_t is the dividend received from home firms, TR_t are lump sum government transfers, and i_t is the nominal interest rate.

The maximization problem of the home household consists in maximizing equation (1) subject to constraint (4), yielding the optimal profile of consumption, holdings of domestic bond and labor supply. The first-order conditions imply:[11]

$$\mathcal{U}_{C,t} = \varepsilon_{p,t} \left(C_t - \gamma \mathcal{H}_t \right)^{-\sigma}, \tag{5}$$

$$(1+i_t)^{-1} = \beta \mathbb{E}_t \left[\frac{\mathcal{U}_{C,t+1}}{\mathcal{U}_{C,t}} \frac{P_t}{P_{t+1}} \right], \tag{6}$$

$$\frac{\mathcal{U}_{L,t}}{\mathcal{U}_{C,t}} = \frac{W_t}{P_t}, \tag{7}$$

[9] Although it has been investigated in a number of recent papers, we do not consider here the presence of imperfect exchange rate pass-through. A reason is that it is not likely to be an important feature across countries within the euro area. In addition, this feature is clearly irrelevant from the euro-area point of view.

[10] More precisely, at date t, home households hold $B\left(s^{t+1}\right) = B_{t+1}$ units of the one-period bond denominated in home currency that pay 1 at date $t+1$ if state s_{t+1} occurs and 0 otherwise, where $s^t = (s_0, \cdots, s_t)$ denotes the story of events up to date t. Foreign households hold $B^*_t\left(s^{t+1}\right) = B^*_{t+1}$ units of such bond. The price of this bond in home currency is denoted $Q\left(s^t, s^{t+1}\right) = Q_{t,t+1}$. The price at date t of the portfolio held by home households is thus given by $E_t\left[Q_{t,t+1} B_{t+1}\right]$. We define the one-period interest rate as $1 + i_t = 1/E_t\left[Q_{t,t+1}\right]$. Note that, since bonds are state-contingent, including bonds denominated in foreign currency would be redundant. For more details, see Chari, Kehoe, and McGrattan (2002).

[11] We abstract here from the optimal intra-temporal allocations between domestic and foreign goods.

where $\mathcal{U}_{X,t}$ denotes the derivative of utility $\mathcal{U}$ with respect to variable X at the period t. Equation (5) defines the marginal utility of consumption. Equation (6) is the usual Euler equation for inter-temporal consumption flows. It establishes that the ratio of marginal utility of future and current consumption is equal to the inverse of the real interest rate. Equation (7) is the condition for the optimal consumption-leisure arbitrage, implying that the marginal rate of substitution between consumption and labor is equated to the real wage.

2.2 Firms

There is a continuum of infinitely living and monopolistically competitive firms indexed by h on the interval $[0, n)$ for the home country and by f on the interval $[n, 1]$ for the foreign country. They produce differentiated goods which are bundled into homogeneous home and foreign goods by a constant returns to scale of the Dixit-Stiglitz form:

$$Y_t = \left[\left(\frac{1}{n}\right)^{1/\theta} \int_0^n Y_t(h)^{\frac{\theta-1}{\theta}} \mathrm{d}h\right]^{\frac{\theta}{\theta-1}} \qquad \text{and} \qquad Y_t^* = \left[\left(\frac{1}{1-n}\right)^{1/\theta} \int_n^1 Y_t^*(f)^{\frac{\theta-1}{\theta}} \mathrm{d}f\right]^{\frac{\theta}{\theta-1}}.$$

The production technology of the representative home firm h combines labor as primary input and a country-specific technology shock.[12]

$$Y_t(h) = A_t L_t(h). \tag{8}$$

Output is normalized by population size, so that it is expressed in per capita terms. We thus deduce that total home labor demand is given by

$$L_t = \int_0^n L_t(h)\,\mathrm{d}h = \frac{Y_t V_t}{A_t} \tag{9}$$

where $V_t = \int_0^n \frac{Y_t(h)}{Y_t} \mathrm{d}h$ represents the dispersion of production across firms in the home economy.

Since input markets are perfectly competitive and technology shocks are country specific, the standard static first-order condition for cost minimization implies that all domestic firms have identical real marginal cost, MC_t, given by,

$$MC_t = \frac{1}{(1+\vartheta)} \frac{W_t}{P_{H,t} A_t} \tag{10}$$

where ϑ is a subsidy for output that offsets the effect on imperfect competition in goods markets on the steady-state level of output $(0 \leq \vartheta < 1)$.

Firms price setting decision is modelled through a modified version of the Calvo's (1983) staggering mechanism. In addition to the baseline mechanism, we allow for the possibility that firms that do not optimally set their prices may nonetheless adjust it to keep up with

[12] We assume that the technology shock A_t follows an AR(1) process: $A_t = (1-\rho_a)\bar{A} + \rho_a A_{t-1} + \eta_{a,t}$.

the previous period increase in the general price level (see Sbordone, 2003, and Christiano, Eichenbaum, and Evans, 2003, for details concerning this assumption). In each period, a firm faces a constant probability, $1-\alpha$, of being able to re-optimize its price and chooses the new price $\tilde{P}_{H,t}(h)$ that maximizes the expected discounted sum of profits

$$\mathbb{E}_t \sum_{k=0}^{\infty} \alpha^k \Upsilon_{t,t+k} \left[\frac{\tilde{P}_{H,t}(h)\, \Psi^H_{t,t+k}}{P_{H,t+k}} - MC_{t+k} \right] Y_{t+k}(h) \tag{11}$$

subject to the sequence of demand equations:

$$Y_{t+k}(h) = \left(\frac{\tilde{P}_{H,t}(h)\, \Psi^H_{t,t+k}}{P_{H,t+k}} \right)^{-\theta} Y_{t+k} \tag{12}$$

where $\Upsilon_{t,t+k} = \beta^k \mathcal{U}_C(C_{t+k}) / \mathcal{U}_C(C_t)$ is the discount factor between time t and $t+k$, and

$$\Psi^H_{t,t+k} = \begin{cases} \prod_{\nu=0}^{k-1} (\bar{\pi}_H)^{1-\xi} (\pi_{H,t+v})^{\xi} & k > 0 \\ 1 & k = 0, \end{cases} \tag{13}$$

where $\bar{\pi}_H$ is the domestic trend inflation and the coefficient $\xi \in [0,1]$ indicates the degree of indexation to past prices, during the periods in which firm is not allowed to re-optimize. $\Psi^H_{t,t+k}$ is a correcting term that accounts for the fact that, if the firm h does not re-optimize its price, it updates it according to the rule:

$$P_{H,t}(h) = (\bar{\pi}_H)^{1-\xi} (\pi_{H,t-1})^{\xi} P_{H,t-1}(h). \tag{14}$$

Consequently, the first-order condition associated to the profit maximization implies that firms set their price equal to the discounted stream of expected future real marginal costs:

$$\mathbb{E}_t \sum_{k=0}^{\infty} \alpha^k \Upsilon_{t,t+k} \left[(\bar{\pi}_H)^{(1-\xi)k} \left(\frac{P_{H,t+k-1}}{P_{H,t-1}} \right)^{\xi} \frac{\tilde{P}_{H,t}(h)}{P_{H,t+k}} - \frac{\theta}{\theta-1} MC_{t+k} \right] Y_{t+k}(h) = 0. \tag{15}$$

If flexible prices is assumed ($\alpha = 0$), this expression gives the optimal relative price $\tilde{P}_{H,t}(h) / P_{H,t} = \mu MC_t$, where $\mu \equiv \theta / (\theta - 1)$ is the optimal markup in a flexible-price economy. As there are no firm-specific shocks in this economy, all firms that are allowed to re-optimize their price at date t select the same optimal price $\tilde{P}_{H,t}(h) = \tilde{P}_{H,t}$, $\forall h$.

Staggered price setting under partial indexation implies the following expression for the evolution of the domestic price index:

$$P_{H,t} = \left[\alpha \left((\bar{\pi}_H)^{1-\xi} (\pi_{H,t-1})^{\xi} P_{H,t-1} \right)^{1-\theta} + (1-\alpha) \left(\tilde{P}_{H,t} \right)^{1-\theta} \right]^{\frac{1}{1-\theta}}. \tag{16}$$

The price setting problem solved by firms in the foreign country is similar and leads to an optimal rule analogous to equation (15). Yet, we allow foreign structural parameters (α^*, ξ^*) and country-specific shocks (A^*_t) to differ from their home country counterparts.

2.3 Real exchange rate dynamics

Under the assumption of complete markets, domestic and foreign households trade in state-contingent claims denominated in the home currency. This implies the following perfect risk-sharing condition (cf. Chari, Kehoe, and McGrattan, 2002):[13]

$$Q_t = \kappa \frac{\mathcal{U}^*_{C^*,t}}{\mathcal{U}_{C,t}} \tag{17}$$

where the real exchange rate, defined as $Q_t \equiv S_t P_t^*/P_t$, is proportional to the ratio of the marginal utility of consumption between the two countries.[14] The assumption of international market completeness insures that, in our model, the real exchange rate and consumption are stationary variables (see also Benigno, 2004).

Since the real exchange rate deviates from PPP because of home bias in preferences, we also have

$$Q_t = \left(\frac{S_t P^*_{H,t}}{P_{H,t}}\right)^{\omega^*} \left(\frac{S_t P^*_{F,t}}{P_{F,t}}\right)^{1-\omega^*} \left(\frac{P_{F,t}}{P_{H,t}}\right)^{\omega-\omega^*} = (\mathcal{T}_t)^{\omega-\omega^*} \tag{18}$$

where $\mathcal{T}_t$ is the home terms of trade, i.e. the relative price between foreign and home bundles of goods as perceived by the home resident. It is defined as[15]

$$\mathcal{T}_t = \frac{P_{F,t}}{P_{H,t}} = \frac{S_t P^*_{F,t}}{P_{H,t}}. \tag{19}$$

This definition implies, using equations (5), (17), and (18):

$$(\mathcal{T}_t)^{\omega-\omega^*} = \kappa \frac{\varepsilon^*_{p,t}\left(C_t - \gamma C_{t-1}\right)^{\sigma}}{\varepsilon_{p,t}\left(C_t^* - \gamma^* C^*_{t-1}\right)^{\sigma^*}}. \tag{20}$$

Equation (20) provides a rather elegant way to escape the exchange rate indeterminacy. Note that, when there is no home bias in preferences ($\omega = \omega^*$), the perfect risk-sharing assumption does not allow to determine the terms of trade anymore.

Combining Euler equation (6) with the perfect risk-sharing equation (17), we obtain the following dynamics for the real exchange rate and the terms of trade:

$$\mathbb{E}_t\left[\frac{Q_{t+1}}{Q_t}\right] = \mathbb{E}_t\left[\frac{\mathcal{U}^*_C\left(C^*_{t+1}\right)\mathcal{U}_C\left(C_t\right)}{\mathcal{U}^*_C\left(C^*_t\right)\mathcal{U}_C\left(C_{t+1}\right)}\frac{P^*_t P_{t+1}}{P_t P^*_{t+1}}\right] = \frac{1+i_t}{1+i^*_t} \tag{21}$$

$$\mathbb{E}_t\left[\frac{\mathcal{T}_{t+1}}{\mathcal{T}_t}\right] = \mathbb{E}_t\left[\frac{P^*_{F,t+1}P_{H,t}}{P_{H,t+1}P^*_{F,t}}\frac{1+i_t}{1+i^*_t}\right]. \tag{22}$$

[13] It is worth emphasizing that the exchange rates between the countries within the euro area have experienced significant changes over the estimation period. For instance, the French Franc and the Italian Lira have been depreciated several times with respect to the German Mark, notably for compensating loss of competitiveness. It is hence necessary to incorporate the relation between exchange rate and price differential in the structural model, even though the exchange rate is now fixed within the area.

[14] $\kappa = \left[S_0 P_0^* \mathcal{U}_{C,0}\right] / \left[P_0 \mathcal{U}^*_{C^*,0}\right]$ is a constant that depicts initial condition.

[15] The foreign terms of trade are simply given by $\mathcal{T}^*_t = P^*_{H,t}/P^*_{F,t} = 1/\mathcal{T}_t$, because the law of one price holds.

Equation (21) is the Uncovered Interest rate Parity (UIP) condition, which states that the expected change in the exchange rate is exactly compensated by the real interest rate differential. It is worth emphasizing that the UIP condition is not an additional implication in the model, but rather a redundant relation.

2.4 Market clearing conditions

Since the sub-indexes of consumption of home and foreign goods are given by equations (3), the home and foreign demands for generic goods h and f are given by

$$\begin{aligned} C_t(h) &= \frac{1}{n}\left(\frac{P_{H,t}(h)}{P_{H,t}}\right)^{-\theta} C_{H,t} \quad \text{and} \quad C_t^*(h) = \frac{1}{n}\left(\frac{P_{H,t}^*(h)}{P_{H,t}^*}\right)^{-\theta} C_{H,t}^* \\ C_t(f) &= \frac{1}{1-n}\left(\frac{P_{F,t}(f)}{P_{F,t}}\right)^{-\theta} C_{F,t} \quad \text{and} \quad C_t^*(f) = \frac{1}{1-n}\left(\frac{P_{F,t}^*(f)}{P_{F,t}^*}\right)^{-\theta} C_{F,t}^*. \end{aligned}$$

In addition, the consumption aggregator (2) implies that home and foreign demands for composite home and foreign goods are given by

$$\begin{aligned} C_{H,t} &= \omega\left(\frac{P_t}{P_{H,t}}\right) C_t \quad \text{and} \quad C_{H,t}^* = \omega^*\left(\frac{P_t^*}{P_{H,t}^*}\right) C_t^* \\ C_{F,t} &= (1-\omega)\left(\frac{P_t}{P_{F,t}}\right) C_t \quad \text{and} \quad C_{F,t}^* = (1-\omega^*)\left(\frac{P_t^*}{P_{F,t}^*}\right) C_t^*. \end{aligned}$$

Then, goods market clearing in the home and foreign countries implies:

$$\begin{aligned} Y_t(h) &= nC_t(h) + (1-n)\, C_t^*(h) \\ &= \left(\frac{P_{H,t}(h)}{P_{H,t}}\right)^{-\theta}\left(\frac{P_t}{P_{H,t}}\right)\left(\omega C_t + \mathcal{T}_t^{\omega-\omega^*}\frac{1-n}{n}\omega^* C_t^*\right) \end{aligned}$$

and

$$\begin{aligned} Y_t^*(f) &= nC_t(f) + (1-n)\, C_t^*(f) \\ &= \left(\frac{P_{H,t}(f)}{P_{F,t}}\right)^{-\theta}\left(\frac{P_t}{P_{F,t}}\right)\left(\frac{n}{1-n}(1-\omega)\, C_t + (1-\omega^*)\, \mathcal{T}_t^{\omega-\omega^*} C_t^*\right) \end{aligned}$$

so that aggregate outputs in home and foreign goods are:

$$Y_t = \omega\,(\mathcal{T}_t)^{1-\omega} C_t + \frac{1-n}{n}\omega^*\,(\mathcal{T}_t)^{1-\omega^*} C_t^* \tag{23}$$

and

$$Y_t^* = (1-\omega)\,(\mathcal{T}_t)^{-\omega}\frac{n}{1-n}C_t + (1-\omega^*)\,(\mathcal{T}_t)^{-\omega^*} C_t^*. \tag{24}$$

2.5 Log-linear equilibrium

In order to estimate this model, we approximate it using a first-order Taylor development around the steady state. The home-block, expressed in terms of percentage deviation around the steady state is given by the following equations:[16]

$$\begin{aligned}
\hat{c}_t &= \frac{\gamma}{1+\gamma}\hat{c}_{t-1} + \frac{1}{1+\gamma}\mathbb{E}_t\hat{c}_{t+1} - \frac{(1-\gamma)}{(1+\gamma)\,\sigma}\left(\hat{\imath}_t - \mathbb{E}_t\hat{\pi}_{H,t+1}\right) && (25)\\
&\quad + \frac{(1-\gamma)\,(1-\omega)}{(1+\gamma)\,\sigma}\mathbb{E}_t\Delta\hat{\tau}_{t+1} + \frac{\left(1-\rho_p\right)(1-\gamma)}{(1+\gamma)\,\sigma}\hat{\varepsilon}_{p,t} \\
\hat{\pi}_{H,t} &= \frac{\xi}{1+\beta\xi}\hat{\pi}_{H,t-1} + \frac{\beta}{1+\beta\xi}\mathbb{E}_t\hat{\pi}_{H,t+1} + \frac{(1-\beta\alpha)\,(1-\alpha)}{(1+\beta\xi)\,\alpha}\widehat{mc}_t && (26)\\
\widehat{mc}_t &= \left(\frac{\sigma}{1-\gamma} + \varphi\omega s\right)\hat{c}_t - \frac{\gamma\sigma}{1-\gamma}\hat{c}_{t-1} + \varphi\,(1-\omega s)\,\hat{c}_t^* - (1+\varphi)\,\hat{a}_t && (27)\\
&\quad + \left[(1-\omega)\,(1+\varphi\omega s) + \varphi\,(1-\omega^*)\,(1-\omega s)\right]\hat{\tau}_t \\
\hat{y}_t &= \left[(1-\omega)\,\omega s + (1-\omega^*)\,(1-\omega s)\right]\hat{\tau}_t + \omega s\hat{c}_t + (1-\omega s)\,\hat{c}_t^* && (28)\\
\hat{\varepsilon}_{p,t} &= \rho_p\hat{\varepsilon}_{p,t-1} + \eta_{p,t} && (29)\\
\hat{a}_t &= \rho_a\hat{a}_{t-1} + \eta_{a,t} && (30)
\end{aligned}$$

where $s = \bar{C}/\bar{Y}$ denotes the home consumption/output ratio at the steady state. We abstract here from the symmetric foreign block.

The terms of trade equation is given by:

$$\hat{\tau}_t = \frac{1}{\omega-\omega^*}\left[\frac{\sigma}{1-\gamma}\hat{c}_t - \frac{\gamma\sigma}{1-\gamma}\hat{c}_{t-1} - \frac{\sigma^*}{1-\gamma^*}\hat{c}_t^* + \frac{\gamma^*\sigma^*}{1-\gamma^*}\hat{c}_{t-1}^* + \hat{\varepsilon}_{p,t}^* - \hat{\varepsilon}_{p,t}\right]$$

Since our first objective is to estimate the model, it is closed by specifying an interest rate rule for each country. This raises some difficulties, because the policy rule is not based on micro-foundations as the other equations of the model. While monetary policies have been constrained by the Exchange Rate Mechanism of the EMU, the exact way this constraint has been implemented by central banks is not clear. In addition, over the sample used for estimation, it is likely that the policy rules have changed over time. Consequently, we proceed as follows. In a first step, we adopt a widely-accepted policy rule such that the nominal interest rate smoothly adjusts to the deviation of inflation to its steady-state value and to the deviation of domestic aggregate output to its natural value.[17] The log-linearized home feedback rules are then given by:[18]

$$\hat{\imath}_t = \psi_i\hat{\imath}_{t-1} + (1-\psi_i)\left[\psi_\pi\hat{\pi}_{H,t} + \psi_y\left(\hat{y}_t - \hat{y}_t^n\right)\right] + \hat{\varepsilon}_{i,t} \tag{31}$$

[16] $\hat{x}_t$ denotes the log-deviation of X_t from the steady-state value $\bar{X}$, i.e. $\hat{x}_t = \log\left(X_t/\bar{X}\right)$.

[17] We define the natural value of output as the value that would prevail in absence of any cost-push shocks, i.e. at the flexible-price equilibrium.

[18] A similar one exists for the foreign country.

where $(\hat{y}_t - \hat{y}_t^n)$ denotes the log-deviation of home output to its natural value, $\hat{\varepsilon}_{i,t}$ is the monetary policy shock, $\psi_i \in (0,1)$, $\psi_\pi > 1$ and $\psi_y > 0$.[19]

In a second step, we also investigated several alternative specifications for the policy rule to check the robustness of our estimates.[20] We also investigated whether a change in the sample period may affect the stability of our estimations. For this purpose, we re-estimated the model over more recent subperiods (from 1983:1 and from 1991:1). In all cases, we found that the parameters of the policy rule were not altered significantly. In the presentation of our results, we will only comment how our results are altered by a change in the sample period.

In the case of an area with more than two countries, the broad structure of the model remains essentially unchanged. The major change is that, in an N-country model, international transmission mechanisms pass through $(N-1)$ independent terms of trade. Consequently, since the Phillips curve depends on the terms of trade through movements in the real marginal cost, inflation dynamics is affected by demand conditions in all countries. Moreover, domestic consumption is affected by the average of real interest rates prevailing in all countries of the area. A complete description of adjustments that are necessary for the extension to more than two countries is provided in a separate technical appendix (available upon request).

2.6 Description of the tests of heterogeneity

In this section, we describe how the various tests of heterogeneity can be performed in our context. We first define, for each country k, the parameter sets. The structural parameters are denoted $\Theta_{Str}^k = (\gamma^k,\ \sigma^k,\ \varphi^k,\ \alpha^k,\ \xi^k)'$. Note that two additional parameters have been introduced in the model above, the discount factor β and the elasticity of substitution across goods produced within a given country θ. To reduce the estimation burden, we calibrate these parameters to their conventional values and we assume that they are equal for each country. Consequently, they are not explicitly introduced in the parameter set. The policy parameters are denoted $\Theta_{Pol}^k = (\psi_i^k,\ \psi_\pi^k,\ \psi_y^k)'$. We therefore assume that the characteristics of the monetary policy shocks are part of the stochastic parameters rather than part of the policy parameters. Finally, the parameters pertaining to shocks in country k are denoted $\Theta_{Sto}^k = (\rho_p^k,\ \rho_a^k,\ \rho_i^k,\ \sigma_p^k,\ \sigma_a^k,\ \sigma_i^k)'$ and parameters pertaining to the correlation matrix are denoted $\Theta_{Sto}^* = (\delta_p^{kj},\ \delta_a^{kj},\ \delta_i^{kj},\ k, j = 1, \cdots, N, j > k)'$, where δ_p^{kj}, δ_a^{kj}, and δ_i^{kj} denote the correlation between country-k and country-j preference, technology, and monetary policy

[19] We assume that $\hat{\varepsilon}_{i,t}$ follows an AR(1) process: $\hat{\varepsilon}_{i,t} = \rho_i \hat{\varepsilon}_{i,t-1} + \eta_{i,t}$. We also estimated a specification with a time-varying inflation objective and an i.i.d. monetary policy shock, along the lines of SW. As in Onatski and Williams (2004), however, we obtained that the variance of the monetary policy shock is essentially null. Consequently, we kept specification (31) that does not resort to a shock with a zero variance.

[20] In particular, we estimated a policy rule with the consumer price index π_t in place of the domestic inflation $\pi_{H,t}$. This is equivalent to introduce the exchange rate or the terms of trade in the reaction function of monetary authorities. We also tested specifications with a different timing for inflation or output in the policy rule. Alternatively, we estimated the model with and without detrending the inflation and interest rate.

shocks, respectively. We also define $\Theta = \left(\Theta_{Str}^{k\prime}, \Theta_{Pol}^{k\prime}, \Theta_{Sto}^{k\prime}, k = 1, \cdots, N, \Theta_{Sto}^{*\prime}\right)^{\prime}$ the vector of all unknown parameters in the model.

Now, we define the different hypotheses concerning heterogeneity as follows: the null hypothesis that the structure of the economy is the same in each country (structural homogeneity) is given by

$$H_0^{Str} : \Theta_{Str}^{k} = \Theta_{Str}^{j} \qquad \text{for all } k, j = 1, \cdots, N, j \neq k.$$

The null hypothesis that central banks respond in the same way to economic developments (policy homogeneity) is given by

$$H_0^{Pol} : \Theta_{Pol}^{k} = \Theta_{Pol}^{j} \qquad \text{for all } k, j = 1, \cdots, N, j \neq k.$$

We also define the joint hypothesis of both structural and policy homogeneity as

$$H_0^{StrPol} : H_0^{Str} \cap H_0^{Pol}$$

Finally, the null hypothesis of stochastic homogeneity is characterized by the fact that the same shocks affect the N national economies. It can be written as

$$H_0^{Sto} : \begin{cases} \Theta_{Sto}^{k} = \Theta_{Sto}^{j} & \text{for all } k, j = 1, \cdots, N, j \neq k \\ \Theta_{Sto,l}^{*} = 1 & \text{for all } l = 1, \cdots, N(N+1)/2. \end{cases}$$

The first equality implies that serial correlations and variances of shocks are equal for all countries, while the second equality implies that all cross-correlations of shocks between country-k and country-j are equal to one.[21] We also consider a "partial stochastic homogeneity", corresponding to the hypothesis that serial correlations and variances are equal for all countries. This hypothesis allows to identify whether the possible rejection of H_0^{Sto} is due to the fact that shocks have different univariate characteristics or to the fact that they are imperfectly correlated.

3 Empirical analysis

In this section, we describe the data used for the estimation of the MCM. Then, we briefly explain how the models are estimated using Bayesian econometrics and specify the priors for the distribution of parameters. Finally, we present the results of the estimation of the model.

[21] This hypothesis obviously reduces the number of shocks from $3N$ to only 3, and it raises the difficulty that the number of observable variables is not equal to the number of shocks anymore. To cope with this problem, we approximate the test that cross-correlations are equal to one by testing $\Theta_{Sto,l}^{*} = 0.995$ for all l and checking that the ex-post cross-correlations are actually very close to one.

3.1 Data

The MCM is estimated for the three largest countries of the area (Germany, France, and Italy). The sample period runs from 1970:1 to 1998:4 at a quarterly frequency. The data are drawn from OECD Business Sector Data Base for individual countries.[22] The estimation of the model is based ultimately on three key macroeconomic variables for each country: real consumption, the inflation rate, and the nominal short-term interest rate. Consumption is defined as real consumption expenditures, linearly detrended.[23] We measure inflation as the annualized quarterly percent change in the implicit GDP deflator. The interest rate is the three-month money-market rate. Figure 1 displays the historical path of the various series under consideration for each country. We observe a downward trend in inflation and interest rate, which mainly corresponds to the convergence process of economic conditions within the euro area. The structural model presented above is clearly not designed to capture such an empirical feature. Therefore, inflation and the nominal interest rate are detrended by the same quadratic trend in inflation.[24] It should be noticed that neither the terms of trade nor the real marginal cost are necessary for the estimation of model, since they are function of the other macroeconomic variables.

3.2 Econometric approach

For estimating the DSGE model described above, we adopt the Bayesian strategy proposed, among others, by Fernandez-Villaverde and Rubio-Ramirez (2003), Schorfheide (2003), and SW.[25] Most alternative approaches are precluded in our context. On one hand, calibration is not a promising avenue, because we focus on the effect of heterogeneity between countries within the euro area. The choice of distinct parameters for the various countries would be largely arbitrary, since the economic differences between these countries are not always clearly established. On the other hand, the FIML has proved to be rather tricky to implement in medium- or large-scale models, and resulted in an unrealistic estimation of some important structural parameters. In particular, the Calvo's probability α that a firm is not allowed to re-optimize its price was found to reach its upper bound of 1, a value that has to be ruled out for theoretical reasons. Consequently, we resort to the Bayesian technique to incorporate some prior information on structural parameters and render the estimation procedure more stable.

[22] Note that, in the case of Germany, we corrected for the mechanical impact of re-unification on GDP and GDP deflator data using data for West Germany for the year 1991.

[23] We also examined a detrended consumption computed using the regression on a quadratic time trend or a Hodrick-Prescott filter, and we obtained very similar results.

[24] We perform the same estimations without detrended inflation and interest rate, and we found essentially the same parameter estimates and the same conclusions for the tests of heterogeneity.

[25] Procedures to compute Bayesian econometrics are available in GAUSS software (see Schorfheide, 2003) and MATLAB software (the pre-processor DYNARE – developed by M. Juillard – includes now a module for estimation. See http://www.cepremap.cnrs.fr/dynare/).

Let $\hat{x}_t = \left(\hat{x}_t^k, k = 1, \cdots, N\right)$ be the vector of observable variables, where $\hat{x}_t^k = \left(\hat{c}_t^k, \hat{\pi}_{H,t}^k, \hat{\imath}_t^k\right)'$ contains the country-k observable variables (consumption, inflation and interest rate). The log-linearized model (25)–(31) is cast in a state-space representation for $\hat{x}_t$

$$\hat{x}_t = C(\Theta)\, \hat{s}_t \tag{32}$$

$$\hat{s}_t = A(\Theta)\, \hat{s}_{t-1} + B(\Theta)\, \eta_t \tag{33}$$

where $\hat{s}_t$ is the vector of state variables. In addition to observable variables, it includes unobservable variables such as marginal cost, natural output, terms of trade or shock processes. Last, η_t is a vector of i.i.d. variables with zero mean and covariance matrix $\Sigma(\Theta)$. The system matrices $C(\Theta)$, $A(\Theta)$, $B(\Theta)$ and $\Sigma(\Theta)$ are all functions of the parameter vector Θ.

A Kalman filter is used to estimate the system (32)–(33). The algorithm preliminary evaluates the number of explosive eigenvalues. Consequently, indeterminate models (that do not satisfy the Blanchard-Kahn conditions) are directly ruled out during the course of the estimation.

For a given structural model $\mathcal{M}_i$ and a set of parameters Θ, we denote $\Gamma(\Theta|\mathcal{M}_i)$ the prior distribution of Θ and $\mathcal{L}(X_T|\Theta, \mathcal{M}_i)$ the likelihood function associated to the observable variables $X_T = \{\hat{x}_t\}_{t=1}^T$. Then, from Bayes rule, the posterior distribution of the parameter vector is proportional to the product of the likelihood function and the prior distribution of Θ,

$$\Gamma(\Theta|X_T, \mathcal{M}_i) \propto \mathcal{L}(X_T|\Theta, \mathcal{M}_i)\, \Gamma(\Theta|\mathcal{M}_i). \tag{34}$$

Given the specification of the model, the posterior distribution cannot be recovered analytically. However, it can be evaluated numerically, using a Monte-Carlo Markov Chain (MCMC) sampling approach. More specifically, we rely to the Metropolis-Hastings (MH) algorithm to obtain a random draw of size 100,000 from the posterior distribution of the parameters.[26] The mode and the Hessian of the posterior distribution evaluated at the mode are used to initialize the MH algorithm.

3.3 Prior distribution

In this section, we describe how we selected the prior distribution for unknown parameters. In most cases, priors have been chosen to be very close to those adopted by SW for the euro area, but we also incorporate some information drawn from Onatski and Williams (2004).[27] Priors are reported in the first column of Table 1. The habit persistence parameter, γ, the fraction of firms that are not allowed to re-optimize their price, α, and the degree of price

[26] The first 50,000 observations are discarded to eliminate any dependence on the initial values.

[27] The latter authors provide an interesting investigation of some shortcomings of the standard Bayesian approach in the context of DSGE models. In particular, they put forward that parameter estimates are very sensitive to the way priors are introduced. In the estimation of the model, we took advantage of some of their results.

indexation, ξ, are assumed to follow a beta distribution, with a mean of 0.7 and a standard error of 0.1. The mean value of 0.7 is close to values found in other studies in the literature. The inverse of the inter-temporal elasticity of substitution of consumption, σ, and the inverse of the elasticity of labor disutility, φ, are assumed to follow a normal distribution, because they may theoretically take rather large values. They have a mean of 2 with a standard error of 0.25. This choice is based on evidence provided by Onatski and Williams (2004) who stress that these parameters may actually be larger than those reported by SW. Parameters pertaining to the monetary policy reaction function are standard: the long-term parameter on inflation ψ_π is 1.5 and the long-term parameter in output gap ψ_y is 0.5, with a standard error of 0.1, corresponding to the plain vanilla Taylor rule (they follow a normal distribution). The smoothing parameter ψ_i and persistence parameters (ρ_p, ρ_a, and ρ_i) are assumed to follow a beta distribution, with a mean of 0.7 and a standard error of 0.1. We opt for a prior uniform distribution between $[0, 2]$ for all standard deviations of the stochastic shocks, σ_p, σ_a, and σ_i.

While the shocks in a given country are assumed to be uncorrelated, we allow a non-zero correlation between a given shock in two countries. We thus denote δ_p, δ_a, and δ_i the correlations between domestic and foreign preference shocks, technology shocks, and monetary policy shocks, respectively. Correlations across countries have a normal distribution with a mean of 0.2 and a standard error of 0.1. We use the same priors for all countries and the euro area in turn.

Finally, we imposed dogmatic priors over the discount factor β and the elasticity of substitution across goods produced in a given country, θ. The values we use ($\beta = 0.99$ and $\theta = 10$) are conventional in the literature. The consumption/output ratio s is set equal to 1 for all countries, assuming that commercial trade is broadly balanced. The selection of the parameters of home bias in preferences (ω) is more tricky since the three countries under study are far from covering the whole external trade. We therefore set these parameters as follows, in order to reflect the weight of each country in the external trade of the others: the weights of German, French and Italian goods in the consumption of German households are $(0.8; 0.11; 0.09)$. For French and Italian households, the weights are $(0.13; 0.8; 0.07)$ and $(0.13; 0.07; 0.8)$ respectively. We checked that marginally altering these values would not change our results in any significant way.

3.4 Estimate of the complete MCM

We begin with the estimation of the complete MCM, since it is a benchmark for testing the different sources of heterogeneity. The joint dynamics of the whole system is estimated simultaneously for Germany, France, and Italy. This is actually a rather time-consuming task, since it involves 9 observable series and 51 unknown parameters. Table 2 provides two sets of information regarding parameter estimates. The first set reports the mean and the standard error of the posterior distribution of parameter estimates. The second set contains the 5, 50, and 95 percentiles of the posterior distribution of parameters. Figure 2 summarizes this information visually by plotting the prior and posterior distributions. As it appears clearly from the figure, the posterior distribution of φ and ψ_π is rather close to the prior distribution. This suggests that these parameters do not strongly affect the likelihood and translates in the rather large associated standard deviations. To check the robustness of our estimates, we also discuss the estimation results over the subperiod 1983-98. These estimates are reported in Table A of the Appendix.

The overall picture that emerges from the table is that the three countries display very similar parameter estimates. Yet, beyond this general finding, some differences are worth emphasizing. As regard the behavior of households, our estimates of the inverse of the consumption elasticity of substitution (σ) range between 1.5 and 2, while the inverse of the elasticity of labor disutility (φ) is equal to 2. Although we select the same priors for all countries, we obtain significant differences for the habit persistence parameter γ. This parameter is estimated to be rather low in Germany (0.63), medium in France (0.69), and large in Italy (0.78). We strongly reject the null hypothesis that the three parameters are equal across countries, suggesting that there is some heterogeneity of structural parameters across countries. These estimates differ slightly from the estimates reported in SW for the euro area. In particular, their estimate of the habit persistence parameter is found be rather low (at 0.6) and the estimate of the inverse of the consumption elasticity of substitution is as low as 1.4.

Turning to the behavior of firms, we obtain that the parameter of price indexation ξ ranges between 0.28 for Germany and 0.43 for Italy. Inspecting the table in the Appendix, we observe that the price indexation parameter has slightly increased over the recent period. This parameter estimate is found to be in the range between 0.4 and 0.5. In addition, the probability that firms are not allowed to re-optimize their price α is very close to 0.8. The degree of price stickiness is rather large, since the average duration of price contracts is about 5 quarters. This figure is somewhat larger than microeconomic evidence, but it is in the range of previous macroeconomic estimates.

Our estimate of the monetary policy rules should be considered as only indicative of how short-term interest rates reacted to macroeconomic developments over the sample period, because it is very likely that the behavior of monetary authorities has changed over the whole period. Reaction-function parameters display rather similar patterns across countries. The long-run reaction of short-term interest rate to inflation and output gap are about 1.5 and 0.5 respectively in the three countries. The interest rate persistence ψ_i is about 0.85, which is lower than the estimate reported by SW for the euro area.

The volatility of the preference and technology shocks are very close for the three countries. The former is within the range 0.05-0.06 while the latter is around 0.04. In contrast, some large differences in the variability of the monetary policy shock are found. While the volatility is low in Germany and Italy (around $0.23 - 0.24\%$), it is very large in France (at 0.42%). This result may be related to some aspects of the French monetary policy, not incorporated in the model, such as the implicit anchoring to the German monetary policy from 1983 on. When the sample period begins in 1983, the volatility of the French monetary policy shock turns out to decrease to 0.29%. Meanwhile, the volatility of the monetary policy shock in the two other countries also decreased to about 0.15%.

Concerning the serial correlation of shocks, the table reveals some significant differences across countries for the preference shock ($\rho_p = 0.51$ in France and 0.80 in Italy) and for the technology shock ($\rho_a = 0.66$ in France and 0.86 in Italy). In contrast, the estimates of ρ_i are all very close to 0.45. We notice that our estimates of the serial correlation of shocks are in general much smaller than those reported in SW (around 0.85). This result suggests that our structural model is able to reproduce most persistence in the data without resorting too heavily to the serial correlation of shocks.

Most cross-country correlations between shocks are significantly positive. Note however that shocks are far from being perfectly correlated across countries. This result is of importance, because it points to a significant source of transmission across the members of the area, suggesting some asymmetry of shocks across countries. When we turn to the 1983-98 subperiod, we obtain an increase in the correlation of some technology shocks, but it does not change the broad picture of a weak correlation between shocks across countries.

Overall, rather large differences in the parameter estimates are found between countries and the euro area as a whole as reported by SW. This suggests that the area-wide estimation of parameters describing the behavior of households may suffer from an aggregation bias. Such aggregation bias has already been pointed out as a possible undesirable outcome of estimating an AWM (Demertzis and Hugues Hallett, 1998).

4 Testing for the sources of heterogeneity

In this section, we first describe how assuming some form of homogeneity affect the parameter estimates of the model. Then we provide some model evaluation based on posterior odds. Finally, we present the results of the various tests concerning the heterogeneity across countries within the euro area based on loss functions.

4.1 Estimates of the constrained models

Table 3 reports the parameter estimates of the MCM when a given source of homogeneity is assumed to hold. First, we estimate an MCM in which *structural homogeneity* holds across countries. This model allows to test formally the null hypothesis that private agents behave in a similar manner in the three countries. Structural parameters are found to be rather close to the complete MCM for the utility function parameters ($\gamma = 0.79$, $\sigma = 1.89$ and $\varphi = 2.20$). Turning to the behavior of firms, our estimates reveal that the parameter of price indexation is significantly below the estimates obtained for the complete MCM, while other parameters are not significantly altered. Overall, this result suggests that, between core countries of the euro area, structural heterogeneity may be neglected at a first approximation.

Second, we estimate an MCM with *policy homogeneity*, so that monetary policy parameters are constant across countries.[28] The common policy rule has parameters equal to $\psi_i = 0.87$, $\psi_\pi = 1.43$ and $\psi_y = 0$. The major change with respect to the complete MCM is that the policy rule does not respond to output gap anymore. Imposing policy homogeneity also alters some structural parameters significantly. First the consumption habit parameter increases to about 0.9 for France and Italy. Second the Calvo probability rises to somewhat implausible values (at about 0.94). Also, the serial correlation of shocks remains at rather low levels (below 0.5), indicating that the persistence in the data is well captured by structural parameters. Finally, we notice a sharp increase in the volatility of the preference and technology shocks. This result may be interpreted as the sign that the constraints imposed to the model imply a loss of adequacy to the data, so that the hypothesis of policy homogeneity has some undesirable outcomes. One reason may be that the monetary rules have themselves suffered from temporal instability. Indeed, when this model is re-estimated over the period 1983-98, we observe that the habit parameter and the Calvo probability are estimated to more plausible values.

When we jointly assume structural and policy homogeneity, we do not observe significantly changes as compared to the model with policy homogeneity. This suggests that combining the two sets of constraints does not imply side effects that would worsen the estimation of structural parameters.

We turn now to the *stochastic homogeneity* hypothesis. As a first step, the partial stochastic homogeneity implies that volatility and serial-correlation parameters are assumed to be equal across countries. The volatility of preference and technology shocks is not significantly affected, while the volatility of the monetary policy shock increases in Germany and France. Turning to serial correlation of shocks, the preference shock turns out to be much more serially correlated under partial stochastic homogeneity. In addition, this hypothesis does not affect the estimation of structural parameters too markedly. Actually, the main change in the parameter estimates is the sharp decrease in the value of the habit parameter that is found to be around 0.5 in Germany and France. Also the Calvo probability decreases slightly in all countries. Finally we consider the full stochastic homogeneity, i.e. the correlation of shocks across countries is assumed to be very close to one. This hypothesis that only area-wide preference, technology and monetary-policy shocks affect the economies alters the dynamics of shocks dramatically. In particular, the preference shock is found to be essentially non-stationary, and the volatility of preference and technology shocks sharply increases, since they are multiplied by a factor 10.

[28] Notice that we consider the asymmetry in the monetary policy shocks as part of stochastic heterogeneity and not as part of policy heterogeneity.

4.2 Model evaluation based on posterior odds

A first way to evaluate the sources of heterogeneity in a Bayesian context relies on comparing the posterior distributions of the alternative models (see Geweke, 1999). Once the likelihood function and the prior distribution are given, the marginal likelihood of a given model $\mathcal{M}_i$ is obtained using the following expression

$$\mathcal{L}\left(X_T|\mathcal{M}_i\right)=\int_{\Theta}\mathcal{L}\left(X_T|\Theta,\mathcal{M}_i\right)\Gamma\left(\Theta|\mathcal{M}_i\right)\mathrm{d}\Theta. \tag{35}$$

Multiple integration is required to obtain the marginal likelihood, making exact computation infeasible. However, using random draws from the posterior distribution, it is possible to evaluate the expression (35) numerically, as shown in Geweke (1999). An advantage of the measure is that it accounts for the model dimensionality and adjusts for the effect of the prior distribution (Chang, Gomez, and Schorfheide, 2002).

Let $\hat{\mathcal{L}}\left(X_T|\mathcal{M}_i\right)$ denote the estimated marginal likelihood of model $\mathcal{M}_i$. Then, we compute the Bayes factor between two models $\mathcal{M}_i$ and $\mathcal{M}_j$ as

$$\mathcal{B}_{i,j}\left(X_T\right)=\frac{\hat{\mathcal{L}}\left(X_T|\mathcal{M}_i\right)}{\hat{\mathcal{L}}\left(X_T|\mathcal{M}_j\right)}.$$

Now, assume that there are $m+1$ competing models $\mathcal{M}_i$ for $i=0,\cdots,m$, with $\mathcal{M}_0$ denoting the reference model. If we denote $\mathcal{P}_{i,0}$ the prior probability of model $\mathcal{M}_i$ (with $\sum_{j=0}^{m}\mathcal{P}_{j,0}=1$), we obtain the posterior odds, which incorporates information on priors, as follows:

$$\mathcal{PO}_{i,T}=\frac{\mathcal{P}_{i,0}\hat{\mathcal{L}}\left(X_T|\mathcal{M}_i\right)}{\sum_{j=0}^{m}\mathcal{P}_{j,0}\hat{\mathcal{L}}\left(X_T|\mathcal{M}_j\right)}.$$

A widely-accepted approach to assess the empirical performances of an estimated DSGE model relies on the comparison of the DSGE models with an a-theoretical VAR model.[29] Such a reference to a VAR model is rather natural, since the reduced form of log-linearized DSGE models can be viewed as a constrained VAR model. Thus, the test is based on whether the constraints imposed by the DSGE models to the VAR model are rejected by the data.

Such an empirical assessment is performed in Table 4, which reports, for the DSGE models and the competing VAR(1) model, the prior probabilities $\mathcal{P}_{i,0}$, the marginal likelihood $\hat{\mathcal{L}}\left(X_T|\mathcal{M}_i\right)$, the Bayes factor $\mathcal{B}_{i,j}\left(X_T\right)$ relative to the VAR model and the posterior odds $\mathcal{PO}_{i,T}$. We selected a VAR model with one lag only, because the number of unknown parameters in a VAR model is very large in our multi-country context.[30] Since the marginal likelihood cannot be computed analytically due to the complexity of the model, it is approximated using the simulation-based modified harmonic mean estimator proposed by Geweke (1999). We assign a prior probability of 1/7 to the seven models under consideration.

[29] See, e.g., Fernandez-Villaverde and Rubio-Ramirez (2003), Schorfheide (2000) or Chang, Gomes, and Schorfheide (2002).

[30] The number of parameters is as high as 126 for a VAR model with one lag and goes up to 207 with two lags.

The first result we observe from the table is that the VAR model overwhelmingly dominates DSGE models. This empirical evidence indicates that the data do not support some strong restrictions imposed by the DSGE models, as compared to VAR models. A similar observation was made by Schorfheide (2000). Another interesting result is the ranking of the DGSE models themselves. As it appears clearly in the table, in contrast, the complete MCM does not dominate all nested models that allow some homogeneity. This result shows up in the Bayes factors that markedly favor the models with structural and policy homogeneity. Of all DSGE models, the best model is found to be the model in which structural as well as policy homogeneity are assumed. On the other hand, the hypothesis of stochastic homogeneity is very strongly rejected in terms of Bayes factor.

4.3 Model evaluation based on loss functions

It may be argued that a reason for the poor performance of DSGE models is that they are not designed to capture all statistical characteristics in the data. In particular, the joint dynamics of shocks is overly constrained, since most cross-correlations are imposed to be zero. In addition, it imposes several constraints on the contemporaneous relationships between model variables, in particular across countries. For instance, international transmission mechanisms do not involve any additional estimated parameter as compared to a close economy set-up. Another important explanation for the poor ability of the DSGE models to reproduce some features of the actual data is the likely occurrence of structural shifts over the sample period. In particular, it may be argued that monetary policy rules have experienced significant changes during this period. From this point of view, a more relevant way to assess the performances of the DSGE model would be whether it is able to replicate some important stylized facts estimated on actual data.

Now, we adopt the approach recently proposed by Schorfheide (2000), and Chang, Gomes, and Schorfheide (2002) to compare the performance of (non-nested) DGSE models. The idea is to evaluate the ability of the competing models to reproduce some characteristics of the data, such as impulse response functions or cross-covariance functions. If one focuses on a given characteristic of the model (say, the cross-covariance functions), one can measure the ability of each model to reproduce this characteristic using for instance a quadratic loss function,

$$L_q\left(\Lambda, \hat{\Lambda}_i\right) = \left(\Lambda - \hat{\Lambda}_i\right)' W \left(\Lambda - \hat{\Lambda}_i\right)$$

where Λ denote the population characteristics, $\hat{\Lambda}_i$ the prediction of model $\mathcal{M}_i$ and W a positive definite weighting matrix. Schorfheide (2000) shows that the ranking of the DSGE model prediction $\hat{\Lambda}_i$ depends only on the value of the risk measure

$$R_q\left(\hat{\Lambda}_i|X_T\right) = \left(\hat{\Lambda}_i - \mathbb{E}\left(\Lambda|X_T\right)\right)' W \left(\hat{\Lambda}_i - \mathbb{E}\left(\Lambda|X_T\right)\right).$$

The posterior distribution of the population characteristics Λ is given by

$$\Pr(\Lambda | X_T) = \sum_{j=1}^{m} \mathcal{PO}_{i,T} \times \Pr(\Lambda | X_T, \mathcal{M}_i)$$

where $\Pr(\Lambda | X_T, \mathcal{M}_i)$ denotes the posterior distribution of Λ under $\mathcal{M}_i$. When the reference model dominates all other models, such that $\mathcal{PO}_{0,T} = 1$, then the posterior distribution of the population characteristics reduces to $\Pr(\Lambda | X_T) = \Pr(\Lambda | X_T, \mathcal{M}_0)$. In addition, under a quadratic loss function, it can be shown that the prediction $\hat{\Lambda}_i$ is simply given by the posterior mean $\hat{\Lambda}_i = \mathbb{E}(\Lambda | X_T, \mathcal{M}_i)$ of Λ under model $\mathcal{M}_i$. Finally, the weighting matrix W is chosen to be the inverse of the covariance matrix of the population characteristics Λ.

Table 5 reports the loss functions evaluated for the cross-covariance functions of all observable variables computed over 20 quarters.[31] The population cross-covariance functions are given by the VAR(1) since its posterior probability is equal to one. The first row of the table gives the value of the overall loss function for each competing DSGE model. Then, the global loss function is broken down by country in order to get a better diagnosis on the ability of the different models to reproduce the characteristics of the various economies. Interestingly, the model that performs worst is the model with partial stochastic homogeneity, since it is simply unable to reproduce the cross-covariance functions of the VAR model. Among other models, the complete MCM does not perform very well. Since this is the less constrained model, this finding suggests that its additional degrees of freedom do not help in reproducing the characteristics of the VAR model. When structural homogeneity is assumed, no improvement in the loss function is obtained. In case of policy homogeneity, one observes a clear improvement, that mainly comes from German cross-covariances and from the interactions of shocks across countries. Finally, the best results are obtained for the model with both structural and policy homogeneity that yields the lowest loss function for each country. This evidence drawn from the loss functions is fully consistent with the analysis based on the posterior odds. A similar result is found when the estimation is performed over the 1983-98 subperiod. In this case, the model with structural homogeneity performs much better than the complete MCM and even than the model with policy homogeneity. This result may be interpreted as a narrowing of the structural parameters across countries over the recent period. In the same spirit, we also observe that the model with partial stochastic homogeneity now yields more plausible results. Although it is still the worst model, it does not seem out of the running anymore. This suggests that the shocks may now have closer characteristics across countries. Nevertheless, the correlation of shocks across countries is still too low to be consistent with full stochastic homogeneity.

[31] We focus here on the cross-covariance functions rather than on impulse response functions, because cross-covariances can be identified from the reference VAR model without any additional assumptions. In the case of impulse response functions, additional restrictions have to be imposed on the VAR model to guarantee identification.

Our results differ to some extent from the recent evidence provided by Verhoef (2003). This author investigates in a structural VAR framework the extent of the asymmetry of shocks within the European Union. He obtains that supply and demand shocks in France and Germany are strongly correlated, while the correlations with shocks in Italy are much weaker. On one hand, our estimations of the cross-correlation of preference and technology shocks are very low (below 0.3), but in addition, we do not observe sizeable differences in the cross-correlations across countries. More specifically, France appears to be the most correlated with the other two countries, but the differences between cross-correlations do not turn out to be significant.

5 Conclusion

In this paper, we investigate the sources of heterogeneity within the euro area. To address this issue, we develop a multi-country DSGE model, which can be used to estimate the dynamics of national economies within the euro area. This model incorporates frictions required to reproduce the persistence in the actual data, including the presence of sticky-price setting and external habit formation in consumption. An additional characteristic of the model is the introduction of heterogenous behaviors across countries that allows to explicitly test the hypothesis of heterogeneity of behaviors across countries. The complete MCM is estimated on German, French, and Italian data, using Bayesian techniques. We provide evidence that the structural and policy parameters in these countries display rather limited differences. In contrast, the hypothesis of stochastic homogeneity is very strongly rejected by the data.

The overall picture that emerges of our investigation is that a parsimonious DSGE model for modeling core countries within the euro area is a model that assumes structural and policy homogeneity. Another way to summarize our results is that heterogeneity within the euro area mainly comes from stochastic heterogeneity, i.e. from the unexplained part of the model (Smets and Wouters, 2004, draw a similar conclusion for the US and the euro area). Our joint modeling of the three economies allows us to be more precise on the source of heterogeneity. Indeed although preference and technology shocks have, to some extent, very similar properties, they are only very weakly correlated across countries. A consequence is that business cycle fluctuations are not likely to be synchronized within the euro area, even between core countries. It is worth emphasizing that our evaluation is based on the three largest countries of the area, that may be viewed as very similar economies. It is likely that including additional economies would widen the discrepancies between the structural as well as policy parameters. In addition, subsample analysis does not suggest that the correlation between country-specific shocks may have increased over the recent period. By the way, our results eventually indicate that an aggregated model for the euro area would reflect the business cycle fluctuations within the euro area in a very imperfect way.

A clear limitation of our results is that the model described in this paper does not incorporate some important mechanisms, in particular concerning the fiscal policy and the structure of the labour market. Indeed, such missing components are likely to under-estimate the actual structural heterogeneity between the countries, while over-estimating the actual stochastic heterogeneity, since non-modeled factors are incorporated in shocks. This issue is left for future research.

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Table 1: Prior distribution for parameters

		Type	Mean	Std.error
Consumption habit	γ	Beta	0.700	0.100
Consumption elasticity of substitution	σ	Normal	2.000	0.250
Labour desutility	φ	Normal	2.000	0.250
Price indexation	ξ	Beta	0.700	0.100
Calvo probability	α	Beta	0.700	0.100
Policy rule: lagged interest rate	ψ_i	Beta	0.700	0.100
Policy rule: inflation	ψ_π	Normal	1.500	0.100
Policy rule: output gap	ψ_y	Normal	0.500	0.100
Volatility of preference shock	σ_p	Uniform	0.000	2.000
Volatility of productivity shock	σ_a	Uniform	0.000	2.000
Volatility of monetary policy shock (x100)	σ_i	Uniform	0.000	2.000
Serial correlation of preference shock	ρ_p	Beta	0.700	0.100
Serial correlation of productivity shock	ρ_a	Beta	0.700	0.100
Serial correlation of monetary policy shock	ρ_i	Beta	0.700	0.100
Cross-correlation of preference shocks	δ_p	Normal	0.200	0.100
Cross-correlation of productivity shocks	δ_a	Normal	0.200	0.100
Cross-correlation of monetary policy shocks	δ_i	Normal	0.200	0.100

Table 2: Posterior distribution of parameter estimates for the complete MCM

		Germany					France					Italy				
		Mean	Std dev.	5%	Median	95%	Mean	Std dev.	5%	Median	95%	Mean	Std dev.	5%	Median	95%
Consumption habit	γ	0.630	0.050	0.553	0.632	0.714	0.688	0.045	0.617	0.691	0.765	0.777	0.029	0.730	0.777	0.823
Consumption elast. of subst.	σ	1.542	0.232	1.162	1.533	1.922	1.851	0.226	1.482	1.851	2.228	2.009	0.218	1.656	2.009	2.373
Labour desutility	φ	1.934	0.253	1.522	1.929	2.349	2.015	0.252	1.595	2.019	2.428	1.922	0.247	1.511	1.919	2.316
Price indexation	ξ	0.290	0.078	0.157	0.283	0.406	0.324	0.083	0.191	0.318	0.455	0.436	0.102	0.257	0.428	0.593
Calvo probability	α	0.839	0.019	0.809	0.840	0.869	0.822	0.017	0.794	0.823	0.848	0.794	0.022	0.759	0.795	0.830
Policy rule: lagged interest rate	ψ_i	0.871	0.020	0.841	0.873	0.901	0.820	0.027	0.778	0.822	0.864	0.906	0.014	0.885	0.908	0.929
Policy rule: inflation	ψ_π	1.507	0.100	1.340	1.510	1.666	1.517	0.101	1.353	1.518	1.681	1.497	0.094	1.344	1.500	1.648
Policy rule: output gap	ψ_y	0.458	0.104	0.288	0.462	0.627	0.482	0.102	0.314	0.480	0.645	0.522	0.091	0.375	0.522	0.670
Vol. preference shock	σ_p	0.048	0.008	0.035	0.047	0.061	0.063	0.010	0.047	0.062	0.078	0.055	0.008	0.043	0.054	0.068
Vol. productivity shock	σ_a	0.037	0.006	0.026	0.036	0.047	0.038	0.007	0.028	0.038	0.050	0.035	0.006	0.026	0.035	0.045
Vol. mon. policy shock (x100)	σ_i	0.244	0.020	0.211	0.243	0.276	0.426	0.034	0.372	0.423	0.482	0.228	0.021	0.196	0.226	0.261
Serial-corr. preference shock	ρ_p	0.640	0.065	0.531	0.643	0.741	0.509	0.077	0.380	0.510	0.633	0.793	0.036	0.739	0.795	0.851
Serial-corr. productivity shock	ρ_a	0.740	0.067	0.635	0.741	0.854	0.660	0.075	0.536	0.661	0.780	0.854	0.035	0.796	0.855	0.911
Serial-corr. mon. policy shock	ρ_i	0.506	0.067	0.395	0.508	0.617	0.447	0.067	0.337	0.445	0.557	0.414	0.071	0.300	0.412	0.534
Cross-correlations across countries																
Preference shock - 1/2	$\delta_{p\,12}$	0.311	0.063	0.201	0.313	0.410										
Preference shock - 1/3	$\delta_{p\,13}$	0.166	0.067	0.059	0.168	0.273										
Preference shock - 2/3	$\delta_{p\,23}$	0.279	0.071	0.166	0.279	0.397										
Productivity shock - 1/2	$\delta_{a\,12}$	0.194	0.067	0.077	0.196	0.300										
Productivity shock - 1/3	$\delta_{a\,13}$	-0.032	0.076	-0.156	-0.032	0.096										
Productivity shock - 2/3	$\delta_{a\,23}$	0.135	0.075	0.018	0.138	0.258										
Monetary policy shock - 1/2	$\delta_{i\,12}$	0.198	0.070	0.087	0.200	0.317										
Monetary policy shock - 1/3	$\delta_{i\,13}$	0.124	0.066	0.016	0.127	0.229										
Monetary policy shock - 2/3	$\delta_{i\,23}$	0.239	0.069	0.132	0.237	0.355										

Table 3: Posterior distribution of parameter estimates under alternative hypotheses

		Complete MCM		Structural hom.		Policy hom.		Struct+Pol hom.		Partial stoch. hom.		Stochast. hom.	
		Mean	Std dev.	Mean	Std dev.	Mean	Std dev.	Mean	Std dev.	Mean	Std dev.	Mean	Std dev.
Germany													
Consumption habit	γ	0.630	0.050	0.792	0.029	0.759	0.045	0.885	0.018	0.479	0.042	0.513	0.039
Consumption elast. of subst.	σ	1.542	0.232	1.894	0.218	2.056	0.221	2.278	0.223	1.358	0.194	1.901	0.192
Labour desutility	φ	1.934	0.253	2.198	0.231	1.882	0.244	1.915	0.228	1.928	0.217	1.567	0.235
Price indexation	ξ	0.290	0.078	0.151	0.037	0.395	0.092	0.206	0.047	0.425	0.111	0.382	0.082
Calvo probability	α	0.839	0.019	0.877	0.013	0.928	0.010	0.950	0.007	0.667	0.047	0.743	0.025
Policy rule: lagged interest rate	ψ_i	0.871	0.020	0.886	0.017	0.870	0.015	0.875	0.014	0.705	0.039	0.749	0.034
Policy rule: inflation	ψ_π	1.507	0.100	1.499	0.102	1.427	0.105	1.361	0.105	1.705	0.076	1.601	0.066
Policy rule: output gap	ψ_y	0.458	0.104	0.361	0.119	0.005	0.005	0.003	0.003	0.544	0.096	0.498	0.092
Vol. preference shock	σ_p	0.048	0.008	0.093	0.014	0.091	0.017	0.191	0.031	0.059	0.010	0.944	0.093
Vol. productivity shock	σ_a	0.037	0.006	0.054	0.010	0.191	0.052	0.314	0.080	0.020	0.002	0.202	0.023
Vol. mon. policy shock (x100)	σ_i	0.244	0.020	0.233	0.019	0.213	0.015	0.210	0.013	0.455	0.033	4.058	0.306
Serial-corr. preference shock	ρ_p	0.640	0.065	0.408	0.070	0.511	0.083	0.310	0.061	0.947	0.014	0.997	0.000
Serial-corr. productivity shock	ρ_a	0.740	0.067	0.671	0.067	0.362	0.076	0.415	0.069	0.872	0.023	0.612	0.039
Serial-corr. mon. policy shock	ρ_i	0.506	0.067	0.570	0.059	0.435	0.059	0.450	0.063	0.356	0.048	0.476	0.048
France													
Consumption habit	γ	0.688	0.045	*0.792*	-	0.898	0.025	*0.885*	-	0.453	0.039	0.476	0.039
Consumption elast. of subst.	σ	1.851	0.226	*1.894*	-	2.161	0.232	*2.278*	-	1.651	0.190	2.214	0.190
Labour desutility	φ	2.015	0.252	*2.198*	-	1.974	0.250	*1.915*	-	1.973	0.238	1.684	0.236
Price indexation	ξ	0.324	0.083	*0.151*	-	0.378	0.084	*0.206*	-	0.442	0.116	0.362	0.090
Calvo probability	α	0.822	0.017	*0.877*	-	0.943	0.009	*0.950*	-	0.648	0.039	0.730	0.027
Policy rule: lagged interest rate	ψ_i	0.820	0.027	0.825	0.027	*0.870*	-	*0.875*	-	0.688	0.041	0.665	0.040
Policy rule: inflation	ψ_π	1.517	0.101	1.497	0.099	*1.427*	-	*1.361*	-	1.487	0.078	1.439	0.055
Policy rule: output gap	ψ_y	0.482	0.102	0.303	0.118	*0.005*	-	*0.003*	-	0.383	0.099	0.239	0.097
Vol. preference shock	σ_p	0.063	0.010	0.089	0.012	0.188	0.042	0.176	0.029	*0.059*	-	*0.944*	-
Vol. productivity shock	σ_a	0.038	0.007	0.059	0.012	0.330	0.065	0.374	0.099	*0.020*	-	*0.202*	-
Vol. mon. policy shock (x100)	σ_i	0.426	0.034	0.427	0.035	0.365	0.024	0.364	0.025	*0.455*	-	*4.058*	-
Serial-corr. preference shock	ρ_p	0.509	0.077	0.402	0.071	0.271	0.061	0.292	0.063	*0.947*	-	*0.997*	-
Serial-corr. productivity shock	ρ_a	0.660	0.075	0.641	0.066	0.409	0.071	0.468	0.066	*0.872*	-	*0.612*	-
Serial-corr. mon. policy shock	ρ_i	0.447	0.067	0.515	0.080	0.337	0.057	0.326	0.058	*0.356*	-	*0.476*	-

Table 3: (end)

		Complete MCM		Structural hom.		Policy hom.		Struct.+Pol. hom.		Partial stoch. hom.		Stochast. hom.	
		Mean	Std dev.	Mean	Std dev.	Mean	Std dev.	Mean	Std dev.	Mean	Std dev.	Mean	Std dev.
Italy													
Consumption habit	γ	0.777	0.029	*0.792*	-	0.903	0.022	*0.885*	-	0.695	0.031	0.697	0.030
Consumption elast. of subst.	σ	2.009	0.218	*1.894*	-	2.040	0.235	*2.278*	-	1.741	0.189	2.421	0.193
Labour desutility	φ	1.922	0.247	*2.198*	-	1.995	0.247	*1.915*	-	1.999	0.220	1.812	0.263
Price indexation	ξ	0.436	0.102	*0.151*	-	0.465	0.100	*0.206*	-	0.421	0.100	0.379	0.089
Calvo probability	α	0.794	0.022	*0.877*	-	0.935	0.011	*0.950*	-	0.646	0.034	0.697	0.028
Policy rule: lagged interest rate	ψ_i	0.906	0.014	0.902	0.018	*0.870*	-	*0.875*	-	0.814	0.028	0.769	0.026
Policy rule: inflation	ψ_π	1.497	0.094	1.466	0.101	*1.427*	-	*1.361*	-	1.642	0.082	1.553	0.065
Policy rule: output gap	ψ_y	0.522	0.091	0.226	0.087	*0.005*	-	*0.003*	-	0.538	0.111	0.495	0.088
Vol. preference shock	σ_p	0.055	0.008	0.058	0.007	0.116	0.027	0.105	0.017	*0.059*	-	*0.944*	-
Vol. productivity shock	σ_a	0.035	0.006	0.054	0.011	0.271	0.095	0.322	0.090	*0.020*	-	*0.202*	-
Vol. mon. policy shock (x100)	σ_i	0.228	0.021	0.231	0.025	0.227	0.018	0.222	0.017	*0.455*	-	*4.058*	-
Serial-corr. preference shock	ρ_p	0.793	0.036	0.812	0.034	0.688	0.058	0.729	0.046	*0.947*	-	*0.997*	-
Serial-corr. productivity shock	ρ_a	0.854	0.035	0.815	0.038	0.532	0.084	0.638	0.061	*0.872*	-	*0.612*	-
Serial-corr. mon. policy shock	ρ_i	0.414	0.071	0.466	0.088	0.510	0.073	0.493	0.068	*0.356*	-	*0.476*	-
Cross-correlations across countries													
Preference shock - 1/2	$\delta_{p\,12}$	0.311	0.063	0.303	0.066	0.272	0.064	0.280	0.065	0.674	0.046	*0.995*	-
Preference shock - 1/3	$\delta_{p\,13}$	0.166	0.067	0.147	0.069	0.136	0.065	0.112	0.061	0.617	0.063	*0.995*	-
Preference shock - 2/3	$\delta_{p\,23}$	0.279	0.071	0.261	0.066	0.190	0.067	0.192	0.066	0.597	0.061	*0.995*	-
Productivity shock - 1/2	$\delta_{a\,12}$	0.194	0.067	0.221	0.073	0.161	0.067	0.167	0.072	0.562	0.056	*0.995*	-
Productivity shock - 1/3	$\delta_{a\,13}$	-0.032	0.076	-0.012	0.068	-0.006	0.069	0.016	0.071	0.511	0.040	*0.995*	-
Productivity shock - 2/3	$\delta_{a\,23}$	0.135	0.075	0.161	0.072	0.187	0.075	0.201	0.071	0.513	0.058	*0.995*	-
Monetary policy shock - 1/2	$\delta_{i\,12}$	0.198	0.070	0.211	0.069	0.274	0.066	0.265	0.066	0.608	0.042	*0.995*	-
Monetary policy shock - 1/3	$\delta_{i\,13}$	0.124	0.066	0.132	0.069	0.148	0.066	0.144	0.067	0.494	0.059	*0.995*	-
Monetary policy shock - 2/3	$\delta_{i\,23}$	0.239	0.069	0.243	0.064	0.226	0.070	0.238	0.067	0.577	0.041	*0.995*	-

Table 4: Performance evaluation

	Complete MCM	Structural hom.	Policy hom.	Struct.+Pol. hom.	Partial stoch hom.	Stochast. hom.	VAR(1) model
Panel A: 1970-98 sample							
Prior probability	*0.143*	*0.143*	*0.143*	*0.143*	*0.143*	*0.143*	*0.143*
Marginal likelihood	3971.93	3985.00	3993.33	4017.55	3819.39	3377.23	4088.99
Bayes factor	1	473923	2.0E+09	6.5.E+19	5.6.E-67	5.3E-259	6.9.E+50
Posterior odds	1.5E-51	6.9E-46	2.9E-42	9.4E-32	8.2E-118	0.0E+00	1
Panel B: 1983-98 sample							
Prior probability	*0.143*	*0.143*	*0.143*	*0.143*	*0.143*	*0.143*	*0.143*
Marginal likelihood	2269.88	2291.18	2261.44	2302.51	2156.78	1971.81	2400.71
Bayes factor	1	1.8E+09	2.2E-04	1.5.E+14	7.6.E-50	3.5E-130	6.6.E+56
Posterior odds	1.5E-57	2.7E-48	3.3E-61	2.3E-43	1.2E-106	5.4E-187	1

Table 5: Loss function based on cross-covariance functions

	Complete MCM	Structural hom.	Policy hom.	Struct.+Pol. hom.	Partial stoch. hom.
Panel A: 1970-98 sample					
Overall	14.79	14.82	12.44	**10.61**	1661.44
Germany	3.12	3.46	2.03	**1.29**	515.91
France	2.63	2.76	2.66	**2.28**	77.82
Italy	0.93	0.58	0.85	**0.51**	17.75
Cross countries	8.11	8.02	6.89	**6.53**	1049.97
Panel B: 1983-98 sample					
Overall	11.35	7.17	7.97	**6.85**	24.54
Germany	1.50	0.64	0.40	**0.24**	8.31
France	2.09	**1.05**	2.16	1.45	2.01
Italy	1.88	0.64	0.71	**0.54**	1.65
Cross countries	5.87	4.84	4.70	**4.63**	12.57

Table A: Posterior distribution of parameter estimates under alternative hypotheses (1983-1998)

		Complete MCM		Structural hom.		Policy hom.		Struct+Pol hom.		Partial stoch. hom.		Stochast. hom.	
		Mean	Std dev.	Mean	Std dev.	Mean	Std dev.	Mean	Std dev.	Mean	Std dev.	Mean	Std dev.
Germany													
Consumption habit	γ	0.620	0.052	0.728	0.037	0.707	0.042	0.818	0.030	0.628	0.046	0.501	0.038
Consumption elast. of subst.	σ	1.479	0.251	1.537	0.271	1.919	0.251	2.056	0.221	1.202	0.197	1.826	0.232
Labour desutility	φ	1.974	0.243	2.102	0.243	1.906	0.244	1.936	0.267	1.959	0.243	1.818	0.236
Price indexation	ξ	0.480	0.114	0.205	0.050	0.486	0.128	0.236	0.060	0.498	0.111	0.569	0.094
Calvo probability	α	0.824	0.022	0.855	0.016	0.901	0.011	0.930	0.008	0.786	0.023	0.689	0.032
Policy rule: lagged interest rate	ψ_i	0.928	0.013	0.927	0.015	0.878	0.014	0.883	0.013	0.912	0.017	0.806	0.026
Policy rule: inflation	ψ_π	1.508	0.106	1.497	0.100	1.419	0.089	1.444	0.105	1.496	0.100	1.562	0.067
Policy rule: output gap	ψ_y	0.448	0.102	0.416	0.101	0.006	0.005	0.005	0.005	0.509	0.097	0.501	0.086
Vol. preference shock	σ_p	0.051	0.010	0.069	0.014	0.077	0.014	0.127	0.022	0.050	0.007	0.878	0.100
Vol. productivity shock	σ_a	0.030	0.005	0.036	0.007	0.091	0.010	0.136	0.032	0.032	0.005	0.133	0.018
Vol. mon. policy shock (x100)	σ_i	0.127	0.014	0.127	0.014	0.117	0.013	0.117	0.013	0.263	0.019	2.661	0.224
Serial-corr. preference shock	ρ_p	0.671	0.070	0.565	0.070	0.560	0.059	0.447	0.068	0.646	0.061	0.997	0.001
Serial-corr. productivity shock	ρ_a	0.716	0.067	0.697	0.074	0.440	0.074	0.464	0.075	0.762	0.056	0.405	0.059
Serial-corr. mon. policy shock	ρ_i	0.537	0.073	0.576	0.068	0.444	0.085	0.515	0.085	0.282	0.062	0.434	0.067
France													
Consumption habit	γ	0.678	0.042	*0.728*	-	0.825	0.036	*0.818*	-	0.578	0.045	0.545	0.041
Consumption elast. of subst.	σ	1.796	0.240	*1.537*	-	2.038	0.209	*2.056*	-	1.712	0.230	2.035	0.189
Labour desutility	φ	1.957	0.266	*2.102*	-	2.016	0.217	*1.936*	-	1.972	0.229	1.954	0.251
Price indexation	ξ	0.428	0.108	*0.205*	-	0.488	0.129	*0.236*	-	0.370	0.097	0.487	0.095
Calvo probability	α	0.835	0.018	*0.855*	-	0.928	0.009	*0.930*	-	0.810	0.021	0.732	0.030
Policy rule: lagged interest rate	ψ_i	0.854	0.026	0.846	0.031	*0.878*	-	*0.883*	-	0.858	0.024	0.668	0.048
Policy rule: inflation	ψ_π	1.489	0.093	1.515	0.105	*1.419*	-	*1.444*	-	1.500	0.098	1.402	0.057
Policy rule: output gap	ψ_y	0.493	0.103	0.467	0.090	*0.006*	-	*0.005*	-	0.516	0.092	0.188	0.086
Vol. preference shock	σ_p	0.055	0.008	0.056	0.011	0.099	0.014	0.098	0.015	*0.050*	-	*0.878*	-
Vol. productivity shock	σ_a	0.036	0.007	0.037	0.007	0.149	0.021	0.152	0.037	*0.032*	-	*0.133*	-
Vol. mon. policy shock (x100)	σ_i	0.294	0.028	0.297	0.029	0.261	0.020	0.264	0.023	*0.263*	-	*2.661*	-
											-		-
Serial-corr. preference shock	ρ_p	0.532	0.069	0.500	0.072	0.426	0.087	0.438	0.079	*0.646*	-	*0.997*	-
Serial-corr. productivity shock	ρ_a	0.513	0.096	0.533	0.096	0.368	0.067	0.374	0.072	*0.762*	-	*0.405*	-
Serial-corr. mon. policy shock	ρ_i	0.389	0.072	0.424	0.076	0.297	0.055	0.297	0.060	*0.282*	-	*0.434*	-

Table A: (end)

		Complete MCM		Structural hom.		Policy hom.		Struct.+Pol. hom.		Partial stoch. hom.		Stochast. hom.	
		Mean	Std dev.	Mean	Std dev.	Mean	Std dev.	Mean	Std dev.	Mean	Std dev.	Mean	Std dev.
Italy													
Consumption habit	γ	0.744	0.041	*0.728*	-	0.853	0.025	*0.818*	-	0.732	0.036	0.638	0.039
Consumption elast. of subst.	σ	1.759	0.246	*1.537*	-	1.851	0.213	*2.056*	-	1.472	0.208	2.331	0.205
Labour desutility	φ	2.034	0.262	*2.102*	-	1.891	0.270	*1.936*	-	1.976	0.251	1.815	0.221
Price indexation	ξ	0.494	0.107	*0.205*	-	0.449	0.072	*0.236*	-	0.484	0.111	0.479	0.097
Calvo probability	α	0.790	0.028	*0.855*	-	0.916	0.008	*0.930*	-	0.768	0.027	0.690	0.030
Policy rule: lagged interest rate	ψ_i	0.911	0.021	0.917	0.019	*0.878*	-	*0.883*	-	0.915	0.015	0.808	0.024
Policy rule: inflation	ψ_π	1.511	0.101	1.506	0.094	*1.419*	-	*1.444*	-	1.509	0.102	1.619	0.072
Policy rule: output gap	ψ_y	0.482	0.095	0.370	0.103	*0.006*	-	*0.005*	-	0.518	0.092	0.511	0.087
Vol. preference shock	σ_p	0.048	0.010	0.043	0.008	0.079	0.013	0.073	0.012	*0.050*	-	*0.878*	-
Vol. productivity shock	σ_a	0.026	0.005	0.036	0.008	0.122	0.008	0.137	0.033	*0.032*	-	*0.133*	-
Vol. mon. policy shock (x100)	σ_i	0.168	0.021	0.171	0.022	0.154	0.013	0.156	0.015	*0.263*	-	*2.661*	-
Serial-corr. preference shock	ρ_p	0.699	0.062	0.724	0.051	0.625	0.065	0.662	0.064	*0.646*	-	*0.997*	-
Serial-corr. productivity shock	ρ_a	0.842	0.058	0.783	0.053	0.537	0.055	0.604	0.064	*0.762*	-	*0.405*	-
Serial-corr. mon. policy shock	ρ_i	0.434	0.076	0.462	0.077	0.488	0.091	0.458	0.081	*0.282*	-	*0.434*	-
Cross-correlations across countries													
Preference shock - 1/2	$\delta_{p\,12}$	0.320	0.077	0.2899	0.0741	0.228	0.077	0.256	0.072	0.659	0.042	*0.995*	-
Preference shock - 1/3	$\delta_{p\,13}$	0.151	0.080	0.1439	0.0794	0.179	0.080	0.141	0.074	0.572	0.065	*0.995*	-
Preference shock - 2/3	$\delta_{p\,23}$	0.298	0.076	0.3102	0.0801	0.195	0.092	0.237	0.081	0.695	0.057	*0.995*	-
Productivity shock - 1/2	$\delta_{a\,12}$	0.169	0.075	0.173	0.0768	0.159	0.054	0.167	0.081	0.529	0.066	*0.995*	-
Productivity shock - 1/3	$\delta_{a\,13}$	0.067	0.077	0.0729	0.0772	0.017	0.074	0.022	0.083	0.554	0.042	*0.995*	-
Productivity shock - 2/3	$\delta_{a\,23}$	0.214	0.080	0.2174	0.0806	0.197	0.084	0.243	0.080	0.560	0.064	*0.995*	-
Monetary policy shock - 1/2	$\delta_{i\,12}$	0.140	0.083	0.141	0.0798	0.205	0.074	0.154	0.079	0.641	0.042	*0.995*	-
Monetary policy shock - 1/3	$\delta_{i\,13}$	0.131	0.076	0.1362	0.0805	0.147	0.071	0.170	0.071	0.646	0.069	*0.995*	-
Monetary policy shock - 2/3	$\delta_{i\,23}$	0.359	0.078	0.3749	0.0791	0.399	0.072	0.425	0.078	0.694	0.037	*0.995*	-

Figure 1: Data

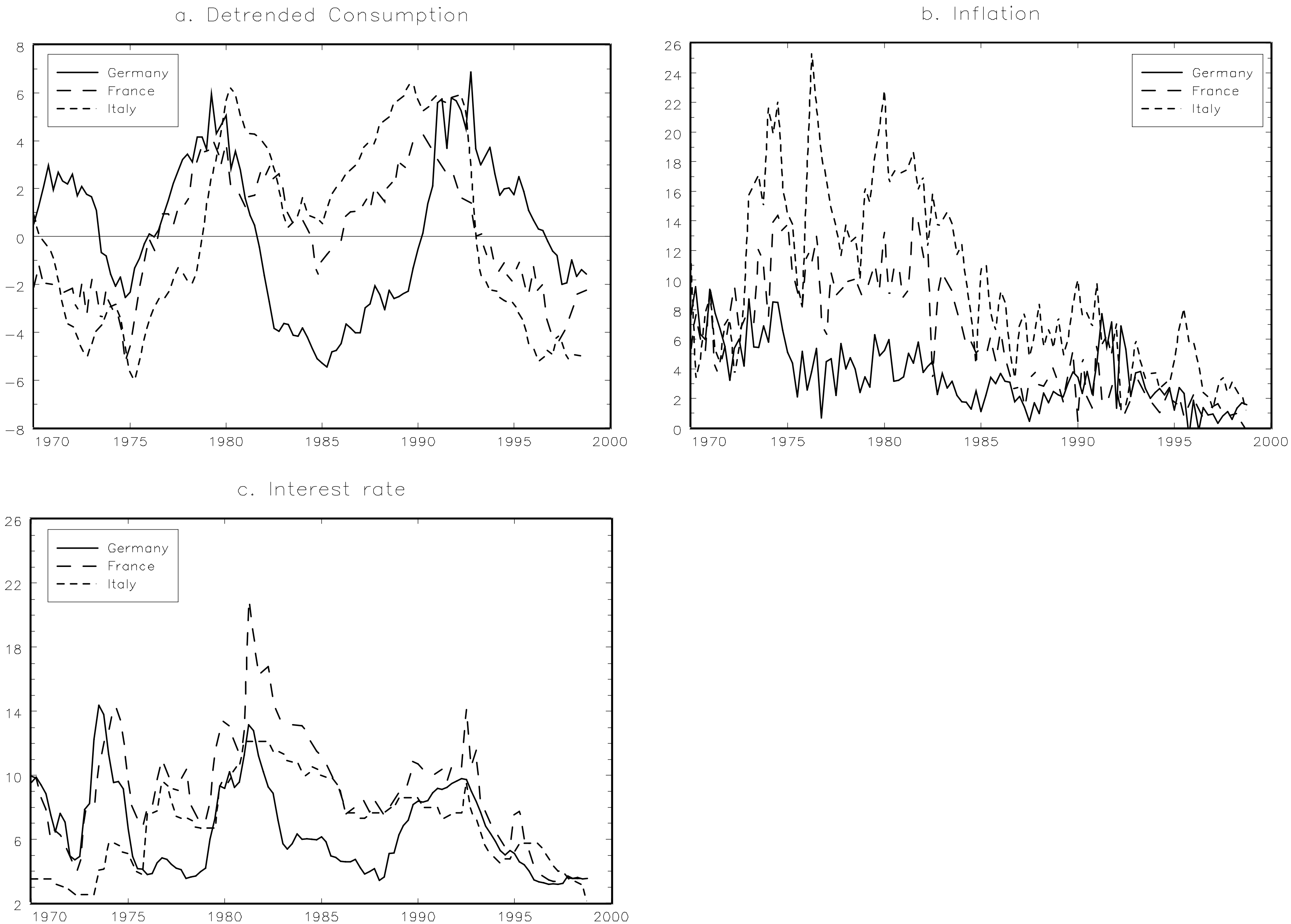
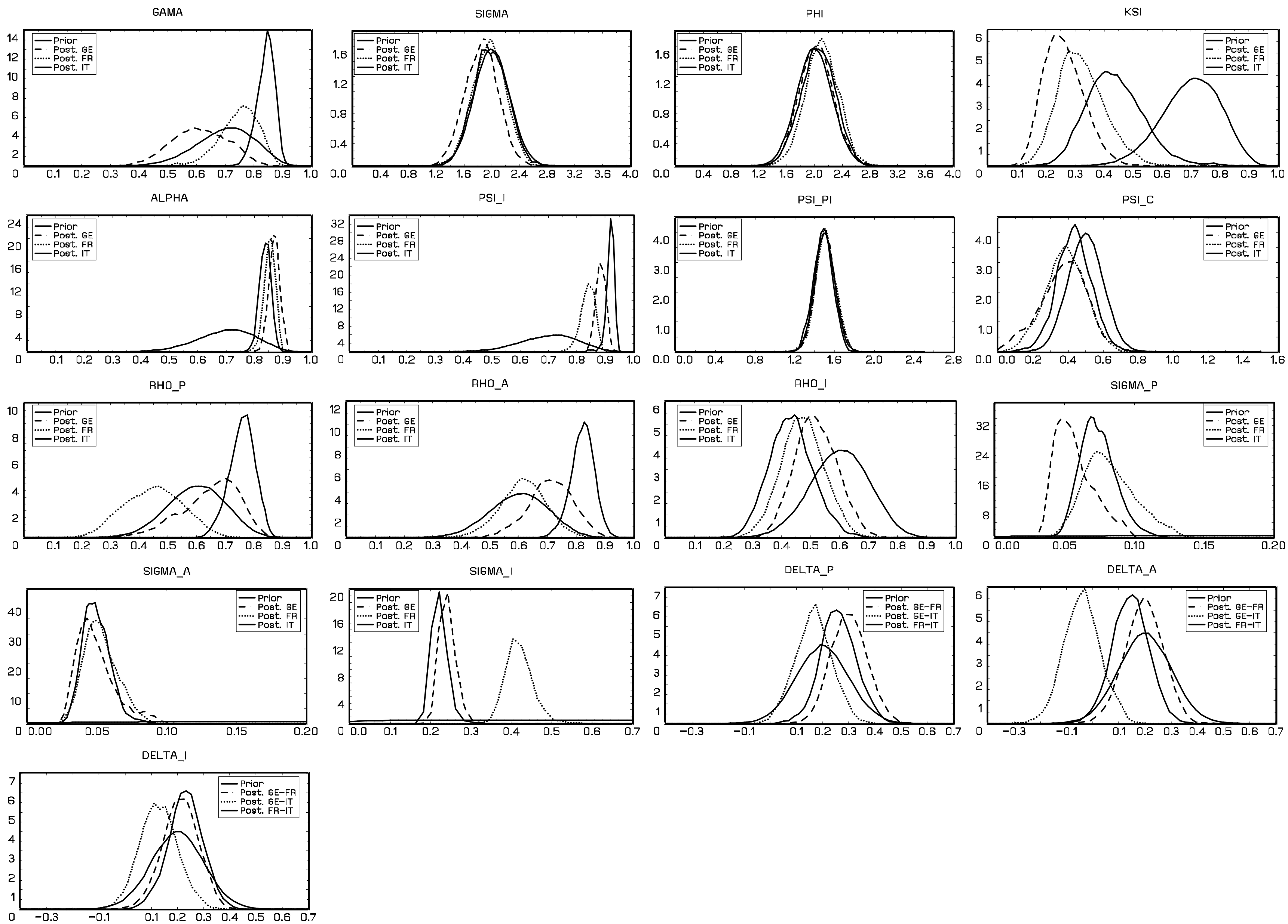


Figure 2: Estimated parameter distribution for the complete MCM



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