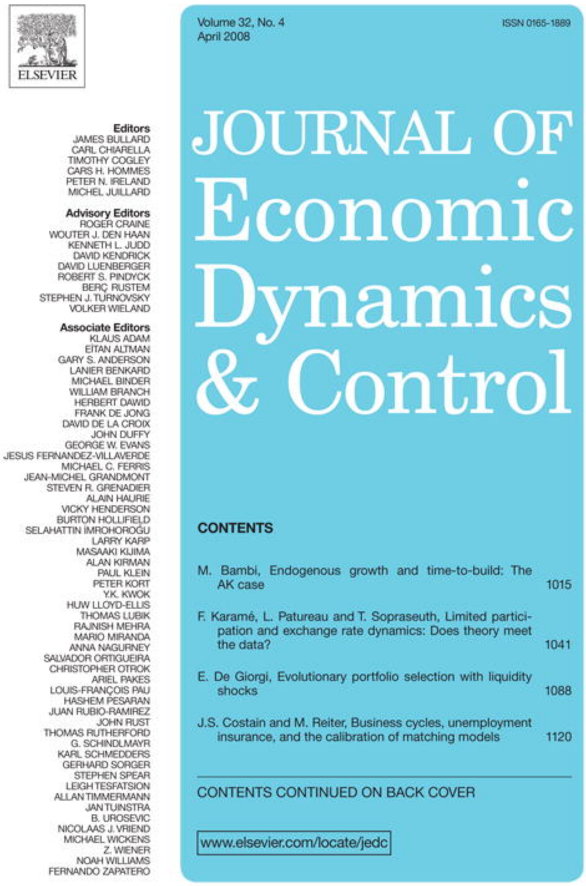
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Limited participation and exchange rate dynamics: Does theory meet the data?

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Abstract

The paper explores the empirical dimensions of a New Open Economy Macronomy model characterized by credit market frictions. We find that these frictions are essential for the model to match a large set of moments of German data. Moreover, the simulated impulse response functions to supply and nominal shocks are consistent with VAR findings. Since the model is estimated on moments rather than on conditional IRFs, this underlines the ability of the model to match the data. Finally, monetary shocks do not seem to be the primary driving force behind the aggregate dynamics, which is consistent with the VAR literature.

JEL classification: F32; F41

Keywords: New Open Economy Macroeconomics; Limited participation model; Simulated method of moments; Structural estimation

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1. Introduction

The field of international macroeconomics has undergone substantial renewal over the last decade with the development of the modelling approach known as 'New Open Economy Macroeconomics' (NOEM). Initiated by Obstfeld and Rogoff (1995), this theoretical strand of literature stresses the introduction of market imperfections into dynamic general equilibrium models based on optimizing agents, in an open-economy setting. The proliferation of models within the NOEM framework illustrates the vivacity of this literature (see Lane, 2001 for a survey). Nevertheless, in spite of the sharp growth in theoretical research in this area, little is known about the empirical performance of such models. While authors typically compare simulated theoretical moments with their empirical counterparts, this procedure does not provide a statistical criterion for assessing the overall goodness of fit of the model. Recent attempts have therefore been made to fill the gap between theory and data. In a closed-economy setting, Ireland (2001), Kim (2000), Collard et al. (2002), among others, use estimation procedures to assess the empirical accuracy of dynamic stochastic general equilibrium models (DSGE). Following on from this line of work, we here use estimation techniques to gauge the goodness of fit of a New Open Economy model.

Ghironi (2000), Smets and Wouters (2003), Bergin (2003) and Lubik and Schorfheide (2007) have also assessed the empirical relevance of NOEM models. While they mainly focus on estimating, and eventually discriminating between different types of nominal rigidities, we choose to estimate an open-economy model with flexible prices and credit market frictions. Pioneered in a closed-economy setting by Fuerst (1992), Lucas (1990) and Christiano (1991), the limited participation assumption models credit market frictions in order to reproduce the persistent fall in the nominal interest rate following a monetary expansion, i.e. the liquidity effect. Christiano et al. (1997) argue that limited participation models may be a more useful starting point than sticky price models to capture the key features of a closed economy's response to monetary shocks. Further, Hairault et al. (2004) show that in a small open-economy model based on credit market frictions, a positive monetary shock generates nominal exchange rate overshooting, which is shown to substantially contribute to nominal exchange rate volatility. In a two-country framework, Schlagenhauf and Wrase (1995) stress the role played by the liquidity effect in the international transmission of economic fluctuations. In a nutshell, the limited participation model has been shown to provide fruitful explanations of the way in which monetary shocks are transmitted to the economy, in particular regarding exchange rate dynamics. The aim of this paper is to go further in the evaluation of its empirical relevance via structural estimation.

We therefore estimate a small open-economy stochastic dynamic model with credit market frictions and an endogenous interest rate rule *à la* Taylor (1993). The model is estimated using the simulated method of moments approach (hereafter SMM) so as to match a set of second-order moments considered as a 'summary' of fluctuations in the German economy during the flexible exchange rate period (1971–1998). Our estimation is based on the matching of 57 moments, which is a large number compared to other papers using SMM. Jonsson and Klein (1996) base their estimation on 31 moments on Swedish data, Collard et al. (2002) use a set of 24 moments, compared to 5 moments in Hairault et al. (1997). With 57 moments, we test the ability of the model to capture a large array of international business cycle features.

Our results confirm the empirical relevance of the limited participation model. First, the estimation procedure suggests that this model is statistically supported by the data. In addition, credit market frictions are found to be essential for making the model consistent with the data. Further, following Christiano et al. (1997), we evaluate the performance of our model with respect to the broad empirical effects of technological and monetary shocks. Our procedure allows us to estimate the behavior of the German Bundesbank and shows that monetary disturbances are not the primary force behind aggregate dynamics. VAR evidence (such as Leeper et al., 1996) has previously underlined that it is the systematic portion of monetary policy that really matters, i.e. the endogenous behavior of the interest rate, rather than exogenous innovations. This has led to the role of exogenous innovations being played down in favor of endogenous behavior. Our results are consistent with this view.

The paper is organized as follows. Section 2 presents the limited participation (LP hereafter). We then describe the estimation procedure in Section 3. Section 4 evaluates the empirical performance of the New Open Economy model. Section 5 illustrates the role of credit market frictions by comparing the predictions of the LP model with those of a simple cash-in-advance (CIA) model. Section 6 concludes.

2. The LP model

We develop a small open-economy NOEM model that is to be estimated on German data. In line with Hairault et al. (2004), the model is characterized by flexible prices and credit market imperfections. Given the time-consuming estimation procedure, we discard the two-country setting in order to lower the number of estimated structural parameters.[1] We then examine the empirical relevance of the LP timing and adjustment costs in accounting for key features of the German economy.

2.1. Timing of the period

The model consists of four types of economic agents (the consumer-household, the perfectly competitive good-producing firm, the financial intermediary and the

[1]Our adoption of a small open-economy, rather than a two-country setting, is motivated by the following reasons. The modelling of a two-country world would require to impose symmetry across countries so as to keep the estimation procedure tractable. This assumption reduces the interest of a two-country framework *per se*, by artificially suppressing cross-country structural differences. Moreover, it is likely to result in misspecification and potentially biased estimates, as estimates could correspond to the mean of the two countries, rather than the two countries separately.

Central Bank) and five markets (goods, labor, loanable funds, foreign assets and money markets). Pioneered in a closed-economy setting by Fuerst (1992), Lucas (1990), and Christiano (1991), the LP assumption models credit market frictions so as to reproduce the persistent fall in the interest rate following a monetary expansion. In our setting, and as is usual in the related literature (Christiano and Eichenbaum, 1992a; Hendry and Zhang, 2001, among others), we model LP by assuming informational asymmetries between lenders and borrowers: households (i.e. lenders) have to make deposits before the monetary policy shock occurs, while firms (i.e. borrowers) have complete information on this shock. We further strengthen the LP transmission mechanism by adding portfolio adjustment costs.

Given the LP assumption, the timing of decisions within a period can be decomposed into three steps.

1. At the beginning of the period, shocks in the model occur. Notably regarding the monetary shock, as the Central Bank acts according to an interest rate rule, monetary policy is conducted by interest rate innovations. The monetary growth factor is endogenously determined by the demand for money and loans.
2. The credit market opens and firms determine their demand for loans. As in Dow (1995), we assume that they have to borrow funds to invest in physical capital. At this stage, the private loan supply is known given the household's deposit choice made in the previous period, according to the LP assumption. The firms also determine their demand for labor and capital. Furthermore, transactions on the various markets (goods, labor, financial assets) take place.
3. At the end of the period, the household chooses the amount M^b_{t+1} to put in the bank the next period and the amount of money-cash M^c_{t+1} for consumption purchases in the next period.

As underlined by Hendry et al. (2003), this timing can be summarized by Fig. 1.

In order to highlight the role of credit market frictions, we will estimate in Section 5 a model without LP. We refer to this model as the CIA model (same model as LP without timing restrictions with respect to the household's portfolio choices and no adjustment costs on money holding). In that framework, the timing of decisions within a period consists of two steps. At the beginning of the period, all shocks occur. Then, decisions and market operations occur. In particular, and in contrast with the LP model, the household optimally decides the amount of money-cash M^c_t and the amount of money-deposits M^d_t in period t for period t, *after* the monetary shock of the current period has occurred.

2.2. Household behavior

The representative household maximizes her expected intertemporal utility:

$$\bar{U}_0 = \mathrm{E}_0 \sum_{t=0}^{\infty} \beta^t U(C_t, L_t), \tag{1}$$

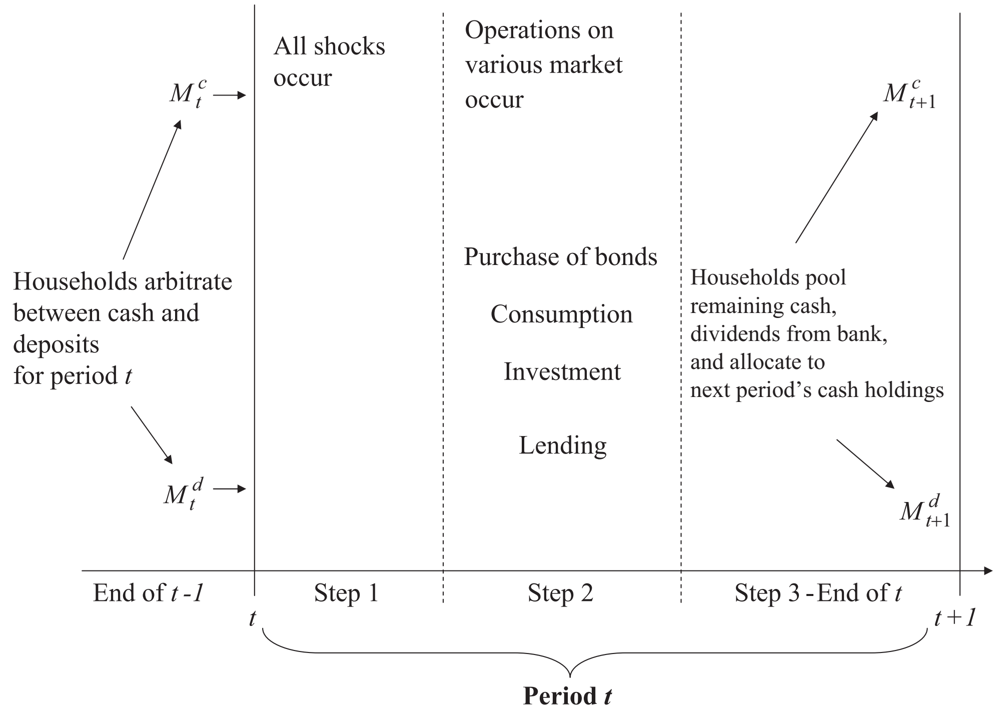

Fig. 1. Timing of the period.

with C_t the consumption bundle and L_t leisure. We assume logarithmic preferences on consumption and leisure

$$U(C_t, L_t) = \ln C_t + \gamma_L \ln L_t, \quad \gamma_L > 0.$$

2.2.1. Allocation of resources

Each economy (home and foreign) is specialized in the production of a single good (H, F). Each variety is produced by an infinite number of firms which perfectly compete with each other. The domestic representative household consumes both types of goods (C_{Ht} and C_{Ft}), the consumption bundle C_t being defined according to the following Cobb–Douglas function[2]:

$$C_t = \left[\frac{C_{Ht}}{\omega_t}\right]^{\omega_t} \left[\frac{C_{Ft}}{1-\omega_t}\right]^{1-\omega_t}.$$

Based on Stockman and Tesar (1995), ω_t is the stochastic relative weight of domestic goods whose law of motion is given by

$$\ln \omega_{t+1} = (1-\rho_\omega)\ln\omega + \rho_\omega \ln\omega_t + \varepsilon_{\omega,t+1}, \tag{2}$$

[2]The Cobb–Douglas specification is a particular case of a general CES function, with the elasticity of substitution across domestic and foreign varieties (θ) equal to 1. This unit value is commonly used in the literature (Corsetti and Pesenti, 2001; Kollmann, 2001b or Anderson and Beier, 2005) and is consistent with the estimated value obtained by Bergin (2006). Our preliminary attempts to estimate θ (that is, assuming a CES function) also yielded a value not statistically different from 1, which led us to calibrate $\theta = 1$.

with ω the mean of the process, $0<\rho_\omega<1$ the persistence parameter and ε_ω an *iid* white noise. The chosen specification for preferences implies that purchasing power parity does not hold (as long as ω is not constant and is different from 0.5). Since an extensive empirical literature[3] concludes that there are sharp deviations from purchasing power parity in the short run, this assumption allows the model to better fit the data in the perspective of structural estimation. C_{Ht} denotes the consumption of the home-produced good and C_{Ft} the consumption of imported goods. The first-order conditions yield the demands for domestic and foreign goods, respectively:

$$C_{Ht} = \omega_t \left[\frac{P_t}{P_{Ht}}\right] C_t, \tag{3}$$

$$C_{Ft} = (1-\omega_t) \left[\frac{P_t}{P_{Ft}}\right] C_t, \tag{4}$$

with P_{Ht} being the price of domestic goods, P_{Ft} the price of imported goods expressed in domestic currency, and P_t the home consumption price index. This leads to the following expression for the domestic consumption price index:

$$P_t = \frac{P_{Ht}^{\omega_t} P_{Ft}^{1-\omega_t}}{\omega_t^{\omega_t}(1-\omega_t)^{1-\omega_t}}. \tag{5}$$

We assume that the law of one price holds, therefore $P_{Ft} = e_t P_t^*$ with P_t^* the foreign consumption price index.[4] Further, as in Kollmann (2001a), we assume that foreign inflation $\pi_t^* \equiv P_t^*/P_{t-1}^*$ is a stochastic process that evolves as

$$\ln \pi_{t+1}^* = (1-\rho_{\pi^*})\ln \pi^* + \rho_{\pi^*} \ln \pi_t^* + \varepsilon_{\pi^*,t+1}, \tag{6}$$

with π^* the mean of the process, $0<\rho_{\pi^*}<1$ the persistence parameter, and ε_{π^*} an *iid* white noise.

2.2.2. Intertemporal program

We introduce money by assuming that the household faces a CIA constraint on her consumption purchases

$$v_t P_t C_t \leqslant M_t^c, \tag{7}$$

where v_t is a velocity shock, supposed to follow an autoregressive process according to

$$\ln v_{t+1} = (1-\rho_v)\ln v + \rho_v \ln v_t + \varepsilon_{v,t+1}. \tag{8}$$

Here v is the mean of the process, $0<\rho_v<1$ the persistence parameter, and ε_v an *iid* white noise. The velocity shock is interpreted as a shock to money demand, as introduced for instance by Ireland and Papadopoulou (2004).

We model the informational asymmetries on the loan market that lie at the heart of the LP assumption as in Andolfatto and Gomme (2003). The amounts of

[3] See Rogoff (1996) for a survey.

[4] We also suppose that the price of goods exported by the foreign country (P_{Ft}^*) is equal to the foreign consumption price index P_t^*, which is consistent with our assumption of a small open-economy.

money-cash M_t^c and money-deposits M_t^b are pre-determined, being chosen in period $t-1$. Furthermore, as noted by Christiano (1991), if such asymmetries allow the model to reproduce the liquidity effect of a money shock, this is far from being as strong and persistent as in the data. To improve the liquidity effect, Christiano and Eichenbaum (1992a) modify the environment so that the financial sector remains more 'liquid' than the real sector for several periods after the shocks. Christiano and Eichenbaum (1992a), followed by King and Watson (1996), Hendry and Zhang (2001) (among others) have thus introduced portfolio adjustment costs. We follow this literature by considering that adjusting money holdings is time-costly. As a result, at period t, when the household chooses her amount of money-cash M_{t+1}^c and its complement (the money-deposit M_{t+1}^b), she takes into account the fact that changing her money holdings M_{t+1}^c is costly: it takes time to reorganize the flow of funds. We assume that the time spent Ω_t on reorganizing the flow of funds is given by

$$\Omega_t = \frac{\xi}{2}\left(\frac{M_{t+1}^c}{M_t^c} - g\right)^2,$$

with $\xi > 0$, the scale parameter of the adjustment costs. In the long-run steady state, M_{t+1}^c/M_t^c is equal to steady-state monetary growth g (equal to the long term inflation rate π). Then, the level of Ω_t as well as its derivative with respect to M_{t+1}^c/M_t^c will be zero in the steady state. Changing M_t^c is costly (in terms of time) with the marginal cost being an increasing function of the parameter ξ. As the time endowment is normalized to unity, leisure L_t is given by the following equation:

$$1 = H_t + L_t + \Omega_t.$$

The household maximizes her expected flow of utility (Eq. (1)) subject to the CIA constraint (Eq. (7)) and the budget constraint

$$\begin{aligned} &M_{t+1}^c + M_{t+1}^b + e_t B_{t+1} + P_t C_t \\ &\quad \leqslant M_t^c + P_t w_t (1 - L_t - \Omega_t) + R_t M_t^b + e_t i_t^F B_t + D_t^f + D_t^b. \end{aligned} \tag{9}$$

The household receives at the end of the period D_t^f and D_t^b, which are the profits of the firm and the bank, respectively. The household's resources also consist of labor income (with w_t the real wage), the amount of money-cash and the amount of money-deposits (both chosen the previous period), the latter yielding some positive return, with R_t the nominal interest rate on bank deposits. The household also saves by holding foreign assets. International financial markets are incomplete, and each period the household buys B_{t+1} foreign asset holdings issued by the rest of the world and denominated in foreign currency. The stock of international assets held in period t (expressed in domestic currency) yields some positive return i_t^F. e_t is the nominal exchange rate, i.e. the home per foreign currency exchange rate. Resources are spent on consumption purchases, financial assets and money demand (both money-cash and money-deposits).

2.3. Firms

The production technology is Cobb–Douglas:

$$Y_t = A_t K_t^{\alpha} H_t^{1-\alpha}, \tag{10}$$

where $\alpha \in [0, 1]$. The technological shock A_t follows a first-order autoregressive process according to

$$\ln A_{t+1} = (1 - \rho_a) \ln A + \rho_a \ln A_t + \varepsilon_{a,t+1}, \tag{11}$$

with A the mean of the process, $0 < \rho_a < 1$ the persistence parameter, and ε_a an *iid* white noise.

The objective of the representative firm is to maximize the stream of dividend payments, where dividends are discounted by their value to the owner of the firm (the consumer). The discount rate is the ratio of the multipliers associated with the household's budget constraint ($\beta(\lambda_{t+1}/\lambda_t)$), which reflects the consumer's change in wealth over time. As in Dow (1995), we assume that the firm has to borrow funds from the bank to invest in physical capital. The cost of borrowing is the nominal interest rate R_t, which equals the return on the household's bank deposits. The program of the firm can be written according to the following Bellman equation:

$$V(K_t) = \underset{\{H_t, K_{t+1}\}}{\mathrm{Max}} \left\{ D_t^f + \mathrm{E}_t \left[\beta \frac{\lambda_{t+1}}{\lambda_t} \right] V(K_{t+1}) \right\}, \tag{12}$$

with the instantaneous profit function given by

$$D_t^f = P_{Ht} Y_t - P_t w_t H_t - P_t R_t I_t - P_t \frac{\Psi_I}{2} \frac{(K_{t+1} - K_t)^2}{K_t} \tag{13}$$

and the law of motion of the capital stock

$$I_t = K_{t+1} - (1 - \delta) K_t, \tag{14}$$

where $\delta \in]0, 1[$ denotes the depreciation rate of capital stock. We suppose that the investment bundle has the same structure as consumption:

$$I_t = \left[\frac{I_{Ht}}{\omega_t} \right]^{\omega_t} \left[\frac{I_{Ft}}{1 - \omega_t} \right]^{1-\omega_t}.$$

The representative firm faces capital adjustment costs that we assume to be quadratic, scaled by the parameter Ψ_I. Adjustment costs are assumed to be paid in terms of the consumption bundle:

$$CA_t^K \equiv \frac{\Psi_I}{2} \frac{(K_{t+1} - K_t)^2}{K_t} = \left[\frac{CA_{Ht}^K}{\omega_t} \right]^{\omega_t} \left[\frac{CA_{Ft}^K}{1 - \omega_t} \right]^{1-\omega_t},$$

with $\Psi_I > 0$. We therefore obtain the domestic demand functions for the domestic and foreign goods, respectively,

$$D_{Ht} = \omega_t \left[\frac{P_t}{P_{Ht}} \right] D_t,$$

$$D_{Ft} = (1 - \omega_t) \left[\frac{P_t}{P_{Ft}} \right] D_t,$$

with D_t being aggregate demand: $D_t = C_t + I_t + CA_t^K$. The firms sell their production to domestic agents and to the rest of the world, that is

$$Y_t = D_{Ht} + X_t.$$

We assume that foreign demand for the domestic good X_t is a decreasing function of the relative price of imported (domestic) goods in the large economy $P_{Ht}/e_t P_t^*$ and a foreign demand term, according to the following equation:

$$X_t = \left[\frac{P_{Ht}}{e_t P_t^*} \right]^{-1} \chi, \tag{15}$$

with the foreign demand term χ assumed to be constant.

2.4. The Central Bank and financial intermediaries

Each period, an amount of money Θ_t is injected into the loanable funds market. The money stock evolves according to

$$M_{t+1} = M_t + \Theta_t, \tag{16}$$

with the monetary injection defined as

$$\Theta_t = (g_t - 1) M_t. \tag{17}$$

Substituting Eq. (17) in Eq. (16), we obtain

$$M_{t+1} = g_t M_t. \tag{18}$$

The Central Bank conducts monetary policy by gradually adjusting the short-term nominal interest rate R_t in response to deviations in output, inflation and money growth from their steady-state values Y, π and g, according to the following policy rule:

$$\ln \frac{R_t}{R} = \rho_r \ln \frac{R_{t-1}}{R} + \rho_\pi \ln \frac{\pi_t}{\pi} + \rho_y \ln \frac{Y_t}{Y} + \rho_g \ln \frac{g_t}{g} + \varepsilon_{Rt}. \tag{19}$$

The parameters ρ_y and ρ_π are strictly positive if the Central Bank is willing to stabilize inflation and output, with inflation defined as $\pi_t = P_t/P_{t-1}$. As standard in the literature, the policy shock ε_{Rt} is assumed to be non-persistent and equal to the serially uncorrelated innovation ε_{Rt}, which is normally distributed with mean zero and standard deviation σ_R. As discussed in Clarida et al. (1998), OECD Central Banks have a tendency to smooth interest rates (which implies $0 < \rho_r < 1$).

In addition, we allow the monetary authority to react to current money growth (through the inclusion of the parameter ρ_g) in order to take into account a distinctive feature of the Bundesbank's monetary policy. Since 1975, the German Central Bank has announced targets for monetary growth and – according to its own descriptions – based monetary policy decisions on deviations of actual money growth from these targets.[5] As in Ireland (2003) and Kim (2000), we retain this specification as it is shown to avoid indeterminacy problems that may arise from the use of an interest rate rule.[6] The *ex ante* specification for the monetary rule is more general than the standard Taylor (1993) rule. The estimation procedure will allow us to go further into the analysis of Bundesbank behavior, since the reaction function parameters will be estimated.

As underlined by Hendry et al. (2003), the Central Bank manipulates g_t to achieve its desired level for the nominal interest rate. More precisely, a desired level for interest rates, along with the state variables of the economy and the optimizing behavior of economic agents, implies a specific value for the demand for real money balances. When monetary authorities set money supply equal to this demand, the desired interest rate is achieved.

The financial intermediary accepts the household's deposits (M_t^b) made at period $t-1$, which are repaid at the end of period t at interest rate R_t. The bank also receives cash injections Θ_t from monetary authorities. The bank's resources are loaned to the firm. The end-of-period profit is redistributed to the household in the form of dividends. The asset balance of the bank implies

$$P_t I_t = M_t^b + \Theta_t, \tag{20}$$

where $P_t I_t$ are loans made to the firm. At the end of the period, the bank dividends are

$$D_t^b = R_t P_t I_t - R_t M_t^b. \tag{21}$$

2.5. Net foreign assets

As in Kollmann (2002), we assume that i_t^F, the interest rate at which the household can borrow or lend foreign assets, equals the exogenous world interest rate i^* plus a spread that is decreasing in the country's real net foreign asset position:

$$i_t^F = i^* - \tau \frac{e_{t-1} B_t}{P_{t-1}}, \quad \tau > 0. \tag{22}$$

[5] The move to monetary targeting came on December 5, 1974 when the Central Bank Council of the Deutsche Bundesbank announced that '*from the present perspective it regards a growth of about 8% in the Central Bank money stock over the whole of 1975 as acceptable in the light of its stability goals*'. The Bundesbank considered this target to '*provide the requisite scope... for the desired growth of the real economy*' (Deutsche Bundesbank, Annual Reports, December 1974).

[6] In the estimation procedure, we ensure that we have determinacy by checking that our estimated model meets Blanchard and Kahn's (1980) conditions.

This assumption eliminates the issue of non-stationarity in the model. As shown by Ghironi (2001), the incomplete asset market assumption makes the model non-stationary since temporary shocks have permanent effects on macroeconomic variables. The literature proposes alternative ways of avoiding non-stationarity. Corsetti et al. (2005) achieve steady-state determinacy by using an endogenous discount factor. Kollmann (2003) introduces adjustment costs on assets for similar reasons. We retain the method adopted by Kollmann (2002).[7]

Since we model a small open-economy, equilibrium quantities are computed at the given foreign price level P^* and the given nominal foreign interest rate i^*. That is, on the foreign asset market, the domestic household can carry out any foreign asset she is willing to hold given the foreign interest rate, only being constrained by her budget constraint. We thus infer from the budget constraint (Eq. (9)) and market equilibria that the household's foreign asset holdings evolve as

$$e_t B_{t+1} - e_t(1 + i_t^F)B_t = P_{Ht} Y_t - P_t D_t. \tag{23}$$

This equation reflects the equilibrium of the balance of payments of the home economy. The small country trades with the rest of the world, depending on the levels of the home production and absorption. If domestic production exceeds absorption ($P_{Ht} Y_t - P_t D_t > 0$), the trade balance is positive while the capital account is negative: the household sells the production surplus abroad and increases her holding of foreign assets. In contrast, if domestic production cannot satisfy the domestic demand for goods, the economy has to import goods from the rest of the world and finance its trade deficit by borrowing from abroad.

2.6. LP and exchange rate dynamics

The first-order conditions on foreign assets B_{t+1} and bank deposits M_{t+1}^b are given by

$$e_t \lambda_t = \beta \mathrm{E}_t(\lambda_{t+1} e_{t+1} i_{t+1}^F),$$
$$\lambda_t = \beta \mathrm{E}_t(\lambda_{t+1} R_{t+1}),$$

where λ_t denotes the Lagrangian multiplier associated with the budget constraint (9). Log-linearizing both conditions around the steady state yields the following equation:

$$\mathrm{E}_t \widehat{R}_{t+1} - \mathrm{E}_t \widehat{i}_{t+1}^F = \mathrm{E}_t \widehat{e}_{t+1} - \widehat{e}_t, \tag{24}$$

where $\widehat{x}_t \equiv (x_t - x)/x$ stands for the percentage deviation from the steady state. An expected decrease in the relative return of domestic deposits ($\mathrm{E}_t \widehat{R}_{t+1} - \mathrm{E}_t \widehat{i}_{t+1}^F < 0$) has to be offset by an expected appreciation of domestic currency to restore the arbitrage condition between domestic and foreign assets. Eq. (24) can be interpreted as the

[7] Schmitt-Grohe and Uribe (2003) investigate the quantitative differences implied by alternative approaches proposed in the literature to induce stationarity. Their main finding is that all versions deliver identical dynamics at business-cycle frequencies.

uncovered interest rate parity (UIRP) condition and accounts for the joint dynamics of the nominal exchange rate and the interest rate.[8]

2.7. Equilibrium

The stochastic environment of our small open-economy consists of the vector of exogenous shocks $\{A, \varepsilon_R, \omega, v, \pi^*\}$.[9] Equilibrium is characterized by the set of prices $\Omega_t^P = \{w_t, R_t, P_t, P_{Ht}, e_t\}_{t=0}^{\infty}$ and the set of quantities[10]

$$\Omega_t^C = \{C_t, H_t, B_{t+1}, M_{t+1}^b, M_{t+1}^c, M_{t+1}\}_{t=0}^{\infty},$$
$$\Omega_t^Q = \{Y_t, H_t, K_{t+1}\}_{t=0}^{\infty},$$

such that

- given the set of prices Ω^P, the vector of exogenous foreign variables $\{i^*, \chi\}$ and the vector of exogenous shocks $\{A, \varepsilon_R, \omega, v, \pi^*\}$, the set of quantities Ω^C maximizes the expected intertemporal utility of the household subject to Eqs. (7) and (9),
- given the set of prices Ω^P, the vector of exogenous foreign variables $\{i^*, \chi\}$ and the vector of exogenous shocks $\{A, \varepsilon_R, \omega, v, \pi^*\}$, the set of quantities Ω^Q maximizes the profits of the representative firm subject to Eqs. (10) and (14),
- given the sets of quantities Ω^C and Ω^Q, the vector of exogenous foreign variables $\{i^*, \chi\}$ and the vector of exogenous shocks $\{A, \varepsilon_R, \omega, v, \pi^*\}$, the set of prices Ω^P ensures that the labor market, the loanable funds market, the goods market, the money market and the current account clear.

Hairault et al. (2004) show that this small open-economy model captures nominal exchange rate volatility. The intuition behind this result is similar to that in Dornbusch (1976). Following a monetary expansion, the domestic interest rate persistently declines due to frictions in the credit market. As the foreign interest rate i^* is constant, the interest rate spread turns out to be negative. Given UIRP, this negative interest rate differential results in an anticipated appreciation of the currency, which generates an immediate depreciation of the nominal exchange rate beyond its new steady-state value. The nominal exchange rate displays overshooting

[8]However, this equation differs from the traditional UIRP equation in two respects. First, the interest rate in Eq. (24) is the future interest rate R_{t+1} (rather than the current one R_t): the LP assumption implies that the relevant interest rate is the *expected* return on deposits, i.e. $E_t R_{t+1}$. Second, the return on foreign assets in Eq. (24) is i^F, which differs from the world interest rate i^* given the risk premium term (see Eq. (22)). The model consequently allows endogenous deviations from UIRP. These deviations are completely endogenous and quantitatively negligible (given our calibration for τ, see Table 2). We will then consider in our subsequent comments that UIRP holds.

[9]The large fiscal expenditures associated with German reunification in 1991 may suggest that fiscal shocks should be included in the estimation procedure. We made the opposite choice by eliminating (as far as possible) the impact of German reunification in the aggregate series used for estimation. We accordingly consider GDP without public expenditures in our data set. Further, we exclude outliers in the macroeconomic series attributable to this particular episode. See Table 11, Appendix A for details.

[10]For the sake of simplicity, we do not distinguish the notations of supply variables and demand variables on all markets.

dynamics as in Dornbusch (1976). Hairault et al. (2004) show that the nominal exchange rate obtained from the model simulated with exogenous monetary shocks displays volatility close to that observed in the data. They conclude that nominal exchange rate overshooting improves our understanding of exchange rate dynamics. However, the authors do not provide a statistical criterion for assessing the overall goodness of fit of their model. The next sections in this paper aim to fill this gap.

3. The structural estimation procedure

Apart from the calibration methodology implemented to evaluate the quantitative performances of RBC models in the spirit of Lucas (1987), two estimation approaches have recently been developed to provide DSGE models with an econometric framework, namely the Bayesian maximum likelihood technique (Kim, 2000; Ireland, 2001 among many others) and the simulated method of moments (developed by Ingram and Lee, 1991; Duffie and Singleton, 1993 and implemented by Collard et al., 2002 among others). Both approaches can be viewed as extending calibration. With both techniques, some parameters have to be calibrated if not potentially estimated for identification purposes or due to a lack of information in the data. In this paper, we adopt the moment-based estimation procedure.

The first reason behind our choice is linked to the issue of stochastic singularity in linearized structural models. The Bayesian maximum likelihood approach requires an identical number of exogenous shocks and endogenous variables in the structural model. This condition is not fulfilled in most cases. Lubik and Schorfheide (2005) and Ruge-Murcia (forthcoming) highlight that adding measurement error terms into the observation equation of the state-space representation, even though widespread in the literature, may make it more difficult to accurately estimate the parameters of structural shocks. There is no such requirement with the SMM, the only (and weaker) constraint being that the number of instrumental parameters (the number of moments to match) is greater than or equal to the number of structural parameters to estimate. A second related argument in favor of a moment-based method for estimating DSGE models, is linked to the information set used in the estimation procedure. As it may be easier to find a large set of linearly independent moments calculated on endogenous variables than linearly independent variables to use in the Bayesian maximum likelihood approach, the SMM may be implemented in a larger and richer informational context. Ruge-Murcia (forthcoming) underlines that the identification of structural parameters may thus be sharper with a moment-based approach. We present more detail of the SMM methodology in the next section.

3.1. *The simulated method of moments*

Consider the following dynamic structural model:

$$\begin{cases} \mu_0(z_t, z_{t-1}, u_t, \varphi_1) = 0, \\ r_0(u_t, u_{t-1}, \varepsilon_t, \varphi_2) = 0, \end{cases}$$

with $\{z_t\}_{t=1}^{t=T}$ the set of observed endogenous variables, $\{u_t\}_{t=1}^{t=T}$ and $\{\varepsilon_t\}_{t=1}^{t=T}$ the sets of unobservable variables and $\varphi = \{\varphi_1, \varphi_2\}$ the set of q unknown structural parameters. $\{\varepsilon_t\}_{t=1}^{t=T}$ is an *iid* process whose distribution is supposed to be known. μ_0 and r_0 are functions capturing the relationship between variables and shocks. In our setting, z_t includes the set of following variables $\{C_t, I_t, Y_t, TB_t, M_t, P_t, e_t, R_t\}$ (where TB_t stands for trade balance), u_t is the set of stochastic variables $\{A_t, \omega_t, v_t, \pi_t^*, \varepsilon_{Rt}\}$ associated with innovations $\varepsilon_t = \{\varepsilon_{at}, \varepsilon_{\omega t}, \varepsilon_{vt}, \varepsilon_{\pi^* t}, \varepsilon_{Rt}\}$.

The SMM dwells on a simple estimation criterion (a quadratic function) that depends on a set of p moments ($p \geqslant q$) calculated on both the observed data and their theoretical counterparts: the idea is to find the value for the q structural parameters that minimizes the distance between the set of p moments calculated on observed data and the same set of p moments estimated on the simulated theoretical series.

The method is implemented as follows. Suppose that the structural model is the 'true' data generating process. Under this assumption, the data simulated from the model ($\{z_t^s(\varphi_0)\}_{t=1}^{t=T}$) for the true set of deep parameters φ_0 and one draw s of structural innovations display the same statistical properties (denoted $sp(.)$) as the observed data (z_t) thus giving

$$\mathrm{E}[sp(z_t^s(\varphi_0))] = \mathrm{E}[sp(z_t)]. \tag{25}$$

However, we only have $SP_T(.)$, the empirical counterpart of this moment condition, calculated on both the observed data and their theoretical counterparts and defined as:

$$SP_T(x) = \frac{1}{T}\sum_{t=1}^{T} sp(x_t),$$

with $x = z$ or $z^s(\varphi)$.

The method can be summarized as follows:

$$\begin{array}{ccc} SP_T(z) & \xrightarrow{\mathrm{p}} & \mathrm{E}[sp(z_t)](= \mathrm{E}[sp(z_t^s(\varphi_0))]), \\ \Uparrow & & \Uparrow \\ SP_T(z^s(\varphi)) & \xrightarrow{\mathrm{p}} & \mathrm{E}[sp(z_t^s(\varphi))]. \end{array}$$

As $SP_T(.)$ converges in probability to $\mathrm{E}[sp(.)]$, and according to Eq. (25), inferring a value for φ_0 means finding the value of deep parameters φ such that the simulated set of moments $SP_T(z^s(\varphi))$ calculated from the structural model is as close as possible (along a particular metric) to $SP_T(z)$ calculated once and for all on the observed data.

The estimation procedure consists of the following steps.

1. From the data, we compute the moments (variance, autocorrelations, cross-correlations) that capture the joint dynamics of all the time series in our sample.
2. The structural model is solved. The solution method relies on a first-order log-linearization around the steady state (Farmer, 1993). Hence for fixed initial conditions z_0 and u_0 and S draws $\{\varepsilon_t^1\}_{t=1}^{t=T}, \ldots, \{\varepsilon_t^S\}_{t=1}^{t=T}$ from the known

distribution,[11] we simulate $S>1$ unique particular paths for the variables of interest $\{z_t^1(\varphi)\}_{t=1}^{t=T}, \ldots, \{z_t^S(\varphi)\}_{t=1}^{t=T}$ through $r_0(.)$ and $\mu_0(.)$ for a particular set of structural parameters φ.

3. We then average the moments calculated on simulated data and noted

$$SP_T^S(z(\varphi)) = \frac{1}{S}\sum_{s=1}^{S} SP_T(z^s(\varphi)).$$

4. The SMM estimator $\widehat{\varphi}_T(W)$ is the value of φ that brings $SP_T^S(z(\varphi))$ (the set of moments computed on simulated data) as close as possible to $SP_T(z)$ (the set of moments computed on observed data). The distance between $SP_T^S(z(\varphi))$ and $SP_T(z)$ is measured according to some metric W. The SMM estimator $\widehat{\varphi}_T(W)$ is then defined as

$$\widehat{\varphi}_T(W) = \arg\min_{\varphi} f(\varphi)' W f(\varphi),$$

with $f(\varphi) = [SP_T(z) - SP_T^S(z(\varphi))]$ and W a positive semi-definite matrix converging to a deterministic matrix. An estimate of the metric W is required for the computation of $\widehat{\varphi}_T(W)$. Ingram and Lee (1991) derived the optimal value W^* for the metric

$$W^* = \left[\left(1 + \frac{1}{S}\right)\Omega\right]^{-1}$$

that depends on the variance–covariance matrix of moments Ω estimated on observed data and based on the generalized method of moments, since this method implies that

$$\sqrt{T}(SP_T(z) - \mathrm{E}[sp(z_t)]) \xrightarrow{L} \mathrm{N}(0, \Omega).$$

Convergent estimation of the optimal metric, noted $\widehat{W}_T^*$, requires a correction for autocorrelation and/or heteroskedasticity for the estimation of Ω. We use the parametric VARHAC correction proposed by Den Haan and Levin (1997).[12] This metric is computed once and for all in step 1.

Steps 2–4 are repeated until a value of the estimated parameters $\widehat{\varphi}_T(\widehat{W}_T^*)$ (noted $\widehat{\varphi}_T^*$) minimizes the objective function.[13]

[11]We suppose that the innovations are *iid* gaussian. These draws remain identical through the estimation procedure.

[12]Alternative corrections are available in the econometric literature. In order to choose the appropriate correction method, we implemented the following counterfactual experiment. First, suppose that the data generating process is the theoretical model and that we know the true value of deep parameters. We then generate fake data with the model. We estimate the structural parameters through SMM. Only the estimation led with the VARHAC corrected metric recognized the complete match between the data (generated by the model by assumption) and the simulated data (also generated by the model through SMM) and led to the global acceptance of the structural model. This experience also allows to check that the deep parameters of the model are well identified.

[13]To minimize the SMM criterion, we use the downhill simplex method (Nelder and Mead, 1965), since this method requires only evaluations and not the derivatives of the function, as do gradient-based methods like BFGS and Newton (among others). The downhill simplex method is used several times in

In order to measure the goodness of fit of our SMM results, the econometric literature proposed three indicators.

- In order to test the significance of the structural parameters, we rely on the following test statistic, based on the asymptotic distribution of $\widehat{\varphi}_T^*$:

$$\sqrt{T}(\widehat{\varphi}_T^* - \varphi_0) \rightarrow \mathrm{N}\left(0, \left[1+\frac{1}{S}\right]\left[\frac{\partial f(\varphi)'}{\partial \varphi} W^* \frac{\partial f(\varphi)}{\partial \varphi'}\right]^{-1}\right)$$

  whose asymptotic variance matrix is easily obtained from numerical derivatives of $f(\varphi)$ evaluated for the solution and from convergent estimation of W^*.
- In order to gauge the overall match between the model and the data, Gourieroux et al. (1993) derived a simple diagnostic test inspired from the over-identification test *à la* Hansen (1982). Hence, the statistics

$$\Psi(\widehat{\varphi}_T^*) = T\left[1+\frac{1}{S}\right]^{-1} [f(\widehat{\varphi}_T^*)]' \widehat{W}_T^* [f(\widehat{\varphi}_T^*)]$$

  tests the global adequacy of the moments computed on simulated data to the moments computed on the observed data set[14] and is asymptotically $\chi^2(p-q)$ distributed.
- This test of global adequacy is completed by a test of goodness of fit of each moment estimated on the simulated data set to its empirical counterpart. This statistic comes from Collard et al. (2002), who adapt Gallant and Tauchen's (1996) statistics developed in the framework of the efficient method of moments, to the simulated method of moments. The test statistic is such that

$$\left\{ diag\left\{ \Omega - \frac{\partial f(\varphi)}{\partial \varphi'}\left[\frac{\partial f(\varphi)'}{\partial \varphi} W^* \frac{\partial f(\varphi)}{\partial \varphi'}\right]^{-1} \frac{\partial f(\varphi)'}{\partial \varphi}\right\}\right\}^{-1/2} \sqrt{T} f(\varphi)$$

  is then asymptotically $\mathrm{N}(0,1)$ distributed.

Finally, we will test restrictions on the value of some deep parameters. A simple test of d restrictions on the structural parameters can then be carried out when computing the estimation of the structural model under the null, $\widehat{\varphi}_T^{0*}$. Gourieroux et al. (1993) show that

$$\Psi(\widehat{\varphi}_T^{0*}) - \Psi(\widehat{\varphi}_T^*)$$

is asymptotically $\chi^2(d)$ distributed.

(*footnote continued*)
order to reach the neighborhood of the solution. Finally, we use a gradient-based method to reach the solution.

[14]This statistic is also an indicator of the robust estimation of the model. Dridi et al. (2007) show that an estimation that passes the over-identification criterion is robust to the mis-specification of the structural model.

The SMM provides a wide range of statistical indicators to measure the goodness of fit of the model, both overall and along each dimension included in the set of moments.

3.2. What dimensions in the data should the model fit?

As Fève (1997) notes, the SMM dwells on the ability of the structural model to match a large set of moments. Our SMM estimates are then such that the model matches the short-run behavior of eight German time series over the post-Bretton Woods (1971:1–1998:4) period. Our choice of Germany is motivated as follows. First, we focus on a small-open developed economy with flexible exchanges rates *vis-à-vis* the US dollar over the period. Secondly, our model is based on credit market frictions; this provides a natural criterion to select the country accordingly. Information from the economic freedom database regarding the degree of credit market constraints in G7 countries led us to choose Germany.[15]

The quarterly time series date are obtained from the OECD BSDB database. The data consist of real GDP (Y), real investment (I) for the business sector, real private consumption (C), the trade balance for goods and services (in volume, deflated by the mean of output calculated over the period, $TB/\bar{Y}$), the monetary aggregate (M_1), the consumer price index (less energy and food) (P), the nominal exchange rate of the Deutschemark *vis-à-vis* the US dollar (E) and the nominal short-run interest rate (R). We focus on the ability of the model to match the statistical moments of the macroeconomic variables that are endogenous in our small open-economy setting, and that are usually found in the related literature (Bergin, 2006, 2003). We also consider investment dynamics since the liquidity effect inherent in the LP model modifies investment choices. Consumption is included in order to gauge the accuracy of the model with regard to other standard business cycle features such as consumption dynamics.

All data, except the nominal interest rate, are seasonally adjusted. We then identify and eliminate outliers in some of the empirical series attributable to oil shocks and German reunification, which are exogenous events that the model will not be able to capture. Empirical moments are then calculated on time series taken in logarithms (except for the trade balance) and detrended using Hodrick and Prescott's (1997) methodology with the adjustment parameter set to 1600. Details on the treatment of the empirical series are provided in Appendix A.

Which moments should the model match? The theoretical model must replicate a set of moments that characterize the key features of the data, which can be captured by standard deviations, autocorrelations and instantaneous as well as dynamic correlations across variables. However, as a model can hardly mimic all the observed moments for all time series, we adopt the pragmatic approach of the partial indirect

[15] The Economic Freedom database (http://www.freetheworld.com/release.html) provides a broad range of indicators regarding various market constraints for a large number of countries and years. Over the period considered, the 5A indicator on credit market frictions suggests that these are most prevalent in Germany as compared to other G7 countries.

inference method developed by Dridi et al. (2007) in the selection of the set of moments to match. The choice of moments is based on two requirements: (i) that they capture a large set of cyclical properties; and (ii) that they can be reproduced by the model. The selection procedure is ultimately validated by the non-rejection of the model according to the global over-identification test *à la* Hansen, which is a sufficient condition to obtain consistent estimates of the structural parameters. This leads us to retain a set of 57 moments, among a large set of empirical features, that our model can fit.[16] We pay particular attention to keep in this set of moments, those 'naturally' imposed by the underlying theory (the moments related to the exchange rate, notably). We thus do not systematically exclude from the set of moments those that are not well statistically matched.

We gauge the ability of the model to match 57 moments, which is a large number compared to other papers using SMM. Jonsson and Klein (1996) base their estimation on 31 moments on Swedish data, and Collard et al. (2002) use a set of 24 moments, versus five moments in Hairault et al. (1997). With 57 moments, if the model is statistically accepted, we can be confident that it captures a large array of dimensions of international business cycles.

The set of moments used for estimation includes variances and autocovariances, as well as instantaneous and dynamic cross-covariances between eight series. These are presented in Table 1.[17]

3.3. Which structural parameters to estimate?

It is widely understood that the estimation results are conditional on the choice of calibrated values (Bergin, 2003). Given this, and to avoid arbitrariness as much as possible, we adopt the following reasoning with regard to the choice of estimated parameters. First, our data set does not deliver enough relevant information to obtain reliable estimates for all parameters. Further, some parameters (such as the capital depreciation rate or the share of capital in production) are commonly pinned down by the Real Business Cycle literature. The calibration of these parameters is then based on standard values, which allows us to estimate the sub-set of parameters whose values are less well-known, or that we consider as critical to match the selected moments. The sub-set of estimated parameters thus includes adjustment cost parameters on money holdings and capital (ξ and Ψ_I, respectively), the coefficients

[16] Dwelling on the partial indirect inference method, we select moments according to the following steps. All moments relative to the time series (i.e. all variances, autocovariances and cross-covariances at lag 0 and 1 for the 8 time series) are initially included in the set of instrumental parameters. After running the structural estimation procedure, we eliminate the moments with the largest positive contribution to the SMM criterion, i.e. that the model reproduces poorly. This elimination procedure explains why some contemporaneous covariances are included in the set of 57 moments while the corresponding dynamic ones are not.

[17] When implementing the SMM procedure, we look for parameter estimates such that the variances and covariances generated by the model match their empirical counterparts. We do not rely on standard deviations and correlations during the estimation process. Indeed, if we had used correlation as a moment to match in this estimation procedure, a potential bias could have resulted as a correlation might be well reproduced even though the underlying covariances and variances are not captured by the model.

Table 1
Which moments to match?

Variance	Autocovariance	Covariance(0)	Covariance(1)	
$\sigma^2(\widehat{e})$	$Cov(\widehat{e},\widehat{e}_{-1})$	$Cov(\widehat{e},\widehat{i})$	$Cov(\widehat{e},\widehat{c}_{-1})$	$Cov(\widehat{m}_1,\widehat{c}_{-1})$
$\sigma^2(\widehat{y})$	$Cov(\widehat{y},\widehat{y}_{-1})$	$Cov(\widehat{e},\widehat{c})$	$Cov(\widehat{e},\widehat{p}_{-1})$	$Cov(\widehat{p},\widehat{e}_{-1})$
$\sigma^2(\widehat{i})$	$Cov(\widehat{i},\widehat{i}_{-1})$	$Cov(\widehat{e},\widehat{p})$	$Cov(\widehat{e},\widehat{TB}_{-1})$	$Cov(\widehat{p},\widehat{i}_{-1})$
$\sigma^2(\widehat{c})$	$Cov(\widehat{c},\widehat{c}_{-1})$	$Cov(\widehat{e},\widehat{TB})$	$Cov(\widehat{y},\widehat{e}_{-1})$	$Cov(\widehat{p},\widehat{c}_{-1})$
$\sigma^2(\widehat{p})$	$Cov(\widehat{p},\widehat{p}_{-1})$	$Cov(\widehat{y},\widehat{i})$	$Cov(\widehat{y},\widehat{i}_{-1})$	$Cov(\widehat{p},\widehat{TB}_{-1})$
$\sigma^2(\widehat{TB})$	$Cov(\widehat{TB},\widehat{TB}_{-1})$	$Cov(\widehat{y},\widehat{c})$	$Cov(\widehat{y},\widehat{p}_{-1})$	$Cov(\widehat{TB},\widehat{e}_{-1})$
$\sigma^2(\widehat{r})$	$Cov(\widehat{r},\widehat{r}_{-1})$	$Cov(\widehat{y},\widehat{m}_1)$	$Cov(\widehat{i},\widehat{e}_{-1})$	$Cov(\widehat{TB},\widehat{y}_{-1})$
$\sigma^2(\widehat{m}_1)$	$Cov(\widehat{m}_1,\widehat{m}_{1,-1})$	$Cov(\widehat{y},\widehat{p})$	$Cov(\widehat{i},\widehat{c}_{-1})$	$Cov(\widehat{TB},\widehat{c}_{-1})$
		$Cov(\widehat{i},\widehat{c})$	$Cov(\widehat{i},\widehat{p}_{-1})$	$Cov(\widehat{TB},\widehat{p}_{-1})$
		$Cov(\widehat{i},\widehat{p})$	$Cov(\widehat{c},\widehat{e}_{-1})$	$Cov(\widehat{r},\widehat{e}_{-1})$
		$Cov(\widehat{c},\widehat{p})$	$Cov(\widehat{c},\widehat{i}_{-1})$	$Cov(\widehat{e},\widehat{r}_{-1})$
		$Cov(\widehat{c},\widehat{TB})$	$Cov(\widehat{c},\widehat{p}_{-1})$	$Cov(\widehat{m}_1,\widehat{e}_{-1})$
		$Cov(\widehat{p},\widehat{TB})$	$Cov(\widehat{c},\widehat{TB}_{-1})$	$Cov(\widehat{e},\widehat{m}_{1,-1})$
		$Cov(\widehat{r},\widehat{e})$		
		$Cov(\widehat{m}_1,\widehat{e})$		

Note: Lower cases denote variables expressed in logarithms and $\hat{x}$ denotes the HP-filter cyclical component of variable x.

Table 2
Calibrated parameters values

α	β	δ	ω	A	v	H	b	π	π^*	i^*	τ	γ_X	γ_P
0.42	0.988	0.025	0.75	1	1	0.2	0.05	1.0083	1.0126	1.0194	0.0019	1.0058	1.0079

of the interest rate rule, and the shock processes. The remainder of the structural parameters are calibrated to standard values found in the literature. Table 2 sums up the calibrated parameter values.

The model is estimated on a quarterly basis. The calibration on German data draws on Kollmann (2001a). The share of capital in production α is set to 0.42. Aggregate data suggest a quarterly capital depreciation rate of about 2.5%, so δ is set to 0.025. The subjective discount factor β is set at $\beta = 0.988$ implying a real interest rate equal to 1.2% per quarter. The parameter ω is set so that the steady-state imports to GDP ratio is 25%, consistent with German data. The steady-state values for the technological level and the velocity shock are set to 1. γ_L is determined from the calibrated value for H, meaning that 80% of time is devoted to non-work activities. The parameter b represents the steady-state indebtedness flow relative to GDP, and is set to 5%. This value is consistent with the average trade balance over GDP ratio in Germany over the period 1971:1–1998:4. We follow Kollmann (2002) by setting $\tau = 0.0019$. Finally, the parameters γ_X and γ_P stand for the deterministic trends in output and prices, respectively, and are determined by a simple OLS regression on corresponding German quarterly data. The values for π, π^* and i^* correspond to the means of German inflation, US inflation and the US short-term nominal interest rate (OECD BSDB database, quarterly basis) over the period 1971:1–1998:4.[18]

The other 15 parameters of the model are estimated via the simulated method of moments. Section 4 discusses the empirical performance of the LP model.

4. Empirical performance of the liquidity model

The standard evaluation method of business cycle models calculates their *ex post* performance by comparing the second-order moments of the theoretical series to those of the corresponding business cycle series. The SMM technique provides a statistical support for this methodology. Furthermore, uncertainty concerning the 'true' values of structural parameters is taken into account as the SMM approach involves a statistical framework permitting significance tests.

4.1. Estimation results

Table 3 shows the estimation results of the LP model over the period 1971:1–1998:4. Column (2), labelled 'Estimation 1' , displays the estimated values of the 15 parameters not previously calibrated. The global fit of the model is captured by the Hansen statistic. For the model to be accepted by the data, this statistic has to be less than the critical value of a $\chi^2(p-q)$, with p being the number of moments to match (57) and q the number of estimated structural parameters (15). The p-value (0.186) shows that the liquidity model is globally accepted by the data.

Column (2) shows that three estimated parameters are close to 0: the reaction function to output in the interest rate rule ρ_y; the persistence parameter of the foreign inflation rate shock ρ_{π^*}, and the velocity shock ρ_v. In addition, the preference shock seems to be very persistent: ρ_ω is very close to 1. We thus constrain these parameters to the values found in Estimation 1 (bold characters in column (3)),[19] and we test the statistical significance of these restrictions. Results are reported in column (3), under the label 'Estimation 2'. The restriction test statistic (with a p-value equal to 0.729) reveals that we strongly accept the null hypothesis that $\rho_y = \rho_{\pi^*} = \rho_v = 0$

[18]We do not model the foreign interest rate and exports as stochastic processes in our estimation. Preliminary attempts to estimate i^* and χ as shocks revealed that this did not significantly improve the empirical performance of the model. We accordingly calibrate them to their mean values over the period for the sake of parsimony. US inflation is measured as CPI less food and energy.

[19]The values for ρ_{π^*} and ρ_v are constrained to 0.0001 and the value for ρ_ω to 0.999 for technical reasons. As pointed out by Bergin (2003), some regions of the parameter space do not imply a well-defined equilibrium in the model. Specifically, the autoregressive coefficients of shocks have to be strictly greater than zero and strictly less than unity.

Table 3
Estimation results, Germany 1971:1–1998:4

(1) Parameters		(2) Estimation 1	(3) Estimation 2
Adjustment costs of money holdings	ξ	9.4194	9.4469
			(2.173)
Adjustment costs of capital	Ψ_I	0.6211	2.4815
			(0.979)
Taylor rule			
Weight on output	ρ_y	0.0000	**0**
Weight on lagged R	ρ_r	0.5230	0.440
			(2.160)
Weight on inflation	ρ_π	0.05931	0.0927
			(1.318)
Weight on money growth	ρ_g	0.4177	0.4706
			(3.279)
Taylor rule shock			
Volatility	σ_R	0.0009	0.00154
			(1.838)
Technological shock			
Persistence	ρ_a	0.9795	0.9711
			(39.08)
Volatility	σ_a	0.00841	0.00845
			(11.749)
Foreign inflation shock			
Persistence	ρ_{π^*}	10^{-6}	**0**
Volatility	σ_{π^*}	0.05158	0.05253
			(10.198)
Taste shock			
Persistence	ρ_ω	0.9999	**1**
Volatility	σ_ω	0.01225	0.01318
			(14.63)
Money demand shock			
Persistence	ρ_v	0.0000	**0**
Volatility	σ_v	0.00535	0.0053
			(5.317)
Hansen statistic			
$\Psi(\widehat{\varphi}_T^*)$		49.97	52.004
		$\chi^2(42)$	$\chi^2(46)$
p-Value		0.186	0.2516
Restriction tests			
$H_0: \rho_y = \rho_{\pi^*} = \rho_v = 0, \rho_\omega = 1$	$\Psi(\widehat{\varphi}_T^*)$		2.036
			$\chi^2(4)$
Accept H_0	p-Value		0.729
$H_0: \Psi_I = \rho_\pi = 0$	$\Psi(\widehat{\varphi}_T^*)$		11.578
			$\chi^2(2)$

Table 3 (*continued*)

(1) Parameters		(2) Estimation 1	(3) Estimation 2
Reject H_0	p-Value		0.003
$H_0 : \sigma_R = 0$	$\Psi(\widehat{\varphi}_T^*)$		13.921
			$\chi^2(1)$
Reject H_0	p-Value		1.91×10^{-4}

Note: Results obtained with $S = 50$. Figures in parentheses are t-statistics (Estimation 2).

and $\rho_\omega = 1$. Moreover, the goodness of fit of the model is statistically better with a p-value of 25.16%.

In column (3), we report the significance of each estimated parameter: the t-statistic is displayed in parentheses below the estimated coefficient.[20] All of the parameters are statistically significant, apart from the adjustment cost of capital Ψ_I and the response of the monetary authorities to inflation ρ_π. Nevertheless, a restricted version of the model with these parameters set to 0 does not pass the restriction test, as indicated in Table 3. The standard deviation of the monetary innovation is only significant at the 10% level. We thus estimate a version of the model with $\sigma_R = 0$, which leads the conclusion that this is not supported by the data.

Column (3) of Table 3 presents a parsimonious estimation of the LP model which is globally accepted by the data. Comparing these estimation results to those found in the literature is difficult as parameter values are highly dependent on the theoretical model, the dimensions in the data that the estimation procedure tries to fit (the set of variables and set of moments or VAR impulse response functions) as well as the country and period used to estimate the model. Nonetheless, in the following paragraphs, we provide references to previous work in order to check that our result lie in intervals considered reasonable by the literature.

Regarding the estimated interest rate rule, the data suggest that the response of monetary policy to output deviations from the steady state should be zero. Both Ireland and Papadopoulou (2004) and Christiano and Gust (1999) find similar results on US data. Our results confirm the presence of interest-rate smoothing behavior by the Bundesbank, even if ρ_r is lower than the estimate of 0.91 found in Clarida et al. (1998).

The Bundesbank's reaction to inflation (ρ_π) is estimated to be low. As such, the estimated interest rate rule does not satisfy the Taylor principle stating that the Central Bank should raise its interest-rate instrument more than one-for-one with increases in inflation to ensure determinacy. However, this does not imply

[20]The t-statistics are not shown in column (2) as they cannot be calculated for technical reasons. In Estimation 1, some parameters are estimated at the bounds of their interval. This implies that the procedure cannot calculate the estimated standard errors. We then discard the parameters ρ_v, ρ_ω, ρ_{π^*} and ρ_y by setting them to their estimated values, which leads us to Estimation 2.

indeterminacy in our setting. We ensure determinacy by checking that the estimated model meets the Blanchard and Kahn (1980) conditions, thereby leading to a unique saddle path for the inflation rate. In an infinite setting context, it is thus sufficient to know the initial price level to pin down the price level path, even though the low estimated weight on inflation in the interest rate rule violates the Taylor principle. This result, similar to that in Ireland (2001), can be rationalized as follows. The Taylor principle was derived in closed-economy models with sticky prices and simple feedback rules (Woodford, 2003). Our model differs along three dimensions: we consider an open-economy with credit market frictions and with a Central Bank that also targets monetary growth. Each of these dimensions modifies the conventional determinacy condition. De Fiore and Liu (2005) stress that the conditions for determinacy crucially depend on the degree of openness to international trade. Christiano et al. (1999) also emphasize that determinacy conditions are modified in a model with credit market frictions (as compared to a sticky price model).[21] Finally, following Ireland (2001) and Kim (2000), determinacy is achieved through the inclusion of the inflation rate *and* the money growth rate in the Central Bank's reaction function. This solution is consistent with the result in Mac Callum (1986) that indeterminacy is avoided if the monetary authorities's objective function involves nominal variables. This conclusion echoes Patinkin's (1965) view that '*a necessary condition for the determinacy of the absolute price level … is that the Central Bank concern itself with some money value and in this sense be willing to suffer from money illusion*'. We have *two* nominal variables to rule out indeterminacy. While the estimated weight on inflation is low, the Central Bank reacts strongly to deviations in monetary growth. Overall our results highlight the conservative behavior of the Bundesbank, as the interest rate reacts to monetary variables (inflation and monetary growth) but not to output fluctuations. They further suggest that the Bundesbank is inflation-adverse while implementing money growth targeting. Finally, the German monetary policy shock displays a lower standard deviation than that found in US data but lies within the bounds in the literature (from 0.017 in the US in Ireland, 2004 to 0.007 in Ireland, 2001).

The adjustment cost of capital Ψ_I is low compared to the values in Ireland and Papadopoulou (2004), Smets and Wouters (2003) (between 5 and 8) and Bergin (2003) (between 250 and 440). We also obtain a high value for adjustment costs of money holdings ξ compared to the value between 0.004 and 0.02 found by Ireland and Papadopoulou (2004). The low Ψ_I and high ξ can be rationalized via the internal propagation mechanisms of the model. The focus of the model is on the credit market where both ξ and Ψ_I play a role. We present in Section 4.2 the economic mechanisms behind this result. The intuition is that both parameters play a significant role in the adjustment on the loan market. The portfolio adjustment cost parameter ξ influences the way households make their deposit choices, that is the

[21]In particular, they show that Taylor's (1999) recommended weight on inflation and output (1.5 and 0.5, respectively) lead to indeterminacy. In a LP model, they show that equilibria in which expectations about inflation are self-fulfilling are eliminated when the Taylor rule responds very little to output. This is consistent with Ireland and Papadopoulou (2004) find similar results on US data.

private supply of loans. The capital adjustment cost parameter Ψ_I is crucial in investment dynamics, that is the demand for loans. It is likely that the estimation procedure has trouble in simultaneously estimating the two parameters; this may explain why we finally obtain intermediate values for both as compared to the literature.

The estimated values for technological process are consistent with those found in the Business Cycle literature. As for the shock to tastes, our estimate is consistent with the values chosen by Hendry et al. (2003) ($\sigma_\omega = 0.01$ and $\rho_\omega = 0.7$) on Canadian data.

The standard deviation of the foreign inflation shock is of magnitude in line with other results in the literature. Hendry et al. (2003) estimate a VAR model for the foreign price process on Canadian data which yields a standard deviation (0.03) close to that we find on German data (0.048). Lubik and Schorfheide (2007) obtain similar values for σ_{π^*} when estimating a small open-economy model on the United Kingdom, Australia, Canada and New-Zealand using Bayesian techniques (from 0.024 to 0.054). Nevertheless, the standard deviation of the foreign inflation shock is larger than that estimated on other shocks. In our view, the estimated shock to π^* is likely to capture elements larger than 'pure' foreign inflation.[22]

The estimates on the velocity shock lie within the intervals found by Ireland (2003) ($\sigma_v = 0.0076$) and Ireland (2004) ($\sigma_v = 0.0012$). In contrast, our results differ from those found in the literature regarding persistence. Previous values hover around 0.9 while we find that the shock is not persistent. This difference might be due to the fact that our model includes mechanisms inducing persistence such as interest smoothing in the Taylor rule, LP and adjustment costs of money holdings.

The relative variances of the exogenous shocks suggest that the monetary shock is not the primary driving force behind the aggregate dynamics, which is consistent with the VAR literature (Faust and Rogers, 2003). In addition, our results suggest that it is the systematic portion of monetary policy that really matters – the endogenous behavior of the interest rate – rather than exogenous innovations. Leeper et al. (1996) also de-emphasize the role of exogenous innovations in favor of endogenous reactions.

To complete our evaluation of the quality of fit of the empirical moments, we present the empirical moments and their simulated counterparts, as well as Gallant and Tauchen's (1996) statistics, in Tables 4–6.[23]

[22]To check this intuition, we run the following experiment. Based on Eq. (15), we rebuilt the π^* series that derives from our modelling, using the observed DM–$ exchange rate for e_t, German exports in volume for X_t and GDP deflator for p_{Ht}. We then obtained a time series for π^*. To investigate what this reconstructed series captures, we regressed π^* in log on a set of explanatory variables (including a constant and a time trend). Lagged US CPI inflation (less energy and food) as well as lagged change in oil prices were significant, suggesting that the foreign inflation shock estimated in the model captures not only the 'true' US foreign inflation but also other elements such as oil price changes.

[23]While the estimation procedure is based on matching variances and covariances (Table 1), we display in Tables 4–6 the 57 moments in terms of standard deviation, correlation and autocorrelation so that these figures are comparable to the values commonly found in the RBC literature. GT statistics in column (3) are calculated from the variances and covariances used in the estimation procedure.

Table 4
Observed moments vs. simulated moments (1)

Moment	Observed (A)	Simulated (B)	GT stat (C)
Standard deviation (in %)			
$\sigma(\widehat{e})$	8.332	7.842 (0.929)	0.426[a]
$\sigma(\widehat{y})$	1.50	1.120 (0.142)	1.434[a]
$\sigma(\widehat{i})$	4.098	3.634 (0.580)	0.781[a]
$\sigma(\widehat{c})$	1.133	1.031 (0.118)	0.826[a]
$\sigma(\widehat{r})$	0.418	0.248 (0.032)	1.608[a]
$\sigma(\widehat{m}_1)$	2.345	1.890 (0.773)	1.782[a]
$\sigma(\widehat{p})$	1.051	1.058 (0.156)	−0.138[a]
$\sigma(\widehat{TB})$	1.042	0.937 (0.109)	1.206[a]
Autocorrelation			
$\rho(\widehat{e})$	0.8452	0.6958 (0.065)	0.935[a]
$\rho(\widehat{y})$	0.8522	0.6925 (0.0596)	2.011[b]
$\rho(\widehat{i})$	0.8194	0.7019 (0.0837)	1.032[a]
$\rho(\widehat{c})$	0.7732	0.5106 (0.100)	1.516[a]
$\rho(\widehat{r})$	0.8881	0.6156 (0.086)	2.104[b]
$\rho(\widehat{m}_1)$	0.7510	0.8425 (0.0337)	0.604[a]
$\rho(\widehat{p})$	0.9504	0.6796 (0.0737)	1.378[a]
$\rho(\widehat{TB})$	0.7746	0.6795 (0.0641)	1.879[a]

Note: a and b, respectively, denote moments that do pass the GT test at the 5% and 1% levels. For instance, for $\sigma(\widehat{e})$ the GT statistic states that the null hypothesis of equality between the simulated moment and its empirical counterpart is accepted, at the 5% level. Figures in parenthesis are standard deviations.

Consider columns (A) and (B) in the three tables, showing the 57 moments on observed data and simulated series, respectively. The comparison of the two columns shows that the quality of estimation is satisfactory, since both the size and sign of a large number of the empirical moments are reproduced. The test of Gallant and Tauchen (GT), which measures the goodness of fit of each moment, confirms the quality of the estimation. Column (C) in Tables 4–6 shows that only 2 out of the 57

Table 5
Observed moments vs. simulated moments (2)

Moment	Observed (A)	Simulated (B)	GT stat (C)
Instantaneous cross-correlations			
$\rho(\widehat{e},\widehat{i})$	−0.186	−0.3511 (0.1596)	0.646[a]
$\rho(\widehat{e},\widehat{c})$	−0.4172	−0.3410 (0.1375)	−0.578[a]
$\rho(\widehat{e},\widehat{p})$	0.555	0.3660 (0.1343)	0.946[a]
$\rho(\widehat{e},\widehat{TB})$	0.3404	0.4439 (0.1591)	−0.198[a]
$\rho(\widehat{e},\widehat{r})$	−0.1099	−0.3493 (0.1263)	0.485[a]
$\rho(\widehat{e},\widehat{m_1})$	−0.1889	−0.0602 (0.1887)	−0.805[a]
$\rho(\widehat{y},\widehat{\imath})$	0.7787	0.6787 (0.0945)	1.379[a]
$\rho(\widehat{y},\widehat{c})$	0.7197	0.5572 (0.1250)	1.200[a]
$\rho(\widehat{y},\widehat{m_1})$	0.0656	0.2140 (0.1925)	−0.479[a]
$\rho(\widehat{y},\widehat{p})$	−0.3918	−0.4207 (0.1253)	−0.271[a]
$\rho(\widehat{i},\widehat{c})$	0.7701	0.6821 (0.0869)	0.997[a]
$\rho(\widehat{p},\widehat{\imath})$	−0.4858	−0.7081 (0.0591)	0.689[a]
$\rho(\widehat{p},\widehat{c})$	−0.6187	−0.6452 (0.0912)	−0.142[a]
$\rho(\widehat{c},\widehat{TB})$	−0.2412	−0.3829 (0.1423)	0.357[a]
$\rho(\widehat{p},\widehat{TB})$	0.2907	0.3203 (0.1476)	−0.029[a]

Note: a and b, respectively, denote moments that do pass the GT test at the 5% and 1% levels. Figures in parenthesis are standard deviations.

observed moments are not significantly well reproduced by the model at the 5% level, that is the persistence of output and that of the nominal interest rate.[24]

Output persistence is not well captured by the model (0.70 in simulated data versus 0.85 in observed data). However, the value predicted by the LP framework lies within the range found in G7 countries (Backus et al., 1995). Moreover, this result is not necessarily surprising (even though disappointing), as International Business

[24]However, these moments are significantly well reproduced at the 1% level, with the critical values of 1.96 and 2.57 at the 5% and 1% levels, respectively.

Table 6
Observed moments vs. simulated moments (3)

Moment	Observed (A)	Simulated (B)	GT stat (C)
Dynamic cross-correlations			
$\rho(\widehat{e},\widehat{c}_{-1})$	−0.422	−0.2223 (0.1429)	−1.079[a]
$\rho(\widehat{e},\widehat{p}_{-1})$	0.5597	0.2956 (0.1397)	1.138[a]
$\rho(\widehat{e},\widehat{TB}_{-1})$	0.2871	0.3054 (0.1730)	0.066[a]
$\rho(\widehat{e},\widehat{m_1}_{-1})$	−0.1312	0.0021 (0.1942)	−0.809[a]
$\rho(\widehat{e},\widehat{r}_{-1})$	−0.1087	−0.2463 (0.1427)	0.220[a]
$\rho(\widehat{y},\widehat{e}_{-1})$	−0.1276	0.0059 (0.1705)	−0.832[a]
$\rho(\widehat{y},\widehat{\imath}_{-1})$	0.7324	0.4826 (0.1309)	1.736[a]
$\rho(\widehat{y},\widehat{p}_{-1})$	−0.4406	−0.3556 (0.1338)	−0.688[a]
$\rho(\widehat{\imath},\widehat{e}_{-1})$	−0.1487	−0.2475 (0.1610)	0.394[a]
$\rho(\widehat{\imath},\widehat{c}_{-1})$	0.6964	0.5225 (0.1131)	0.980[a]
$\rho(\widehat{\imath},\widehat{p}_{-1})$	−0.5148	−0.6126 (0.0844)	0.258[a]
$\rho(\widehat{c},\widehat{e}_{-1})$	−0.3754	−0.2397 (0.1511)	−0.832[a]
$\rho(\widehat{c},\widehat{\imath}_{-1})$	0.7131	0.4978 (0.1192)	1.594[a]
$\rho(\widehat{c},\widehat{p}_{-1})$	−0.6238	−0.5020 (0.1097)	−0.820[a]
$\rho(\widehat{c},\widehat{TB}_{-1})$	−0.1952	−0.2416 (0.1524)	0.054[a]
$\rho(\widehat{m_1},\widehat{c}_{-1})$	0.3222	0.3017 (0.2547)	0.799[a]
$\rho(\widehat{p},\widehat{e}_{-1})$	0.5016	0.1980 (0.1492)	1.412[a]
$\rho(\widehat{p},\widehat{\imath}_{-1})$	−0.4041	−0.2944 (0.0902)	−0.527[a]
$\rho(\widehat{p},\widehat{c}_{-1})$	−0.5456	−0.3047 (0.1062)	−1.153[a]
$\rho(\widehat{p},\widehat{TB}_{-1})$	0.2586	0.2045 (0.1688)	0.257[a]
$\rho(\widehat{TB},\widehat{e}_{-1})$	0.3787	0.3193 (0.6174)	0.410[a]
$\rho(\widehat{TB},\widehat{y}_{-1})$	0.1999	0.3592 (0.1630)	−0.310[a]
$\rho(\widehat{TB},\widehat{c}_{-1})$	−0.2182	−0.1917 (0.1479)	−0.319[a]

Table 6 (*continued*)

Moment	Observed (A)	Simulated (B)	GT stat (C)
$\rho(\widehat{TB},\widehat{p}_{-1})$	0.2659	0.2492 (0.1497)	0.123[a]
$\rho(\widehat{r},\widehat{e}_{-1})$	−0.1186	−0.2253 (0.1348)	0.082[a]
$\rho(\widehat{m_1},\widehat{e}_{-1})$	−0.2287	−0.1436 (0.1866)	−0.719[a]

Note: a and b, respectively, denote moments that do pass the GT test at the 5% and 1% levels. Figures in parenthesis are standard deviations.

Cycle papers have also had difficulty in replicating the high level of GDP persistence found in the data (Chari et al., 2002).

The model does not capture the strong degree of persistence in the short-term nominal interest rate. This echoes Ireland's (2003) results on US data. Over the post-1979 period, he finds that a flexible-price DSGE model has difficulty in matching the interest rate persistence, and notably as compared to a sticky-price model. This suggests to amend the model to include nominal rigidities, which is left for further research.

On the whole, Tables 4–6 confirm the quality of our estimation, since the model is able to match the vast majority of the 57 empirical moments. In particular, the estimated LP model is able to reproduce the cyclical properties of the exchange rate: the Gallant and Tauchen statistics for the variance and autocovariance of the exchange rate are far below the critical values, as well as the sub-set of covariances between the exchange rate and the other variables. The large estimated volatility of the foreign inflation shock suggests that this shock has a significant impact on exchange rate volatility, which will confirmed in the following sections.

4.2. Impulse response functions

This section describes the dynamics of some key variables following technological and monetary shocks. We evaluate the ability of the liquidity model to generate the dynamics implied by each type of shock consistent with the empirical effects identified in the VAR literature.

4.2.1. Dynamics following technological shocks

Fig. 2 shows the impulse response functions of the nominal exchange rate and the nominal interest rate (in level), as well as the real exchange rate and domestic output (in % deviation from steady state) following a 1% increase in the technological growth rate occurring in period 1. Fig. 3 completes the description of the effects of a positive supply shock by illustrating the IRFs of consumption, investment, hours and the trade balance (in % deviation).

As shown in Fig. 2, the nominal exchange rate depreciates on impact and the nominal interest rate immediately jumps beyond its reference level, then declines.

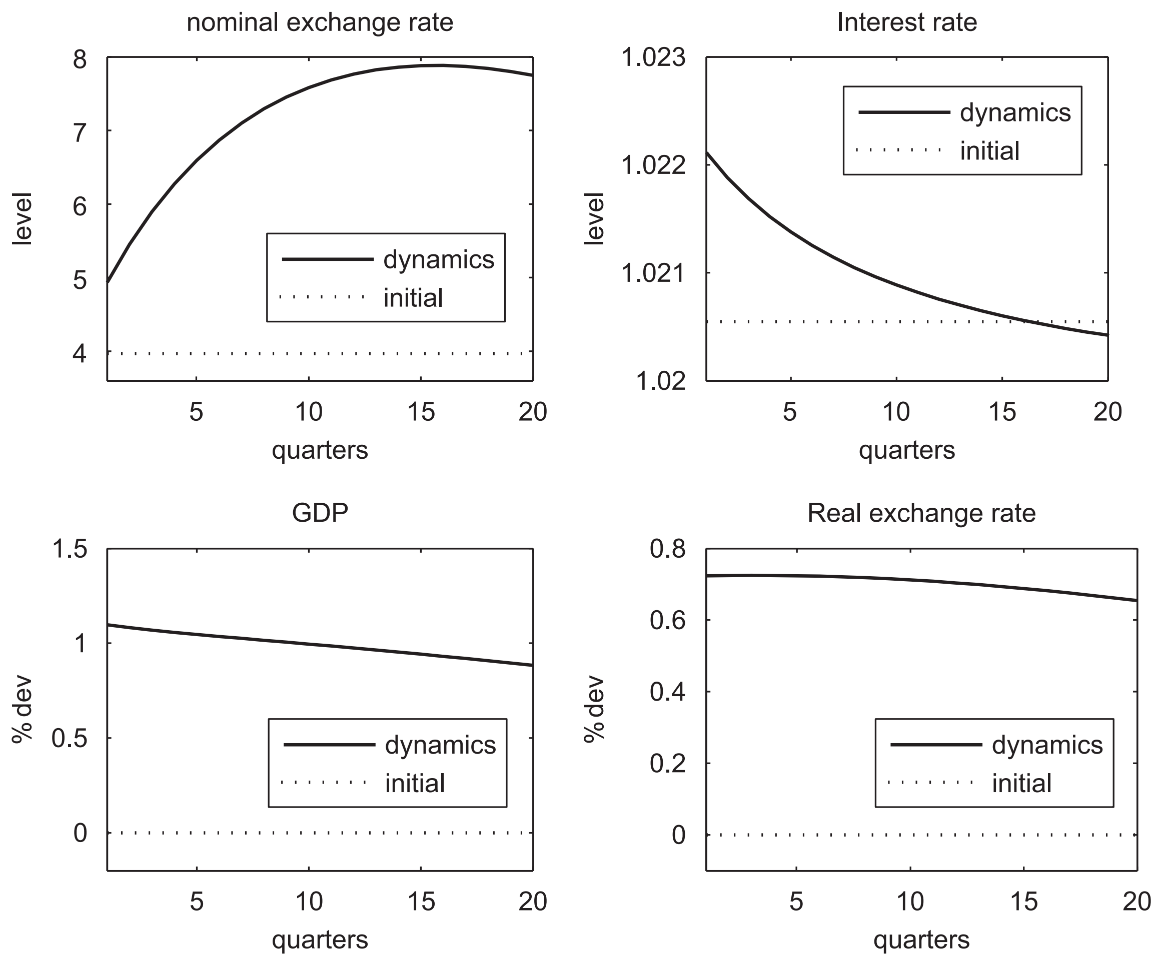

Fig. 2. Effects of a positive technological shock (1).

The interest rate response mainly stems from the behavior of monetary growth (Eq. (19)). The estimation results show that the monetary authorities react to an increase in g by raising the nominal interest rate. As the technological change yields a positive wealth effect for the household, consumption persistently increases (see Fig. 3) and the demand for money (M^c_{t+1}) rises. Following the technological shock, the endogenous monetary g therefore increases (Eq. (18)), and this drives the Central Bank to raise the interest rate in accordance with the estimated interest rate rule. The joint dynamics of e and R are consistent with UIRP: the persistent increase in the domestic interest rate, and consequently in the interest rate spread, has to be offset by an expected depreciation of domestic currency.

The model displays IRFs regarding output and the real exchange rate that are consistent with the predictions of the standard Mundell–Fleming–Dornbusch model, i.e. a real depreciation of domestic currency and an increase in production.[25]

[25]The model also consistently predicts a fall in domestic prices, not displayed here. For reasons of space, we do not present all the IRFs to real and nominal shocks in the paper; but these are available on request from the authors.

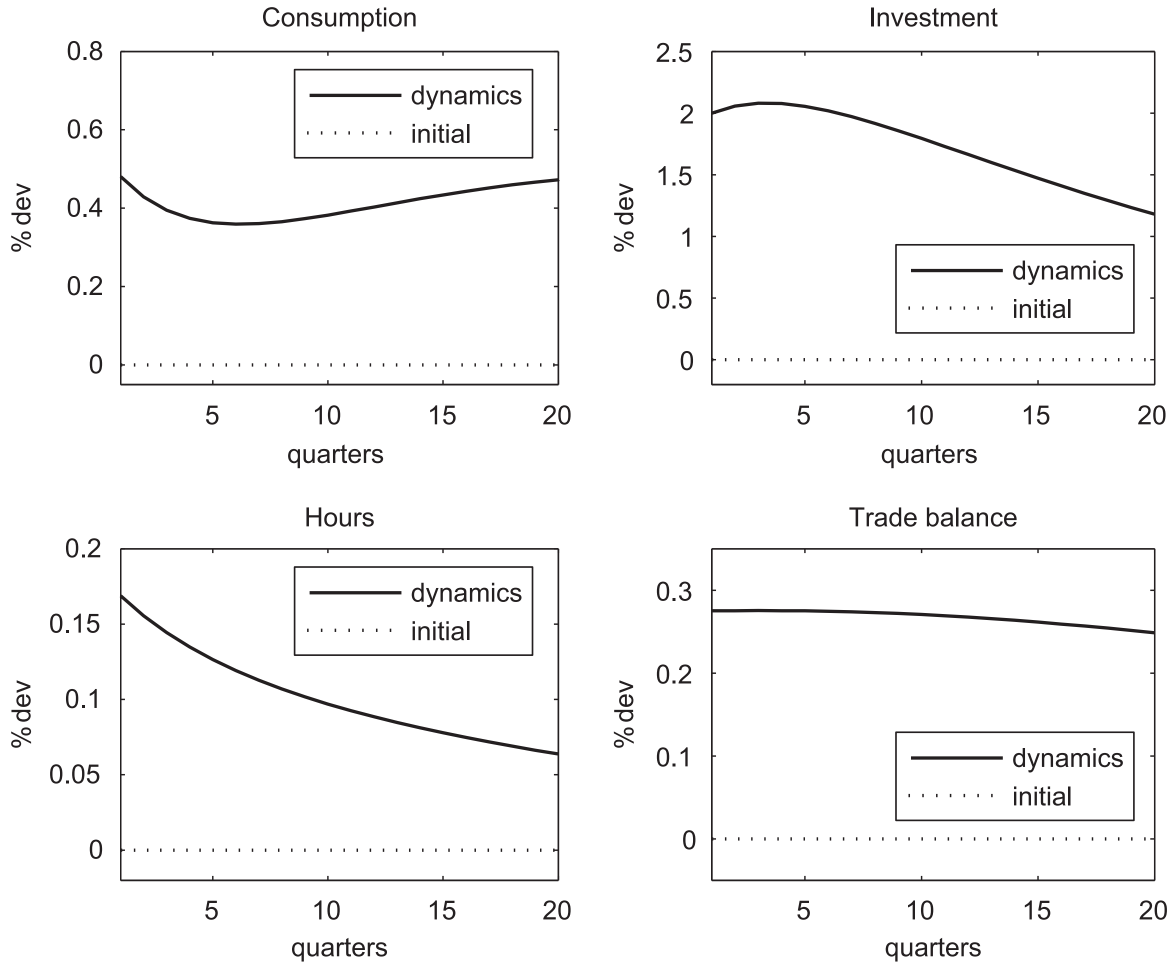

Fig. 3. Effects of a positive technological shock (2).

The nominal exchange rate depreciation (combined with the reduction in prices) results in a depreciation of the real exchange rate.

As shown in Fig. 3, a positive technological shock generates an increase in employment, investment and consumption. Investment rises after the technological shock despite the rise in the nominal interest rate: the positive effect of the supply shock (greater marginal productivity of capital) dominates the negative effect of the jump in borrowing costs. On the labor market, labor demand increases for similar reasons, while the household tends to reduce its labor supply due to the positive wealth effect. As a result, hours worked increase in equilibrium (with a corresponding rise in the real wage).

How well do these IRFs correspond to the empirical facts? Corsetti et al. (2004) estimate a VAR on US data relative to the aggregate of other OECD countries. They find that, following a home technological improvement, US output and consumption (relative to foreign) rise. This is consistent with our findings. The rise in consumption is as large as that in output in the data. We do not match this feature because of the consumption smoothing inherent in our model.

If the impact of supply shocks on GDP is well established in the empirical literature, this is not the case regarding employment. Gali's (1999) result of a negative impact is at the heart of renewed debate on the topic. As noted by Chang and Hong (2006), this negative short-run impact of technology shocks on employment is interpreted as evidence in favor of the sticky-price model versus the flexible-price model; however, it is still an open issue and leads to the related question of the 'good' way of measuring productivity shocks (total factor productivity, TFP, versus labor productivity). Chang and Hong (2006) (among others) claim that TFP, rather than labor productivity, is the natural measure of technology. With this measure, they find that a positive supply shock positively affects hours worked in US data. In our flexible-price setting, we obtain results consistent with theirs.

The empirical literature is scarce on the effects of supply shocks on trade and the real exchange rate and gives ambiguous results. Clarida and Gali (1994), using VAR decomposition, find that the impact of a positive supply shock on the real exchange rate of the US depends on the countries considered (depreciation for US–Japan, and appreciation for US–UK and US–Germany). Corsetti et al. (2004) find that a positive supply shock leads to an appreciation of the US real exchange rate and a fall in net exports. On the contrary, our model predicts real exchange rate depreciation and an improvement of net exports. However, it is worth noting that Corsetti et al. (2004) themselves underline that their empirical results for the US (a very large and relatively closed economy) do not necessarily generalize to smaller and/or more open-economies.[26] The absence of any clear-cut result in the area suggests that more empirical research is needed regarding the open-economy effects of supply shocks.

4.2.2. Dynamics following a positive monetary shock

We now evaluate the dynamics of key variables in the model following a positive monetary shock, that is 1% *decrease* in the interest rate innovation ε_R, occurring in period 1. We present the impulse response functions of the nominal interest rate R and the nominal exchange rate e (in levels) in Fig. 4.

As shown in Fig. 4, the nominal interest rate initially falls with the nominal shock, and then slowly rises back towards its reference level. Since the estimated value of ρ_π is rather small, the Central Bank focuses more on smoothing the interest rate and reacting to money growth g, both of which contribute to the persistence of the liquidity effect.

- Following the fall in the nominal interest rate, the Bundesbank operates a smooth adjustment ($\rho_r = 0.440$).
- Furthermore, adjustment costs scaled by ξ limit the increase in money holdings that could counteract the persistent decline in the nominal interest rate. In the period of the monetary shock, the household chooses the amount of money that it

[26]Their underlying structural model suggests that the wealth effects at the heart of the negative transmission of supply shocks are fairly unlikely in such economies.

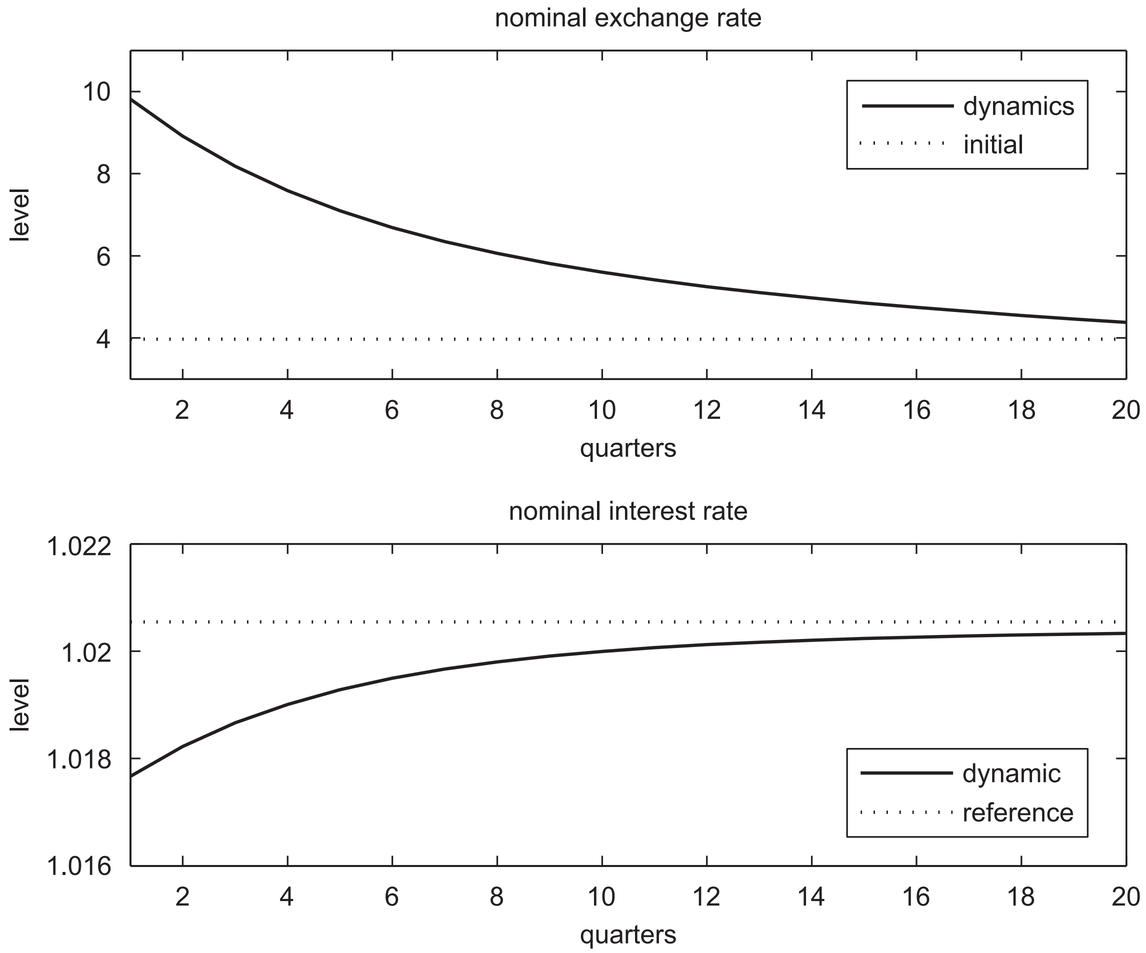

Fig. 4. Effects of a positive monetary shock (1).

wants to consume tomorrow (M^c_{t+1}). After the money shock, the agent anticipates inflation: the household would like to preserve future consumption by increasing the current amount of nominal money balances. However, it is costly for the household to raise the ratio M^c_{t+1}/M^c_t substantially. Changing M^c_{t+1} deprives the agent of time available for leisure or labor. With larger adjustment costs, it is more expensive to modify money holdings today and the household would rather wait. As a result, in the first period, the household raises the amount of money-cash M^c_{t+1} by a small amount. Adjustment costs on money holdings limit the changes in M^c_{t+1} for consumption purchases, thereby slowing down the rise in g. This latter could increase the nominal interest rate R via the Taylor rule, thus reducing the magnitude of the liquidity effect. In a nutshell, adjustment costs scaled by ξ contribute to the persistent fall of the nominal interest rate.

- In addition, the magnitude of the liquidity effect also depends on the demand for loans, i.e. investment decisions. If firms invest more, the resulting increase in the demand for loans might push the interest rate upward, thereby reducing the liquidity effect. This phenomenon is scaled back by adjustment costs of capital Ψ_I. The higher is Ψ_I, the larger is the liquidity effect.

The analysis of impulse response functions is consistent with the estimation findings, which showed that interest smoothing (ρ_r) and credit market frictions (ξ) are statistically significant.

Consider now the IRF of the nominal exchange rate (Fig. 4). On impact, the nominal exchange rate depreciates beyond its steady-state value, then appreciates. The LP model generates nominal exchange rate overshooting, the empirical relevance of which has been largely investigated in the literature. With the seminal paper by Eichenbaum and Evans (1995), the first results were not in favor of the standard overshooting assumption. They showed that, in response to a tighter US monetary policy, the US dollar exhibits a delayed overshooting pattern of 2–3 years *vis-à-vis* the major currencies. Positive interest rate differentials are associated with persistent appreciation of the US dollar: this delayed overshooting is not consistent with the Dornbusch model and UIRP. This topic has generated a large amount of debate in the empirical literature, which does not offer any clear-cut consensus. Faust and Rogers (2003) (applying Faust's, 1998 identification techniques to the US/UK and US/DM exchange rates) find that delayed overshooting is quite sensitive to dubious identifying assumptions. While they reject delayed overshooting as a robust finding, they also reject standard overshooting *à la* Dornbusch. Other papers, whether using alternative methods of measurement of monetary policy (Bonser-Neal et al., 1998, using an event-study methodology, or Kalyvitis and Michaelides, 2001, using the monetary policy indicator proposed by Bernanke and Mihov, 1998) or alternative VAR decomposition schemes (Kim and Roubini, 2000) obtain results more in line with Dornbusch. A related question is whether evidence of delayed overshooting should be considered as evidence of a failure of UIRP. Mac Callum (1994) and Bjornland (2005) point out that endogenous interest-rate setting can generate delayed overshooting in a environment where UIRP holds. To summarize, the empirical literature has uncovered evidence of overshooting, but there is little consensus on the timing of this overshooting or on whether delayed overshooting can be considered as a stylized fact which casts doubt on the validity of UIRP.

Leaving to one side the question of delay, we focus here on rationalizing the stylized fact that the nominal exchange rate overshoots after monetary shocks. In a setting where interest rate parity holds, plausible monetary models should thus be consistent with the following facts: a positive domestic money shock generates a strong and persistent liquidity effect, a persistent negative interest rate differential and nominal exchange rate overshooting. As illustrated in Fig. 4, the dynamics of the nominal interest rate and the nominal exchange rate predicted by the model are consistent with the effects of monetary innovations observed in the data.

The mechanisms behind nominal exchange rate overshooting are straightforward. This stems from the interest rate response combined with UIRP (Eq. (24)). The persistent decline in the interest rate results in a negative interest rate spread that is offset by an expected appreciation of the home currency. The nominal exchange rate accordingly displays overshooting dynamics *à la* Dornbusch, even though our model differs from the traditional one. The seminal Dornbusch (1976) model and the New Open Economy Macroeconomics literature (Obstfeld and Rogoff, 1995) stress price rigidities combined with UIP as the main ingredients of exchange rate dynamics.

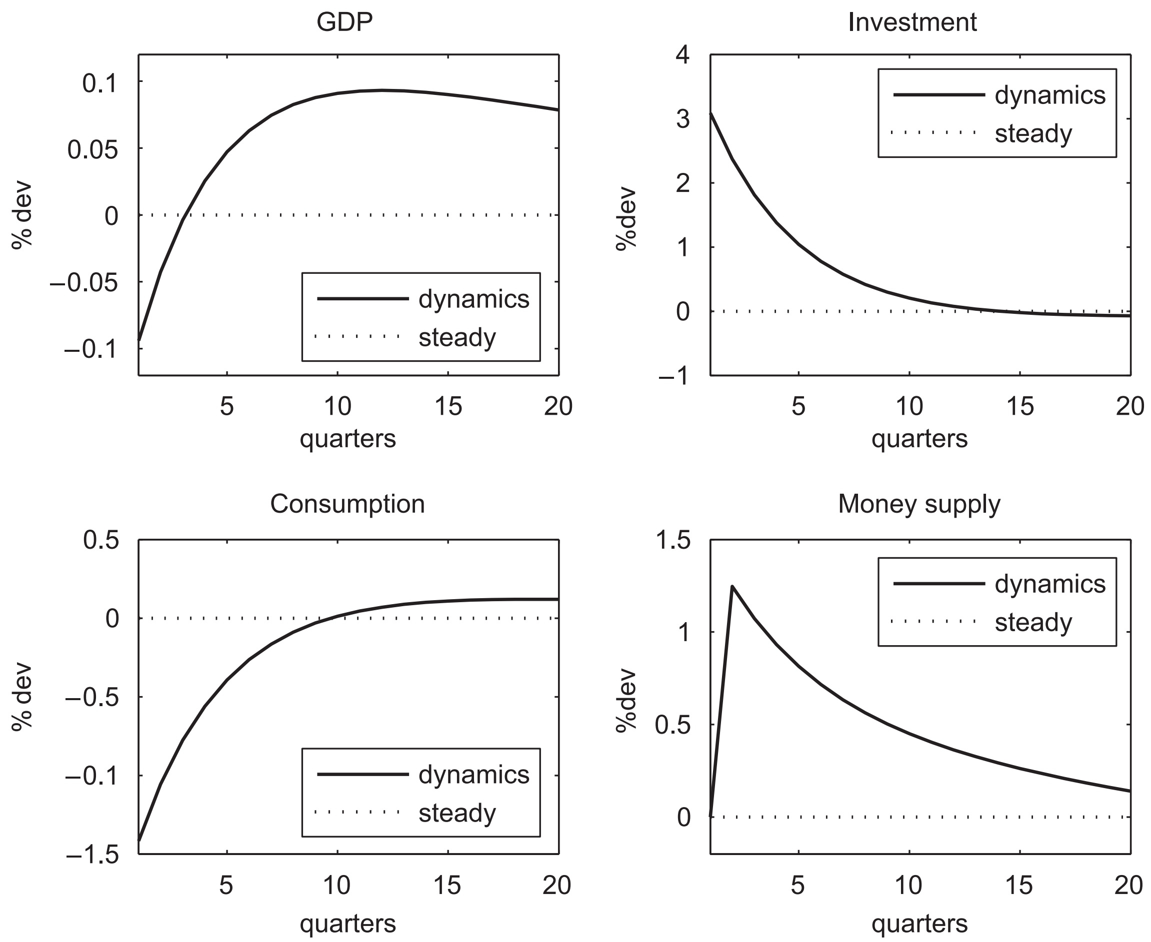

Fig. 5. Effects of a positive monetary shock (2).

In contrast, in our flexible-price setting the key rigidity lies in the credit market through LP and portfolio adjustment costs. However, both models introduce rigidities on the market where the nominal interest rate is determined, so that a home monetary expansion generates a large and persistent decrease in the domestic interest rate (and the interest rate differential) and subsequent nominal exchange rate overshooting, given UIRP. In addition, the traditional definition of overshooting *à la* Dornbusch is tied to the larger response of the nominal exchange rate as compared to that of the money supply after a money shock. This is the case in our model, as can be seen in Figs. 4 and 5.

Then, next step is to focus on the effects of nominal shocks on a larger set of macroeconomic variables. Fig. 5 shows the impulse response functions of output, investment, consumption and money supply to a positive home money shock (in deviations from steady state).

On impact, the output response is negative. This comes from the instantaneous decline in employment (not shown here), the capital stock being fixed in the first period. With the monetary shock, the household escapes the inflationary tax by consuming less in order to have more leisure; employment decreases and so does output. Nevertheless, after a few periods, the output response becomes positive, since the rise in capital (associated with the investment boom) compensates for the decline in labor. Investment reacts positively to the monetary shock: since the cost of borrowing persistently remains below the reference level, firms are enticed to invest more. Consumption is negatively affected by the money shock due to the combination of the Fisher effect (as prices increases with the nominal shock) and the CIA constraint (Eq. (7)). Money supply rises substantially above its initial value due to the exogenous decrease in the interest rate. It then decreases towards its steady-state value as the effects of the *iid* monetary innovation vanish over time.

The survey of Christiano et al. (1999) concludes that, despite the various techniques of identifying money shocks, there is reasonable consensus on the following effects of shocks: after a fall in the nominal interest rate, output ultimately rises, and money supply increases as does consumption and investment, the latter being more responsive to the initial shock. Schlagenhauf and Wrase (1995) and Christiano and Eichenbaum (1992b) identify the output impulse response function following a monetary expansion in G7 countries. They show that for all countries output initially falls but then increases six or nine months after the nominal interest rate drop. Kashyap and Mojon (2003) obtain similar VAR results in European countries. In quantitative terms, the model predicts that the GDP response is equal to one tenth of the interest rate change: for a 100 basis point fall in the (annualized) nominal interest rate, output increases by around 0.1%. On the one hand, the model accounts for the hump-shaped GDP response to the interest rate shock, with a peak at 6 quarters in German data versus 8 quarters in the model. On the other hand, the magnitude of the relative output to interest rate response is lower than the empirical results obtained by Kashyap and Mojon (2003) on the GDP effects of money shocks in Europe.[27] The peak GDP value is found in Kashyap and Mojon (2003) to be equal to half the interest rate change in Germany and France, and one third in the Euro area (Peersman and Smets, 2003; Mojon and Peersman, 2003).[28] This result suggests that the LP model should be modified to include other propagation mechanisms, such as price rigidities, in order to magnify the output response. This is left for further research.

Our results are consistent with the view that the interest rate channel is important in Germany and Europe, as work on microeconomic and macroeconomic data has shown that a decrease in nominal interest rates leads to a rise in investment (Von Kalckreuth, 2001; Chatelain et al., 2001). Empirically, consumption is found to increase as well (even though it reacts less than investment), while the model predicts the opposite. This counterfactual result is a well-known limit of models that introduce money through a CIA constraint, as previously shown in a closed-economy setting by Christiano (1991).

[27]We particularly consider the effects in Europe since our model is based on German data. Kashyap and Mojon's (2003) book is available as a series of working papers of the European Central Bank. We refer here to some of them.

[28]Also note that, in Locarno et al. (2001) and Mac Adam and Morgan (2001), macroeconometric models find GDP responses of 0.29–0.33 after a 100 basis point nominal interest rate change for France and Germany, with the peak response occurring 2 years after the shock.

To summarize, in spite of the model's shortcomings with respect to the consumption and output responses to money shocks, the LP model is able to reproduce the key empirical effects of monetary and technological shocks, regarding both nominal and real variables. This section confirms the empirical relevance of the model, since it was estimated from moments rather from conditional responses. The next section goes further in assessing the performance of the liquidity model in terms of variance decomposition.

4.3. Forecast error variance decompositions

Forecast error variance decompositions measure the relative weight of each shock in accounting for the dynamics of the key variables in the model. Table 7 presents

Table 7
Variance decomposition

h	Shocks				
	A	π^*	v	ω	ε_R
Nominal exchange rate e					
1	0.0011	0.6817	0.0001	0.3159	0.0013
4	0.0036	0.7044	0.000	0.2912	0.0009
8	0.0073	0.7302	0.000	0.2619	0.0006
12	0.0106	0.7525	0.000	0.2365	0.0004
16	0.0130	0.7721	0.000	0.2145	0.0003
20	0.0145	0.7897	0.000	0.1955	0.0003
Nominal interest rate R					
1	0.3923	0.000	0.0685	0.4951	0.0442
4	0.3553	0.000	0.0250	0.5840	0.0357
8	0.3017	0.000	0.0170	0.6527	0.0286
12	0.2630	0.000	0.0143	0.6977	0.0249
16	0.2373	0.000	0.0129	0.7270	0.0228
20	0.2218	0.000	0.120	0.7447	0.0214
Inflation π					
1	0.2898	0.0001	0.0034	0.6227	0.0840
4	0.2895	0.0001	0.0129	0.6100	0.0784
8	0.2951	0.0001	0.0121	0.6187	0.0740
12	0.2877	0.0001	0.0116	0.6292	0.0714
16	0.2810	0.0001	0.0113	0.6379	0.0696
20	0.2760	0.0001	0.0112	0.6443	0.0684
Consumption C					
1	0.2048	0.0001	0.2956	0.4401	0.0594
4	0.2098	0.0001	0.0990	0.6512	0.0398
8	0.1805	0.0001	0.0490	0.7495	0.0210
12	0.1640	0.0001	0.0300	0.7931	0.0128
16	0.1555	0.0001	0.0203	0.8153	0.0088
20	0.1502	0.0001	0.0148	0.8285	0.0064

our results on the exchange rate, the interest rate, inflation and consumption over the short and medium run (from quarter $h = 1$ to 20).

Ireland and Papadopoulou (2004) estimate a closed-economy model with LP and sticky prices on US data. Our results on consumption are consistent with their findings on variance decomposition. Shocks to velocity, technology and preferences mostly drive consumption changes, while this is barely affected by money shocks. All nominal variables are affected by the shock to the Taylor rule. However, the quantitative importance of this shock remains limited. The policy shock explains only a very small fraction of interest rate changes, as well as changes in consumption and inflation. With respect to Ireland (2004) or Lubik and Schorfheide (2007), the role of Taylor rule innovations in variance decompositions is limited. Yet, this is not necessarily a surprising result given the relatively low standard deviation of the shock, compared to the others. Moreover, the result of a limited quantitative importance of the policy shock is also found by Ireland (2004) and Ireland (2003) in the post-1980 period on the US interest rate. Our estimates confirm Leeper et al. (1996) view that '*most movements in monetary policy instruments are responses to the state of the economy, not random deviations by the monetary authorities from their usual patterns of behavior*'.

Table 7 shows that monetary shocks play a minor role in accounting for exchange rate dynamics, which is in line with the VAR literature (Faust and Rogers, 2003). The foreign inflation shock plays a key role in accounting for exchange rate variance, in line with Lubik and Schorfheide's (2007) findings on Canadian data. This phenomenon may be related to the law of one price hypothesis that prevails in our setting. In spite of flexible prices, in order to match exchange rate volatility as well as moments on German inflation, the shock to foreign inflation must be very volatile, thereby playing a critical role in matching the exchange rate standard deviation. This echoes our previous analysis (Section 4.1) that the estimated process on π^* captures elements beyond 'pure' foreign inflation that helps to account for exchange rate movements.

In order to explore the other potential mechanisms that might be important in explaining exchange rate changes, the model should discard the law of one price hypothesis by, for instance, exploring pricing-to-market behavior in addition to the LP model. This is left for future research.

5. Assessing the role of credit market frictions

This section is devoted to the analysis of the role of credit market frictions. To do so, we compare the predictions of the benchmark model (with adjustment costs $\xi > 0$ and timing restrictions on household's portfolio allocations) to those of the model without frictions and $\xi = 0$, i.e. a simple CIA model. We gauge the role of credit market frictions via two criteria: (i) the dynamics implied by a money shock; and (ii) the global fit of the models to the data.

5.1. Credit market frictions as a transmission channel of money shocks

As detailed in Section 2.1, in the CIA model both the timing restrictions on household's portfolio decisions and portfolio adjustment costs are suppressed.

In period t, the household now decides the amount of money-deposit M_t^b in the current period, given the amount of money holdings M_t inherited from period $t-1$. This determines the amount of money-cash (M_t^c).

Hairault et al. (2004) study the performances of the CIA model in a small open-economy DSGE model. In line with the results of closed-economy papers (Christiano, 1991 among others), they show that the simple CIA model is unable to reproduce the stylized facts of a money shock, namely the liquidity effect of the interest rate. As well, the nominal exchange rate displays no overshooting dynamics. When investigating the IRFs of a interest rate shock in the CIA framework, we get very similar results. Moreover, the quantitative effects are negligible. This is also the case regarding real aggregates. Following monetary shocks, there is virtually no effect on real variables. This confirms the widespread view in the literature, that a model without monetary frictions cannot account for the effects of monetary policy shocks obtained in the VAR literature.

Next step is to assess the performances of the CIA model and the LP model in terms of global adequacy for the set of 57 moments to match.

5.2. The global (in)adequacy of the CIA model

In order to compare the fit of the LP and CIA models to the data, we apply the simulated method of moments to the CIA model, for the set of moments and the set of structural parameters identical to those considered for the model with credit market frictions (Estimation 2, Table 3). Table 8 sums up the results of the structural estimation.

As shown in Table 8, the two standard deviations tied to the behavior of money, σ_r and σ_v are estimated as equal to 0. This conveys the idea that they do not contribute to the global fit of the CIA model to the empirical data. These results fit with the conclusion drawn from the IRF analysis: money hardly affects real aggregates in the CIA model.

Table 8
Structural parameter estimation, CIA model

Adj. costs on capital Ψ_I	7.2563	Taylor rule volatility σ_R	0.00014
	(2.729)		(0.0096)
Technology shock		Taylor rule	
Persistence ρ_a	0.9993	Money growth ρ_g	0.6172
	(58.31)		(1.900)
Volatility σ_a	0.00643	Interest rate ρ_r	0.3846
	(5.541)		(1.152)
Pref. shock volatility σ_ω	0.01104	Inflation ρ_π	0.0000
	(13.29)		(0.000)
Foreign inflation volatility $\sigma_{\pi*}$	0.04807	Velocity shock volatility σ_v	0.0000
	(8.999)		(0.000)

Note: t-stats. are displayed in parentheses.

Tables 9 and 10 present the set of empirical moments simulated by each model (in the columns denoted 'LP' and 'CIA', respectively) altogether with the empirical moments (in the column 'Data').[29]

As shown in Table 9, nominal exchange rate volatility is well reproduced by the CIA model, which confirms the role of foreign inflation shocks in our framework. However, we reject the CIA model as appropriate in capturing exchange rate dynamics. First, the exchange rate does not respond at all to money shocks, and the CIA model does not yield overshooting *à la* Dornbusch (1976). Both elements are counterfactual. Beyond the single dimension of exchange rate dynamics, the comparison of the global adequacy of the CIA and the LP models confirms this conclusion. First, we reject the CIA model as appropriate, since it does not satisfy the over-identification test *à la* Hansen with a p-value equal to 1.6%. Secondly, the CIA model does not outperform the LP model along the dimensions where the LP model is worse, i.e. the two moments that do not satisfy the GT test at 5%. In addition, 11 other moments simulated by the CIA model are not statistically close enough to their empirical counterparts (at 5%). Notably, as shown in Tables 9 and 10, performance is worse regarding investment and the trade balance.

This delivers an interesting message. In the absence of any shock to the Taylor rule (and in money demand as well), the CIA model misses one key source of investment and trade balance dynamics. The underlying mechanisms are the following. As underlined by Hairault et al. (2004), in the CIA framework the response of the nominal interest rate obeys its Fisherian fundamentals. Unlike the LP framework, the household is free to choose between money-cash and money-deposits within the current period. As a result, in the period of the monetary shock, the household reduces her bank deposits to favor money-cash, so as to preserve consumption from expected inflation. Given the withdrawal of private funds, the supply of loans ultimately falls and the interest rate, i.e. the cost of investment, rises, leading to a decrease in investment. If any, the effects of the money shock are thus counterfactual. As a result, the CIA framework is not able to capture the role of money shocks on investment. Yet, the importance of this typical Keynesian transmission mechanism of monetary policy is confirmed empirically: according to the VAR literature (see Kashyap and Mojon, 2003), monetary shocks have a larger impact on investment (as compared to other real aggregates, notably consumption). Given the smoothing behavior of consumption, the dynamics of trade balance may be mostly attributable to that of investment. The failure of the CIA model with respect to investment dynamics implies similar difficulty regarding trade balance.

Overall, our results suggest that the absence of credit market frictions worsens the empirical performance of the small open-economy model. This allows us to conclude that credit market frictions play a significant role in accounting for the dynamics of

[29]In Tables 9 and 10, we present the 'true' moments used in the estimation procedure (i.e. variance, autocovariance, covariances) so as not to bias the comparison of results between the two models. Specifically, we base this comparison on the confidence interval around the statistical moment simulated by each model (plus and minus one standard deviation).

Table 9
Models comparison (1)

Moment	Data	LP	CIA
Variance (10^{-3})			
$\sigma^2(\widehat{e})$	6.9916	6.2580	**5.1477**
$\sigma^2(\widehat{y})$	0.2268	0.1452	**0.0745***
$\sigma^2(\widehat{i})$	1.6870	1.3578	**0.4301***
$\sigma^2(\widehat{c})$	0.1292	0.1054	**0.0904**
$\sigma^2(\widehat{p})$	0.1115	0.1147	**0.1280**
$\sigma^2(\widehat{TB})$	0.1095	0.0891	**0.0613***
$\sigma^2(\widehat{r})$	0.0176	0.0062	**0.0035***
$\sigma^2(\widehat{m_1})$	0.5516	0.3254	**0.1094***
Autocovariance (10^{-3})			
$Cov(\widehat{e},\widehat{e}_{-1})$	5.8676	4.4150	**3.6389**
$Cov(\widehat{y},\widehat{y}_{-1})$	0.1918	0.1024*	**0.0541***
$Cov(\widehat{i},\widehat{i}_{-1})$	1.3762	0.9844	**0.3064***
$Cov(\widehat{c},\widehat{c}_{-1})$	0.0993	**0.0568**	0.0664
$Cov(\widehat{p},\widehat{p}_{-1})$	0.1051	**0.0820**	0.0962
$Cov(\widehat{TB},\widehat{TB}_{-1})$	0.0841	0.0615	**0.0433***
$Cov(\widehat{r},\widehat{r}_{-1})$	0.0155	0.0040*	**0.0025***
$Cov(\widehat{m}_1,\widehat{m}_{1,-1})$	0.4129	0.3507	**0.1185***
Instantaneous covariance (10^{-3})			
$Cov(\widehat{e},\widehat{i})$	−0.6509	**−1.0456**	−0.6569
$Cov(\widehat{e},\widehat{c})$	−0.3952	**0.2856**	−0.2879
$Cov(\widehat{e},\widehat{p})$	0.4915	**0.3095**	0.3220
$Cov(\widehat{e},\widehat{TB})$	0.2976	**0.3379**	0.2614
$Cov(\widehat{e},\widehat{r})$	−0.0380	**−0.0696**	−0.0606
$Cov(\widehat{e},\widehat{m}_1)$	−0.3659	−0.0730	**0.0030**
$Cov(\widehat{y},\widehat{i})$	0.4845	0.3041	**0.1071***
$Cov(\widehat{y},\widehat{c})$	0.1231	0.0695	**0.0507**
Instantaneous covariance (10^{-3})			
$Cov(\widehat{y},\widehat{m}_1)$	0.0223	0.0452	0.0158
$Cov(\widehat{y},\widehat{p})$	−0.0625	−0.0543	−0.0470
$Cov(\widehat{i},\widehat{c})$	0.3625	0.2636	0.1985
$Cov(\widehat{i},\widehat{p})$	−0.2105	−0.2780	−0.2023
$Cov(\widehat{c},\widehat{p})$	−0.0745	−0.0718	−0.0852
$Cov(\widehat{c},\widehat{TB})$	−0.0286	−0.0373	−0.0363
$Cov(\widehat{p},\widehat{TB})$	0.0322	0.0327	0.0377

Note: * Denotes moments that do not pass the GT test at the 5% level. In bold characters, simulated moments that do significantly worse in one model than the other in matching the empirical ones.

a money shock. Beyond that dimension, they are a key ingredient in correctly reproducing the cyclical moments of a large set of variables, notably investment and trade balance, as shown in Tables 9 and 10.

Table 10
Models comparison (2)

Moment	Data	LP	CIA
Dynamic cross-covariance (10^{-3})			
$Cov(\widehat{e},\widehat{c}_{-1})$	−0.3984	**−0.1901**	−0.1957
$Cov(\widehat{e},\widehat{p}_{-1})$	0.4903	**0.2549**	0.2732
$Cov(\widehat{e},\widehat{TB}_{-1})$	0.2492	0.2369	**0.1802**
$Cov(\widehat{e},\widehat{r}_{-1})$	−0.0379	**−0.0501**	−0.0441
$Cov(\widehat{e},\widehat{m}_{1,-1})$	−0.2562	0.0101	**0.0556**
$Cov(\widehat{y},\widehat{e}_{-1})$	−0.1594	**0.0070**	−0.0111
$Cov(\widehat{y},\widehat{i}_{-1})$	0.4502	0.2226	**0.0793***
$Cov(\widehat{y},\widehat{p}_{-1})$	−0.0695	−0.0473	**−0.0398**
$Cov(\widehat{i},\widehat{e}_{-1})$	−0.5079	**−0.7521**	−0.4698
$Cov(\widehat{i},\widehat{c}_{-1})$	0.3234	0.2066	**0.1351**
$Cov(\widehat{i},\widehat{p}_{-1})$	−0.2218	−0.2478	**−0.1698**
$Cov(\widehat{c},\widehat{e}_{-1})$	−0.3544	**−0.2041**	−0.2138
$Cov(\widehat{c},\widehat{i}_{-1})$	0.3311	0.1984	**0.1453***
$Cov(\widehat{c},\widehat{p}_{-1})$	−0.0743	**−0.0574**	−0.0735
$Cov(\widehat{c},\widehat{TB}_{-1})$	−0.0230	−0.0244	**−0.0260**
$Cov(\widehat{m}_1,\widehat{c}_{-1})$	0.0856	0.0503	**0.0472**
$Cov(\widehat{p},\widehat{e}_{-1})$	0.4394	**0.1694**	0.1856
$Cov(\widehat{p},\widehat{i}_{-1})$	−0.1741	−0.1214	**−0.1119**
$Cov(\widehat{p},\widehat{c}_{-1})$	−0.0650	**−0.0349**	−0.0426
$Cov(\widehat{p},\widehat{TB}_{-1})$	0.0283	0.0212	**0.0199**
$Cov(\widehat{TB},\widehat{e}_{-1})$	0.3287	0.2471	**0.1951**
$Cov(\widehat{TB},\widehat{y}_{-1})$	0.0312	0.0418	**0.0182**
$Cov(\widehat{TB},\widehat{c}_{-1})$	−0.0258	**−0.0197**	−0.0255
$Cov(\widehat{TB},\widehat{p}_{-1})$	0.0291	0.0259	**0.0325**
$Cov(\widehat{r},\widehat{e}_{-1})$	−0.0413	**−0.0455**	−0.0423
$Cov(\widehat{m}_1,\widehat{e}_{-1})$	−0.4467	−0.1837	**−0.0673**

Note: * Denotes moments that do not pass the GT test at the 5% level. In bold characters, simulated moments that do significantly worse in one model than the other in matching the empirical ones.

6. Conclusion

This paper explores the empirical dimensions of NOEM models. Rather than focusing on monopolistic competition models with price and wage stickiness as estimated in previous papers (Ghironi, 2000; Bergin, 2003; Smets and Wouters, 2003 among others), we assess the empirical goodness of fit of a flexible-price LP model. We estimate the model on German data over the period 1971–1998 using the simulated method of moments.

First, we find that the LP model is statistically supported by the data. Credit market frictions are critical in the ability of the model to match a large set of moments. Further, the impulse response functions predicted by the model are consistent with the VAR findings regarding responses to real and nominal shocks, with the exception of consumption dynamics after an interest rate shock. Since the model is estimated on moments rather than on conditional IRFs, this point further underlines the ability of the model to match the data. Finally, monetary shocks do not seem to be the primary driving force behind the aggregate dynamics, which is consistent with the VAR literature. In order to illustrate the key role of credit market frictions, we estimate a CIA model without credit market frictions. We show that the CIA model is statistically rejected by the data. Moreover, real aggregates and exchange rate responses to a monetary shock are inconsistent with VAR evidence.

Foreign inflation shocks are found to greatly affect the exchange rate. This phenomenon is due to the fact that, with the law of one price, the model can match exchange rate volatility only through large changes in foreign inflation. In order to explore the impact of alternative shocks to the exchange rate, one could extend the model to include nominal rigidities and pricing-to-market within the LP model, thereby extending Ireland and Papadopoulou (2004) to an open-economy setting with price discrimination. Nominal rigidities would reduce the Fisher effect, thereby possibly generating a positive consumption response to the nominal interest rate fall, a dimension that our model fails to match. This is left for future research.

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Appendix A. Description of the data set

The original series come from the following sources.

- the OECD BSDB database, over the period 1970:1–1998:4, for GDP, investment, consumption, the short-term nominal interest rate, exports, imports, the nominal exchange rate and the consumption price index;
- the OECD Main Economic Indicators database, over the period 1970:1–1998:4 for the money supply.

The series for real aggregates are taken in volume (GDP, consumption, investment, imports and exports). We build the trade balance using German exports and imports of goods and services as detailed in Table 11. The nominal interest rate (coming from BSDB) is constructed on an annualized basis, while the model

Table 11
Data

Variables		Seasonality	Outliers	Treatment
Nominal exchange rate ($US per DM)	E	–	None	In log
GDP (at factor cost) less government wages	Y	SA	1991:1	In log
Gross fixed cap. form. business sector (broad definition)	I	SA	1991:1	In log
Private consumption	C	SA	1991:1	In log
Quarterly short term nominal interest rate +1	R	–	None	In log
Exports goods and services, N.A. basis	X	SA	1991:1	In log
Imports goods and services, N.A. basis	D_F	SA	1993:1	In log
Trade balance	TB	SA	None	–
M_1	M	SA	1973:2–1990:2	In log
Deflator of private consumption	P	SA	None	In log

Note: SA, seasonally adjusted. The trade balance has been calculated as exports less imports of goods and services, N.A. basis, volume.

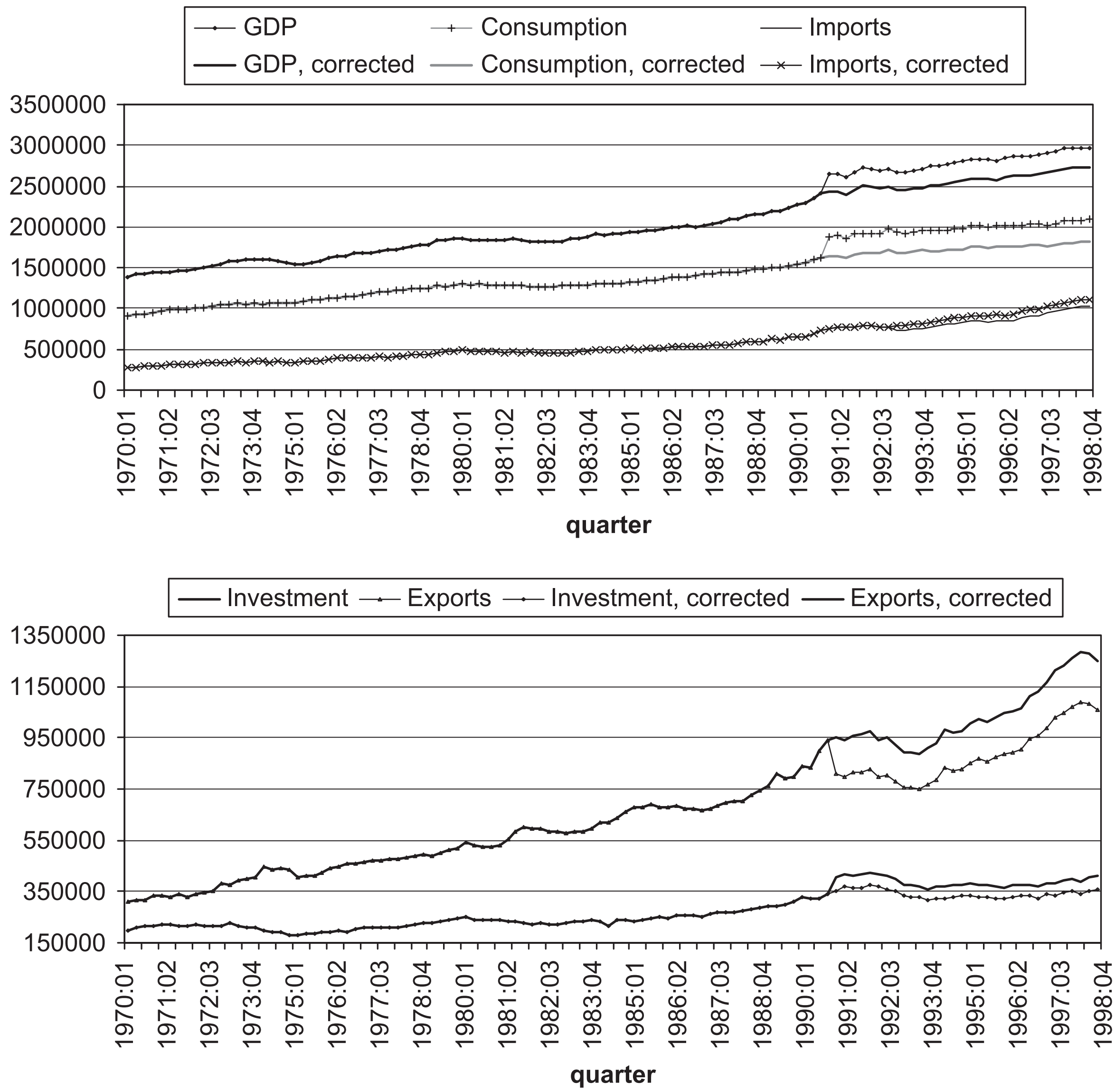

Fig. 6. Before and after outliers treatment.

delivers the quarterly interest rate. We accordingly modify the data to build the corresponding quarterly interest rate.

The set of 57 empirical moments was obtained after the following treatment of the original macroeconomic series. First, we handle outliers as follows. We convert the original series into growth rates. These series are then centered and normalized. As in Milliard et al. (1997), we consider outliers in these series as observations that do not lie in the interval $[-3, 3]$. The corresponding point in the series taken in growth rate is therefore replaced by averaging the closest growth rates. Finally we reconvert the series into levels. All aggregate time series have at least one outlier due to German reunification. Second, empirical moments are calculated on time series taken in logarithms (except trade balance) and de-trended using Hodrick and Prescott's (1997) methodology with the adjustment parameter set to 1600. Table 11 summarizes the details regarding the macroeconomic series and the treatment of outliers.

Fig. 6 illustrates the empirical original series for GDP, consumption, investment, imports and exports (in volume) over the period 1970:1–1998:4, as well as the corresponding series after the identification and removal of outliers, that are used in the estimation procedure.

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