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ESTIMATING THE EVOLUTION OF MONEY'S ROLE IN THE U.S. MONETARY BUSINESS CYCLE

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Abstract

We assess the *time-varying* money's role in the post-WWII U.S. business cycle by estimating a new-Keynesian framework featuring nonseparability in real balances and consumption, portfolio adjustment costs, and a systematic reaction of policymakers to money growth. Rolling-window Bayesian estimations *a la* Canova (2009) are contrasted to a full sample fixed-coefficient investigation. Our results suggest that the assumption of stable parameters is unwarranted. The omission of money may induce biased assessments on the impact of structural shocks to the U.S. macroeconomic aggregates, especially during the great inflation period.

JEL classification: E31, E51, E52.

Keywords: Money in new-Keynesian frameworks, time-varying effects, Bayesian structural estimation, Taylor rules.

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1 Introduction

Modern monetary New-Keynesian models of the business cycle typically consider money as a sideshow, i.e. the equilibrium values of inflation and output are determined without any reference to the stock of money.[1] In fact, a variety of recent empirical contributions challenge this view. Single-equation estimations supporting the role of money in explaining inflation and/or output for the U.S. are provided by Koenig (1990), Meltzer (2001), Nelson (2002), Hafer, Haslag, and Jones (2007), Reynard (2007), Hafer and Jones (2008), and D'Agostino and Surico (2009). Canova and de Nicoló (2002), Leeper and Roush (2003), Sims and Zha (2006), and Favara and Giordani (2009) employ multivariate SVARs models and find that 'LM' shocks exert significant effects on prices and the business cycle. Also in the light of the recent liquidity easing implemented by a variety of central banks in the attempt to tackle the real effects of the financial turmoil, a reconsideration of the role of money in monetary policy frameworks is clearly needed.

This paper estimates a *structural* DSGE monetary model of the business cycle in which money is allowed, but not necessarily required, to play a relevant role. In our model, money may exert 'nonseparability', 'direct', and 'policy' effects. Nonseparability between consumption and real balances affects intratemporal choices, the real wage (via labor supply) and, consequently, marginal costs and inflation. It also affects households' intertemporal rate of substitution of consumption, so modifying the Euler equation for output (Ireland (2004)). The direct effect arises when portfolio adjustment costs, which are modeled as a direct loss of agents' utility, are present. They justify a lag and enhance the role of expectations in the money demand equation. Moreover, they trigger an informational role for contemporaneous real balances as regards future realizations of the natural real interest rate, so rendering money relevant at low frequencies (Nelson

[1] For a detailed exposition of the New-Keynesian monetary policy model of the business cycle, see Woodford (2003) and Galì (2008).

(2002)). The policy effect captures the systematic reaction by policymakers to the evolution of the growth rate of nominal money, a reaction that may be welfare-enhancing if money concurs to determine the equilibrium values of inflation and output.[2]

In particular, our exercise is designed to pin down the possibly *time-varying* role played by money in shaping the post-WWII U.S. macroeconomic dynamics. Indeed, preferences over money-consumption non-separability may very well be unstable over time. Structural relationships involving money and the natural interest rate are likely to have been affected by financial innovations. A drifting emphasis on monetary aggregates by the FOMC may have taken place in the attempt of moving from the great inflation occurred in the 1970s to a more stable macroeconomic environment. Accounting for the possibly *evolving* role played by money is then key to achieve a correct identification of the (time-dependent) drivers of U.S. inflation and output. We tackle this issue by recursively estimating a small scale new-Keynesian DSGE model with Bayesian techniques, an approach recently proposed by Canova (2009).[3] This methodology allows us to investigate parameter instabilities without appealing to the combination of perturbation methods/particle filter recently proposed by Fernández-Villaverde and Rubio-Ramírez (2007). While being potentially very powerful and econometrically neat, their methodology forces the econometrician to stick to a limited number of time-varying parameters. In contrast, Canova's (2009) strategy is suited to account for instabilities in (possibly) *all* the estimated parameters. We see our approach as a first exploration of parameter instabilities in a small scale DSGE model with money.

Our results read as follows. A post-WWII full sample exercise conditional on *fixed coefficients* offers support to the role of money. This support mainly comes from portfolio adjustment costs and the Fed's systematic reaction to money growth. Interest-

[2]The systematic reaction to the money growth rate by policymakers may also be justified with money growth targeting *per se* (Svensson (1999)).

[3]Canova (2009) explores instabilities in the post-WWII U.S. sample with a small-scale DSGE model in which, by assumption, money does not play any active role.

ingly enough, a much richer picture emerges when analyzing the data through the rolling-window lenses. Indeed, our recursive estimations reveal that *money's role is time-dependent*. In particular, we find support for non-separability in the 1970s, along with a strong(er) monetary policy reaction to movements in monetary aggregates. The preference for non-separability, which calls for the structural presence of money in the price and quantity schedules, drops dramatically when entering the 1980s, and plays a marginal role afterwards. Similarly, monetary policy turns out to be less reactive to money growth in the 1980s and 1990s. The relevance of the portfolio adjustment costs is estimated to be fairly stable over time. We also find time-dependence for other 'structural' parameters, a notable example being the degree of habit formation. Furthermore, we find sample-dependence of the impact of money demand as well as other structural shocks. 'LM' shocks are estimated to significantly influence inflation and output in the 1970s. The presence of the stock of money also affects the identification of the other structural shocks (to households' preferences, technology, and monetary policy). Differently, money's impact appears to be moderate in the great moderation sample. Overall, our estimates support the role of money as an important ingredient to describe the post-WWII U.S. macroeconomic dynamics.

Before moving to the next Section, we make contact with some strictly related literature. Working with a microfounded new-Keynesian framework, Ireland's (2004) seminal paper relaxes the typically imposed nonseparability assumption by allowing the cross-derivative of the utility function with respect to consumption and real balances to be non-zero. Dealing with 1980s and 1990s U.S. data, he cannot reject the null of separability, and concludes that the role of money, if any, is negligible. With a richer model embedding habit formation and a systematic reaction of the Fed to money, Canova and Menz (2009) perform an international analysis involving the U.S., the U.K., the Euro area, and Japan, and find support for nonseparability in these countries.

Andrés, López-Salido, and Vallés (2006) consider a model with habit formation and price indexation, and confirm Ireland's (2004) results with Euro-data. Andrés, López-Salido, and Nelson (2009) find empirical support in favor of portfolio adjustment costs for the U.S. and the Euro Area with a model encompassing Andrés et al's (2006) and Canova and Menz's (2009).

There are several differences between these contributions and ours. First of all, our investigation is designed to detect the possible (in)stability of money's role over time. In the light of the institutional and technological changes occurred in the sample under scrutiny, information on the time-dependence of the role of money is clearly of high interest for a better understanding of the drivers of the post-WWII U.S. macroeconomics dynamics. Second, in conducting our analysis we employ Bayesian techniques, which allow for model comparison even in case of misspecified models (An and Schorfheide (2007) and Canova (2007)), a likely scenario when dealing with small-scale DSGE models. Moreover, they are superior to alternatives such as indirect inference, as least as regards new-Keynesian frameworks (Canova and Sala (2009)). Finally, in our investigation we employ the model recently put forward by Andrés, López-Salido, and Nelson (2009), which encompasses most of the previously scrutinized frameworks.

The structure of the paper is the following. Section 2 presents the new-Keynesian monetary policy framework with money we focus on in conducting our empirical analysis, briefly discusses the theoretical relevance of the restrictions of interest, and offers details on our empirical strategy. Section 3 discusses our estimation strategy and interprets our empirical results. Section 4 concludes.

2 A sticky-price New-Keynesian model with money

We sketch the log-linearized monetary policy model with money recently proposed by Andrés, López-Salido, and Nelson (2009) - the reader may refer to their paper for a detailed derivation.

2.1 Model's description

The economy consists of a representative household, a continuum of producing firms, and a monetary authority. Firms' problem is symmetrical, which allows us to focus on the behavior of a representative goods-producing firm.

Households

Each period households have an initial endowment of nominal money holdings M_{t-1} and risk-free bonds B_{t-1} - whose steady-state gross rate is $\overline{R}$ and net rate is r_t, receive a lump-sum nominal transfer T_t, labor income $W_t N_t$ - where N_t is the amount of supplied labor, and a nominal dividend from the firms operating in the economic system D_t. Households choose the sequences $\{C_{t+i,} M_{t+i}, B_{t+i}, N_{t+i}\}_{i=0}^{\infty}$ to maximize the discounted stream of utility

$$E_0 \sum_{t=0}^{\infty} \beta^t \left[a_t U(\widetilde{C}_t, \frac{M_t}{e_t P_t}) - \frac{N_t^{1+\varphi}}{1+\varphi} \right] - G\left(\bullet\right).$$

The utility is maximized subject to the contemporaneous budget constraint

$$\frac{M_{t-1} + B_{t-1} + W_t N_t + T_t + D_t}{P_t} = C_t + \frac{B_t / r_t + M_t}{P_t},$$

where $\widetilde{C}_t \equiv C_t / C_{t-1}^h$ is aggregate consumption (of different goods's quantities) adjusted for habit formation, β is the discount factor, h is the parameter identifying habits in consumption, a_t is the preference shock, P_t is the price aggregator, e_t represents the money velocity shock, and φ is the inverse of the Frisch labor supply elasticity.

The cost function $G\left(\bullet\right)$ takes the form

$$G\left(\bullet\right) = \frac{d}{2} \left\{ \exp\left[c \left(\frac{M_t / P_t}{M_{t-1} / P_{t-1}} - 1 \right) \right] + \exp\left[-c \left(\frac{M_t / P_t}{M_{t-1} / P_{t-1}} - 1 \right) \right] - 2 \right\},$$

where $c, d > 0$ regulates the portfolio adjustment costs. This specification is able to induce important effects on the money demand equation while maintaining the size of direct costs quite small. In particular (as discussed below), positive portfolio adjustment costs render the money demand equation dynamic and forward looking, so creating a link between current real balances and future, expected natural rates, possibly interpretable as long-term rates (Nelson (2002)). Importantly, this holds true also under nonseparability.[4]

Firms

The supply side of the economy is composed by several monopolistically competitive firms. Each firm j produces resources according to the following function: $Y_t^j = z_t N_t^{j(1-\alpha)}$, where Y_t^j is output, N_t^j represents hours hired from the household (such that $\int_0^1 N_t^j dj = N_t$), z_t is a common supply shock and $(1-\alpha)$ is the elasticity of labor with respect to output. The market clearing condition implies $\left(\int_0^1 Y_t^{j\frac{\varepsilon-1}{\varepsilon}} dj\right)^{\frac{\varepsilon}{1-\varepsilon}} = Y_t = C_t$, where ε is the elasticity of substitution between types of goods, and $\frac{\varepsilon-1}{\varepsilon}$ is the steady-state price mark-up. Prices are re-set according to a Calvo-type lottery, i.e. each period each firm has got a probability $(1-\theta)$ of reoptimizing its price. Firms who do not re-optimize simply adjust their prices at the pace of steady-state inflation $\overline{\pi}$. Moreover, a share $(1-\omega)$ of reoptimizing firms set prices optimally, while a fraction ω sets prices according to a rule of thumb, i.e. the new price is set according to the one-period lagged inflation rate.[5]

Equilibrium

As shown by Andrés, López-Salido, and Nelson (2009), after log-linearization of the symmetric equilibrium conditions around the steady-state values the economy's

[4] As pointed out by Nelson (2002) and Andrés, López-Salido, and Nelson (2009), a forward-looking money demand term would appear also if we modeled portfolio adjustment costs in terms of nominal balances. But real balances, besides offering algebraic convenience, captures the notion that portfolio adjustment costs are not literally transaction costs, but instead capture the convenience of maintaining, *ceteris paribus*, some purchasing power in the form of money - e.g. as a 'reserve against contingencies'.

[5] We follow Andrés, López-Salido, and Vallés (2006) and allow rule-of-thumb firms to re-optimize by indexing their price level to past inflation. Differently, Andrés, López-Salido, and Nelson (2009) allow non-reoptimizing firms to implement such indexation.

representation is the following:

$$\begin{aligned}
\widehat{y}_t \;=\;& \frac{\phi_1}{\phi_1+\phi_2}\widehat{y}_{t-1} + \frac{\beta\phi_1+\phi_2}{\phi_1+\phi_2}E_t\widehat{y}_{t+1} - \frac{1}{\phi_1+\phi_2}(\widehat{r}_t - E_t\widehat{\pi}_{t+1}) \\
& -\frac{\beta\phi_1}{\phi_1+\phi_2}E_t\widehat{y}_{t+2} + \frac{\psi_2}{\psi_1(1-\beta h)(\phi_1+\phi_2)}(\widehat{m}_t - \widehat{e}_t) \\
& -\frac{\psi_2(1+\beta h)}{\psi_1(1-\beta h)(\phi_1+\phi_2)}E_t(\widehat{m}_{t+1} - \widehat{e}_{t+1}) \\
& +\frac{\psi_2\beta h}{\psi_1(1-\beta h)(\phi_1+\phi_2)}E_t(\widehat{m}_{t+2} - \widehat{e}_{t+2}) + \frac{(1-\beta h\rho_a)(1-\rho_a)}{(1-\beta h)(\phi_1+\phi_2)}\widehat{a}_t,
\end{aligned} \tag{1}$$

$$\widehat{\pi}_t = \gamma_f E_t\widehat{\pi}_{t+1} + \gamma_f\widehat{\pi}_{t-1} + \lambda\widehat{mc}_t, \tag{2}$$

$$\begin{aligned}
\widehat{mc}_t \;=\;& (\chi+\phi_2)\widehat{y}_t - \phi_1\widehat{y}_{t-1} - \beta\phi_1 E_t\widehat{y}_{t+1} - \frac{\psi_2}{\psi_1(1-\beta h)}(\widehat{m}_t - \widehat{e}_t) \\
& +\frac{\psi_2\beta h}{\psi_1(1-\beta h)}E_t(\widehat{m}_{t+1} - \widehat{e}_{t+1}) - \frac{\beta h(1-\rho_a)}{(1-\beta h)}\widehat{a}_t - (1+\chi)z_t
\end{aligned} \tag{3}$$

$$\begin{aligned}
(1+\delta_0(1+\beta))\widehat{m}_t \;=\;& \gamma_1\widehat{y}_t - \gamma_2\widehat{r}_t + [\gamma_2(\overline{r}-1)(h\phi_2-\phi_1) - h\gamma_1]\widehat{y}_{t-1} - [\gamma_2(\overline{r}-1)\beta\phi_1]E_t\widehat{y}_{t+1} \\
& +\delta_0\widehat{m}_{t-1} + \left[\frac{\psi_2(\overline{r}-1)\beta h\gamma_2}{\psi_1(1-\beta h)} + \delta_0\beta\right]E_t\widehat{m}_{t+1} - \frac{(\overline{r}-1)\beta h(1-\rho_a)}{(1-\beta h)}\gamma_2\widehat{a}_t \\
& +\left\{1 - (\overline{r}-1)\gamma_2\left[\frac{\psi_2\beta h\rho_e}{\psi_1(1-\beta h)} + 1\right]\right\}\widehat{e}_t,
\end{aligned} \tag{4}$$

where $\widehat{x}_t \equiv \log(X_t/\overline{X})$ identifies a variable in log-deviation with respect to its steady-state value, and the following convolutions hold:[6]

[6] In fact, ψ_1, ψ_2, γ_1, and γ_2 are also convolutions of deep parameters. However, one would need to assume the exact form of the non-separability between consumption and real balances to pin down ψ_1 and ψ_2, a step that might bias our estimates in case of wrong specification of the utility function. Moreover, γ_1 and γ_2 have a clear interpretation as elasticity and semi-elasticity of money demand to real GDP and the nominal interest rate. Following Ireland (2001) and (2004), Andrés et al (2006), and Andrés et al (2008), we treat ψ_1, ψ_2, γ_1, and γ_2 as free parameters.

$$
\begin{aligned}
\psi_1 &\equiv -\frac{U_1}{\overline{y}^{(1-h)}U_{11}} \qquad\qquad \psi_2 \equiv -\frac{U_{12}}{\overline{y}^{(1-h)}U_{11}}\left(\frac{m}{e}\right) \\
\gamma_f &\equiv \beta\theta\left\{\theta+\omega[1-\theta(1-\beta)]\right\}^{-1} \quad \gamma_b \equiv \omega\left\{\theta+\omega[1-\theta(1-\beta)]\right\}^{-1} \\
\lambda &\equiv (1-\theta)(1-\beta\theta)(1-\omega)\xi \qquad \chi \equiv \frac{\varphi+\alpha}{1-\alpha} \\
\xi &\equiv \frac{(1-\alpha)}{1+\alpha(\varepsilon-1)}\left\{\theta+\omega[1-\theta(1-\beta)]\right\}^{-1} \\
\phi_1 &\equiv \frac{(\psi_1^{-1}-1)h}{1-\beta h} \quad \phi_2 \equiv \frac{\psi_1^{-1}+(\psi_1^{-1}-1)\beta h^2-\beta h}{1-\beta h} \\
\delta_0 &\equiv -\frac{c^2 d}{U_{22}m^2}.
\end{aligned}
$$

Eq. (1) is the Euler equation for consumption obtained with the imposition of the aggregate resource constraint. It displays leads and lags of real GDP because of households' rational expectations and habit formation. Notably, in case of nonseparability - i.e. $\psi_2 \neq 0$, real balances enter the aggregate demand schedule both in current and expected terms due to their impact on consumption's marginal utility. The impact of real balances on output is magnified by habit formation in consumption due to the link between current real balances and lagged consumption. As anticipated, real balances enter the IS equation given their impact on households' intertemporal choices.

Eq. (2) is a Phillips curve (NKPC) enriched with real balances. Firms' marginal costs, the forcing variable capturing the demand push in the NKPC, are defined as in eq. (3). Again, the pressure exerted by real balances - operative only under $\psi_2 \neq 0$ - is magnified by habit formation. The presence of money in firms' marginal costs is due to the effect exerted by real balances on households' labor supply decisions and, consequently, on real wages. An alternative interpretation of money in the NKPC is the cost-channel (Ravenna and Walsh (2006)), with money acting as a proxy of banks' lending rate.

Importantly, the log-linearized first order conditions feature real balances in deviation with respect to the money demand shock $\widehat{e}_t$. When a money demand shock hits, real balances move according to the money demand equation (4), but the Fed may neutralize the effect exerted on the short-term policy rate by varying money supply to keep the federal funds rate target constant. Consequently, real balances may oscillate even without causing any movements of output and inflation (Reynard (2007)). Then, one has to take into account oscillations of real money *on top of* money demand shocks.

Eq. (4) is a dynamic money demand equation featuring the presence of output leads and lags as well as the contemporaneous opportunity cost of holding money and future expected real balances. The parameter γ_1 represents the money-income elasticity, while the parameter γ_2 stands for the money-interest rate semi-elasticity. Importantly, portfolio adjustment costs play a key-role here. In fact, the money demand equation remains dynamic even under separability - i.e. $\psi_2 = 0$ - as long as portfolio adjustment costs affect households' utility - i.e. $\delta_0 > 0$. In this case, money enters neither the IS nor the NKPC, and impulse responses of output and inflation to a money demand shock are flat (as long as $\rho_m = 0$). Crucially, however, real balances act as leading indicators of future oscillations of the natural real interest rate, i.e. there is a 'direct effect' of the stock of money as stressed by Andrés, López-Salido, and Nelson (2009).

Monetary policy authorities

We model policymakers' decisions with a (log-linearized) augmented Taylor rule

$$\widehat{r}_t = \rho_r \widehat{r}_{t-1} + (1-\rho_r)(\rho_y \widehat{y}_t + \rho_\pi \widehat{\pi}_t + \rho_\mu \widehat{\mu}_t) + \varepsilon_{r_t}, \tag{5}$$

where

$$\widehat{\mu}_t = \widehat{m}_t - \widehat{m}_{t-1} + \widehat{\pi}_t \tag{6}$$

is the nominal money growth rate. The remainder of the rule is standard, in that it postulates a systematic reaction to output ρ_y and inflation ρ_π in percentage deviations with respect to their trend and steady-state values. The Taylor rate is implemented with gradualism, captured by the parameter ρ_r. A similar rule has been estimated by Ireland (2001), Sims and Zha (2006), Canova and Menz (2009), and Andrés, López-Salido, and Nelson (2009).

We close the model with four stochastic processes, which identify respectively the structural shocks to households' preferences, money demand, technology, and monetary policy:

$$\widehat{a}_t = \rho_a \widehat{a}_{t-1} + \varepsilon_{a_t}, \varepsilon_{a_t} \sim N(0, \sigma_{\varepsilon_a}) \tag{7}$$

$$\widehat{e}_t = \rho_e \widehat{e}_{t-1} + \varepsilon_{e_t}, \varepsilon_{e_t} \sim N(0, \sigma_{\varepsilon_e}) \tag{8}$$

$$\widehat{z}_t = \rho_z \widehat{z}_{t-1} + \varepsilon_{z_t}, \varepsilon_{z_t} \sim N(0, \sigma_{\varepsilon_z}) \tag{9}$$

$$\varepsilon_{r_t} \sim N(0, \sigma_{\varepsilon_r}). \tag{10}$$

These shocks feature mutually independent, uncorrelated innovations.

2.2 Model estimation

We estimate the model (1)-(10) using U.S. quarterly data on real output, real money balances, inflation, and the short-term nominal interest rate spanning the sample 1966:I-2007:II.[7] Given the clear historical upward trend displayed by real per-capita output and money, and the change in trends experienced by inflation and the federal funds rate in the post-WWII sample, we treat such series (log-series as for real output and real money) by applying the Hodrick-Prescott filter (weight: 1,600). The reason of this choice is twofold. First, it extracts the cyclical component of the series at hand, which allows us to focus on the frequencies the new-Keynesian model is designed to

[7] The data set is the same as in Ireland (2004). Output is measured by real GDP, real balances are constructed by dividing the M2 money stock by the GDP deflator, inflation is the quarterly change of the GDP deflator, and the interest rate is measured by the federal funds rate (quarterly counterpart). All data but the interest rate are seasonally adjusted. Output and real balances are expressed in per-capita terms (computed by employing the civilian non-institutional population, over 16). We feed the measurement equation with carefully demeaned series. The source of the data is the Federal Reserve Bank of St. Louis' website.

replicate.[8] Second, it enables us to compare our results to the literature that worked with *detrended* series (Ireland (2004), Andrés, López-Salido, and Vallés (2006), Canova and Menz (2009), Andrés, López-Salido, and Nelson (2009)). Alternatively, one could write a model with a unit root in technology in order to implement a model-consistent detrending of the observables, which would be employed in growth rates (e.g. Smets and Wouters (2007), Justiniano and Primiceri (2008)). While being theoretically appealing, this approach would force output and money to display a common (possibly stochastic) growth rate, an assumption not necessarily squaring up with the data.[9] Moreover, it is unclear if low frequencies come from the technological process or, instead, by random walk-type of preferences (Chang, Doh, and Schorfheide (2006)). Our agnostic filtering naturally endows each detrended series with its own flexible trend. It is worth stressing that, given that we filter our series over the entire sample, breaks in the low frequency component of our observables are not responsible for the parameter instability we may find in our empirical exercise.

We conduct our econometric analysis as follows. As a benchmark exercise, we estimate the model over the whole 1966:I-2007:II with a fixed-coefficient strategy.[10] This allows us to have results comparable to those present in the previously mentioned contributions, which hinge upon the assumption of stability of the structural parameters. We then move to the investigation of the possible instabilities affecting this model's relationships by implementing a rolling-window approach *a la* Canova (2009). In particular, we start from the 1966:I-1982:IV window and estimate the model, then we move the first and last observation of the window by four years and repeat the estimation. We keep the size of the window fixed (at 16 years) to minimize the differences in the preci-

[8] Canova (1998) compares the business cycle properties of real GDP extracted with the Hodrick-Prescott vs. alternatives, and discusses them in depth.

[9] The annualized, percentualized growth rate of the real per capita GDP in our sample is 1.56 per cent, while that of money growth reads 1.20.

[10] Estimations performed over the sample 1966:I-2006:IV, which we consider when conducting our recursive estimates, deliver very similar results.

sion of our estimates due to the sample-size. Our last window reads 1990:I-2006:IV, i.e. we consider seven different windows, which allow us to assess seven different posterior densities for all the parameters of interest.

As anticipated in the Introduction, we estimate the model with Bayesian techniques. Some dogmatic priors are imposed on a subset of parameters. We set the discount factor β to 0.9925, corresponding to an annual steady-state real rate of approximately 3%, and we calibrate the gross steady-state quarterly nominal interest rate $\overline{R}$ to 1.0138. Both values are in line with Smets and Wouters' (2007) estimates. We also set the capital-output elasticity α to 1/3, a very standard value in the literature. The elasticity of substitution between goods ε is fixed to 6, which implies a price mark-up equal to 1.2 as in Christiano, Eichenbaum, and Evans (2005).

We assume prior densities for the remaining 20 parameters. As previously stressed, ψ_2, ρ_μ, and δ_0 are key-parameters in this study. As far as nonseparability is concerned, we assume $\psi_2 \sim N(0, 0.5)$, i.e. a zero-mean, symmetric distribution.[11] The prior mean is centered to the value obtained by Andrés, López-Salido, and Vallés (2006) and Andrés, López-Salido, and Nelson (2009), and it lies between the maximum likelihood point estimate by Ireland (2001) - i.e. -0.0199 - and his calibration of the same parameter - i.e. 0.25. As for the Federal Reserve Bank's reaction to nominal money growth fluctuations, we assume $\rho_\mu \sim Gamma(0.8, 0.4)$, a diffuse prior centered at the point estimate obtained by Ireland (2001) and statistically in line with that proposed by Andrés, López-Salido, and Nelson (2009). Notice that we do not discard *a-priori* the scenarios featuring $\psi_2 = 0$ (separable utility function) and/or $\rho_\mu = 0$ (no reaction of the Fed to fluctuations in the money growth rate). In terms of portfolio adjustment costs, we assume $\delta_0 \sim Gamma(6, 2.85)$, i.e. a prior whose mean is very close to the point estimate by Andrés, López-Salido, and Nelson (2009), and whose variance is large

[11] For each assumed density (i.e. also for Beta and Gamma distributions), we indicate mean and standard deviation in brackets.

enough to lead the data to 'reject' the relevance of adjustment costs if that is the case. For the parameter ψ_1, which regulates the impact of money on inflation and output in case of nonseparability, we assume a $Gamma(0.8, 0.1)$, which is consistent with the calibration by Ireland (2004). As regards money demand elasticities, we assume $\gamma_1 \sim Gamma(0.5, 0.25)$ (elasticity to output) and $\gamma_2 \sim Gamma(0.2, 0.15)$ (semi-elasticity to the nominal interest rate), so lining up with the estimates proposed by Ball (2001).[12]

The remaining priors are fairly standard. For the habit formation parameter, we assume $h \sim Beta(0.7, 0.10)$. The priors for the policy parameters read $\rho_r \sim Beta(0.5, 0.1)$, $\rho_y \sim Beta(0.15, 0.05)$, $\rho_\pi \sim Beta(1.5, 0.05)$. The Calvo parameter is assumed to be $\theta \sim Beta(0.65, 0.1)$, price indexation $\omega \sim Beta(0.5, 0.15)$, and the inverse of the labor elasticity $\varphi \sim Gamma(1, 0.25)$. All the roots of the autoregressive shocks are assumed to follow a $Beta(0.75, 0.1)$, while the standard deviations of the structural shocks are assumed to be $InverseGamma(0.01, 1.5)$. All our priors are collected in Table 1.

We estimate the posterior distribution of the model as follows. Given the vector of parameters $\xi = [\beta, \alpha, \overline{r}, \varsigma, \psi_1, \psi_2, h, \gamma_1, \gamma_2, \theta, \omega, \varphi, \rho_R, \delta_0, \rho_y, \rho_\pi, \rho_\mu, \rho_a, \rho_e, \rho_z, \sigma_a, \sigma_e, \sigma_z, \sigma_r]'$, endogenous variables $z_t = [\widehat{y}_t, \widehat{r}_t, \widehat{\pi}_t, \widehat{m}_t]'$, exogenous shocks $\eta_t = [\widehat{a}_t, \widehat{e}_t, \widehat{z}_t]'$, innovations $\varepsilon_t = [\varepsilon_{a_t}, \varepsilon_{e_t}, \varepsilon_{z_t}, \varepsilon_{r_t}]'$, and observable variables we aim at tracking $Y_t = [\widehat{y}_t^{obs}, \widehat{r}_t^{obs}, \widehat{\pi}_t^{obs}, \widehat{m}_t^{obs}]'$, we write the model in state space form, we relate the latent processes to the observable variables via the measurement equation, we employ the Kalman filter to evaluate the likelihood $L(\{Y_t\}_{t=1}^T \mid \xi)$, and we estimate the posterior distribution $p(\xi \mid \{Y_t\}_{t=1}^T)$, which is proportional to the product of the likelihood function $L(\{Y_t\}_{t=1}^T \mid \xi)$ and the priors $\Pi(\xi)$, by employing a standard random-walk Metropolis-Hastings algorithm. We add serially and mutually independent $InverseGamma(0.01, 1.5)$ distributed measurement errors to control for high-frequency oscillations in the data that the business cycle

[12] Given that we employ the quarterly (as opposed to annual, or annualized) short-term interest rate in our empirical analysis, we rescaled the estimated value of the semi-elasticity γ_2 obtained by Ball (2001) - i.e. 0.05 in absolute value - by a factor of 4.

model at hand might not be able to capture.[13]

3 Empirical findings

We first present the results stemming from our fixed-coefficient investigation. This exercise is conducted to get baseline results to perform comparisons with the existing literature. Then, we move to the rolling-window analysis, and concentrate on i) the evolution of the key-structural parameters of the model, and ii) the estimated, subsample specific impulse response functions of the macroeconomic aggregates to the four identified structural shocks.

3.1 Fixed coefficients ...

Table 1 displays the posterior median along with the $[5th, 95th]$ posterior percentiles of the estimated structural parameters. In so doing, we contrast the standard New-Keynesian model estimated under $\psi_2 = \delta_0 = \rho_\mu = 0$ - i.e. nonseparability, no direct effect, no policy reaction to monetary aggregates - to the model that allows, but does not necessarily require, money to shape the macro-dynamics of interest. Several results are

[13] To perform our Bayesian estimation we employed Dynare 4.0, an set of algorithms developed by Michel Juillard and collaborators. Dynare is freely available at http://www.cepremap.cnrs.fr/dynare/. The model is estimated by implementing a two-step strategy. First, we estimate the mode of the posterior distribution by maximizing the log-posterior density, which combines our priors on the parameters of interest with the likelihood function. Second, we employ the random-walk Metropolis-Hastings algorithm to estimate the posterior distribution. The mode of each parameter's posterior distribution was computed by using the 'csminwel' algorithm elaborated by Chris Sims. A check of the posterior mode, performed by plotting the posterior density for values around the mode for each estimated parameter in turn, confirmed the goodness of our optimizations. We then exploited such modes for initializing the random walk Metropolis-Hastings algorithm to simulate the posterior distributions. In particular, the inverse of the Hessian of the posterior distribution evaluated at the posterior mode was used to define the variance-covariance matrix of the chain. The initial VCV matrix of the forecast errors in the Kalman filter is set to be equal to the unconditional variance of the state variables. We initialized the state vector in the Kalman filter with steady-state values. We simulated two chains of 200,000 draws each, and discarded the first 50% as burn-in. To scale the variance-covariance matrix of the random walk chain we used factors implying an acceptance rate belonging to the [23%,40%] interval. We verified the convergence towards the target posterior distribution via the Brooks and Gelman (1998) convergence checks. As typically done in the literature, we discarded all the draws not implying a unique equilibrium of the system.

worth commenting. First and foremost, the marginal likelihood clearly favors the model with money, with a deterioration associated to the restricted framework of about 12 log-points, which translates in a Bayes factor equal to $exp(2615.1 - 2603.2) = 147,240$.[14] This is very strong evidence in favor of the model with money. Digging deeper, it turns out that the deterioration of the fit is mainly due to the restriction imposed on the portfolio adjustment cost parameter. In fact, under the restriction $\psi_2 = 0$ (only), the model's fit, in terms of Marginal Likelihood, increases to 2616.3. This may be explained by the negligible role played by nonseparability, which is 'rejected' by the automatic penalization of overparameterization embedded in the computation of the marginal likelihood. By contrast, when imposing $\delta_0 = 0$ (only), the model's fit dramatically drops to 2600.8, clearly 'rejecting' the imposition of no portfolio adjustment costs. These comparisons squares up with the posterior densities of the key parameters. The posterior median of ψ_2 is very small, i.e. 0.05, and its credible set clearly contains the zero value. The posterior of δ_0 reads 3.2. This values is slightly smaller than the point estimate proposed by Andrés, López-Salido, and Nelson (2009), but it is statistically in line with it. As for the reaction of the Fed to money, the posterior median reads 0.10, a value lower than that found in previous contributions. Indeed, in this last case, the marginal likelihood favors the restricted model with a standard Taylor rule displaying no monetary aggregates *a la* Ireland (2004), with a value equal to 2621.9.

As regards other money-related parameters, ψ_1, which affects the impact of money on output and inflation, has an estimated posterior distribution equal to 0.69, which resembles the estimates by Andrés, López-Salido, and Nelson (2009), and with a $95th$ percentile close to unity, which is the calibration proposed by Ireland (2004). The posterior median of the money-output elasticity is 0.88, slightly lower than the point-

[14] We compute the marginal likelihood *via* the modified harmonic mean estimator by Geweke (1998). In computing model comparisons via the Bayes factor, we keep the priors on the common parameters fixed across models, as done by e.g. Rabanal and Rubio-Ramírez (2005), Rabanal (2007), and Canova (2009). For a different strategy, see Del Negro and Schorfheide (2008).

estimate provided by Ball (2001).[15] As regards the money-interest rate semi-elasticity, our estimated figure, normalized in order to account for the quarterly (vs. annualized) nominal interest rate, amounts to about 0.35, larger than the point estimate provided by Ball (2001) but statistically in line with the one by Andrés, López-Salido, and Nelson (2009).

The posterior distributions of the remaining parameters propose values typically found in the literature. In particular, the posterior median of the habit formation parameter reads 0.86, a value close to those in Rabanal (2007), Christiano, Eichenbaum, and Evans (2005), and Smets and Wouters (2007). The median of the Calvo parameter is 0.66, quite a standard figure in the macroeconomic literature. Also the inverse of the Frisch labor elasticity assumes the conventional value of 1. Taylor rule coefficients suggest an aggressive, gradually implemented long-run reaction of the Fed to inflation swings, in line with some previous literature (Clarida, Gali, and Gertler (2000)), at least as regards the post-1982 sample. Interestingly, the autoregressive parameters of the structural shocks are all below 0.9, which suggests that the model features an internal propagation mechanism able to capture the persistence of the observed macroeconomic series.

Wrapping up, our full sample *fixed coefficients* estimates i) offer clear statistical support to the role of portfolio adjustment costs, ii) reject the relevance of nonseparability, and iii) cast doubts on the role played by monetary aggregates in the post-WWII U.S. monetary policy conduct.

3.2 ... vs. recursive estimates

Intriguingly, a quite richer picture arises when relaxing the conventional fixed coefficient assumption. Figure 1 displays the evolution of (selected) structural parameters

[15] In making this comparison one should take into account the fact that our model is estimated with a *detrended* measure of output, as opposed to the undetrended log-output measure Ball (2001) focuses on.

constructed by considering seven different (partly overlapping) windows. Top-row parameters are those characterizing money's role in the estimated model. First of all, differently with respect to the indications coming from the full sample estimates, nonseparability (namely, complementarity) is clearly supported in subsamples heavily influenced by the 1970s. Focusing on the first window as the reference for the 1970s, it is interesting to note that the (log) marginal likelihood of the unrestricted model, which reads 936.1, drops (moderately) when forcing separability between consumption and money to take place (934.2), remarkably deteriorates when assuming no adjustment costs (925.0), and collapses to 921.6 in correspondence to the standard, 'cashless' new-Keynesian framework. This signals that the impact of monetary aggregates is pervasive when conditioning to the great inflation observations, a result in line with Canova and Menz's (2009), at least as regards nonseparability and policy effects. A quite different picture emerges when conditioning to the last window, which we take as representative of the dynamics during the great moderation. The estimated median of the nonseparability parameter reads 0.13, a quite smaller value than 0.62, i.e. that of the first window. The posterior median of the adjustment costs moves from 1.98 to 4.00, but the uncertainty surrounding it is very large. Also the systematic reaction of the Fed to money growth declines from 0.61 to 0.26, signalling a lower attention to monetary aggregates. Overall, the restricted model performs better in the last window, with a (log) marginal likelihood reading 1081.6 vs. 1080.0 (the latter being that of the money-endowed model).

Other parameters also display a significant evolution over time. In particular, the money demand elasticity to output displays a clear downward trend. In contrast, the money-interest rate semi-elasticity is estimated to be fairly stable. Habits remarkably increase over time, a result that may signal breaks in preferences by American households and/or capture the effects of financial innovations, which have possibly rendered consumption smoothing less costly in the last 25 years. The Taylor rule parameters do not display much instability, a finding in line with Smets and Wouters (2007) and Justiniano and Primiceri (2008). As for shocks' volatilities, we record a non-monotonic pattern for preference shocks, which contrasts the somewhat declining path followed by both policy rate and technological shocks, and the more stable evolution of money demand innovations.

While assuming a-priori independence among parameters' densities, ex-post correlation is typically the case when conducting Bayesian estimations. Our exercises represent no exceptions. When comparing sets of common coefficients under two versions of the model, i.e. unrestricted with money vs. restricted, standard new-Keynesian without money, interesting findings arise. Figure 2 shows how money may be of help for spotting instabilities of some parameters that would not otherwise arise. In particular, when estimating a money demand having no feedbacks on the remaining part of the system, one finds a quite stable elasticity to output. Also the degrees of habit formation appears to be constant if money is omitted from the model. As regards the parameters of the Taylor rule, one may notice some mild differences between the two scenarios but, given the large uncertainty surrounding the estimated Taylor parameters, such differences are hardly meaningful from an economic standpoint. This might be due to our decision to discard draws leading to multiple equilibria, a choice widely adopted in this empirical literature.[16] Interestingly, the absence of money induces a monotonic decline in the preference shock's volatility, which features instead an inverted U-shape when money is allowed to enter the picture.

To summarize, the interactions between money and the remaining aggregates strongly influence the evolution of some key-structural parameters. However, this mainly occurs when observations coming from the 1970s are dominant in the windows considered in our

[16] Notice that, following most of the literature, we do not allow for multiple equilibria to arise. For two notable exceptions, see Lubik and Schorfheide (2004) and Boivin and Giannoni (2006).

analysis. Indeed, for our last window, i.e. 1990:I-2006:IV, differences in the estimated parameters appear, if present, negligible. One may then wonder if the instability of the estimated parameters we found is reflected in the model-consistent impulse response functions.

3.3 Impulse response function analysis

We stick to the comparison involving the first and last windows, and we plot the estimated impulse responses of the benchmark vs. money-endowed frameworks. Indeed, time-dependent parameters imply window-specific impulse responses. The responses associated to the first window 1960:I-1982:IV are depicted in Figure 3. Evidently, the omission of money may indeed bias in an economically relevant manner the estimated responses. In terms of magnitude, the model without money clearly dampens the effects of a monetary policy shock to output, inflation, and real balances, of the preference shock to inflation and the policy rate, and of the technological shock to all our endogenous variables. Moreover, money demand shocks, which have (by construction) zero effects on all variables (except money) in the restricted model, are estimated to induce quantitatively important reactions by the modeled variables, output *in primis*.

This picture dramatically changes when moving to the sample 1990:I-2006:IV, whose estimated responses are depicted in Figure 4. Evidently, the role of money appears to be much milder, if not absent at all. Moreover, the effects of money demand shocks are also moderate. A change in the transmission mechanism of all structural shocks is likely to have occurred, with money losing much of its influence on U.S. output and inflation. However, also in the last window one may appreciate the reaction of output to money demand shocks. This suggests that money may still be important in empirical analysis conducted over the great moderation sample, possibly to control for omitted information-induced biases otherwise affecting the structural parameters of the

Euler-equation for output (Hafer, Haslag, and Jones (2007)).

4 Conclusions

We estimate a new-Keynesian model of the business cycle to assess the possibly time-varying role played by money in the post-WWII U.S. sample. We conduct Bayesian estimations with the rolling-window approach recently put forward by Canova (2009), a strategy suited to deal with the (possible) instability of the structural parameters of the DSGE model at hand. Our results reveal that money is actually a relevant aggregate to understand the U.S. output and inflation dynamics in the 1970s. In particular, impulse responses reveal a strong effect exerted by money demand shocks on inflation and output. Moreover, the presence of money clearly affects the estimated dynamics in response to other identified structural innovations. In contrast, money turns out to play a much more moderate role for the description of the great moderation dynamics. Our results suggest that a fixed-coefficient analysis may severely bias the time-dependent role that money has played in shaping the U.S. macroeconomic dynamics for the last four decades.[17]

Our results rely upon the investigation of an extended version of the model proposed by Ireland (2004). Indeed, it may very well be that the stock of money in this model serves as a proxy for financial wealth, an hypothesis empirically validated by Canova and Menz (2009). Christiano, Motto, and Rostagno (2007) and Castelnuovo and Nisticò (2009) offer fresh evidence on the relevance of stock prices for the U.S. economy. Therefore, while giving money a chance to play an active role in the determination of inflation and the business cycle, current monetary models do not explicitly embed ingredients such as asymmetric information in the lending market, imperfect substitutability between financial assets, and so on. We then agree with Nelson (2008), who calls for a

[17] For a recent application of this reasoning to the validity of the expectations hypothesis for the U.K. term structure of interest rates, see Bianchi, Mumtaz, and Surico (2009).

new generation of models based on a more satisfactory microfoundation of the role of money. In the light of the current liquidity boom triggered by a variety of central banks to tackle the real effects of the financial turmoil, this call appears to be warranted.

From an empirical standpoint, our analysis has (necessarily) dealt with the identification of the cyclical components of the aggregates under investigation. This identification issue is crucial in empirical work, and an econometrician's choices along this dimension may very well be as important as those of the modelers building up the framework to be taken to the data. In a recent paper, Canova and Ferroni (2009) show how to tackle this 'filtering uncertainty' by dealing with a variety of 'contaminated proxies' of output and money. Interestingly, their application to actual data shows exactly that the role of money may turn out to be downplayed by the choice of the 'wrong' filter. We see Canova and Ferroni's (2009) methodology as very promising to detect the role of money in monetary models of the business cycle. About frequency-decompositions, more attention should be paid on the possible link between systematic policy drifts and the money-inflation low-frequency relationship as dictated by, say, the quantity theory (Sargent and Surico (2009)). Indeed, we see the improvement of our current understanding of money's role in monetary business cycle models as an exciting area for future research.

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Param.	*PriorDistrib.*	*Prior Mean* (*St.dev.*)	NK model with money *Posterior Median* [*5th, 95th*]	Baseline NK model *Posterior Median* [*5th, 95th*]
ψ_1	Gamma	0.80 (0.10)	0.69 [0.58,0.80]	0.68 [0.57,0.79]
ψ_2	Normal	0.00 (0.50)	0.05 [−0.03,0.19]	–
h	Beta	0.70 (0.10)	0.86 [0.76,0.95]	0.88 [0.78,0.96]
θ	Beta	0.65 (0.10)	0.66 [0.53,0.78]	0.67 [0.54,0.79]
ω	Beta	0.50 (0.15)	0.76 [0.65,0.86]	0.77 [0.66,0.86]
φ	Gamma	1.00 (0.25)	0.95 [0.55,1.33]	0.94 [0.59,1.34]
γ_1	Gamma	0.50 (0.25)	0.88 [0.29,1.51]	0.39 [0.22,0.59]
γ_2	Gamma	0.20 (0.15)	0.35 [0.02,0.86]	0.37 [0.04,0.73]
δ_0	Gamma	6.00 (2.85)	3.20 [1.22,5.61]	–
ρ_R	Beta	0.50 (0.10)	0.44 [0.32,0.56]	0.40 [0.27,0.52]
ρ_y	Gamma	0.15 (0.05)	0.13 [0.08,0.18]	0.11 [0.07,0.16]
ρ_π	Gamma	1.50 (0.25)	1.67 [1.38,1.96]	1.63 [1.36,1.92]
ρ_μ	Gamma	0.80 (0.40)	0.10 [0.03,1.18]	–
ρ_a	Beta	0.75 (0.10)	0.74 [0.65,0.83]	0.75 [0.66,0.84]
ρ_e	Beta	0.75 (0.10)	0.79 [0.71,0.88]	0.88 [0.83,0.93]
ρ_z	Beta	0.75 (0.10)	0.71 [0.57,0.84]	0.72 [0.59,0.84]
σ_a	Inv_Gamma	0.01 (1.5)	0.0105 [0.0062,0.0152]	0.0104 [0.0062,0.0153]
σ_e	Inv_Gamma	0.01 (1.5)	0.0175 [0.0115,0.0250]	0.0077 [0.0049,0.085]
σ_z	Inv_Gamma	0.01 (1.5)	0.0091 [0.0050,0.0146]	0.0081 [0.0064,0.0131]
σ_r	Inv_Gamma	0.01 (1.5)	0.0020 [0.0016,0.0024]	0.0019 [0.0016,0.0023]
Marg.Lik.			2615.1	2603.1

Table 1: **Model Comparison: Full Sample Estimates.** Sample: 1966:I-2007:II. The computation of the Marginal Likelihoods was performed by employing the Modified Harmonic Mean estimator proposed by Geweke (1998). The Table reports posterior medians and the [5th,95th] posterior percentiles. The posterior summary statistics are calculated from the output of the Metropolis algorithm. Details on estimation procedure are reported in the text.

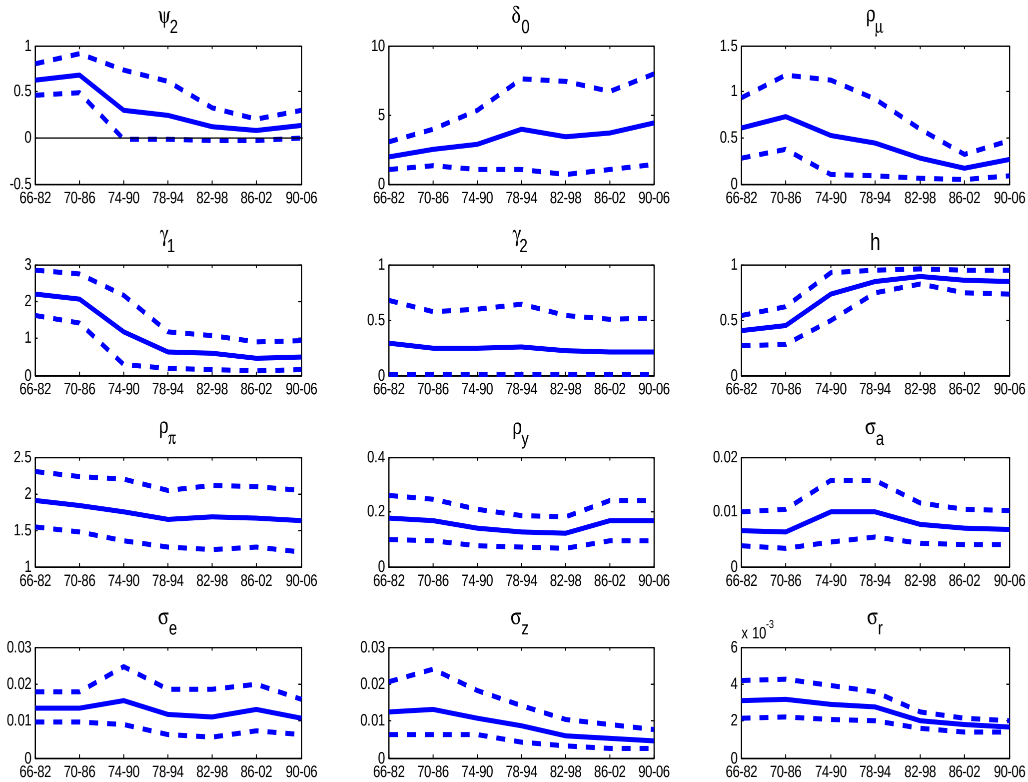

Figure 1: **Evolution of Structural Parameters over Time.** Solid line: Posterior medians. Dotted line: 5th and 95th posterior percentiles. Evolution of the parameters constructed by employing seven rolling windows of 16-year constant length. Windows: [1966:I-1982:IV, 1970:I-1986:IV, ..., 1990:I-2006:IV].

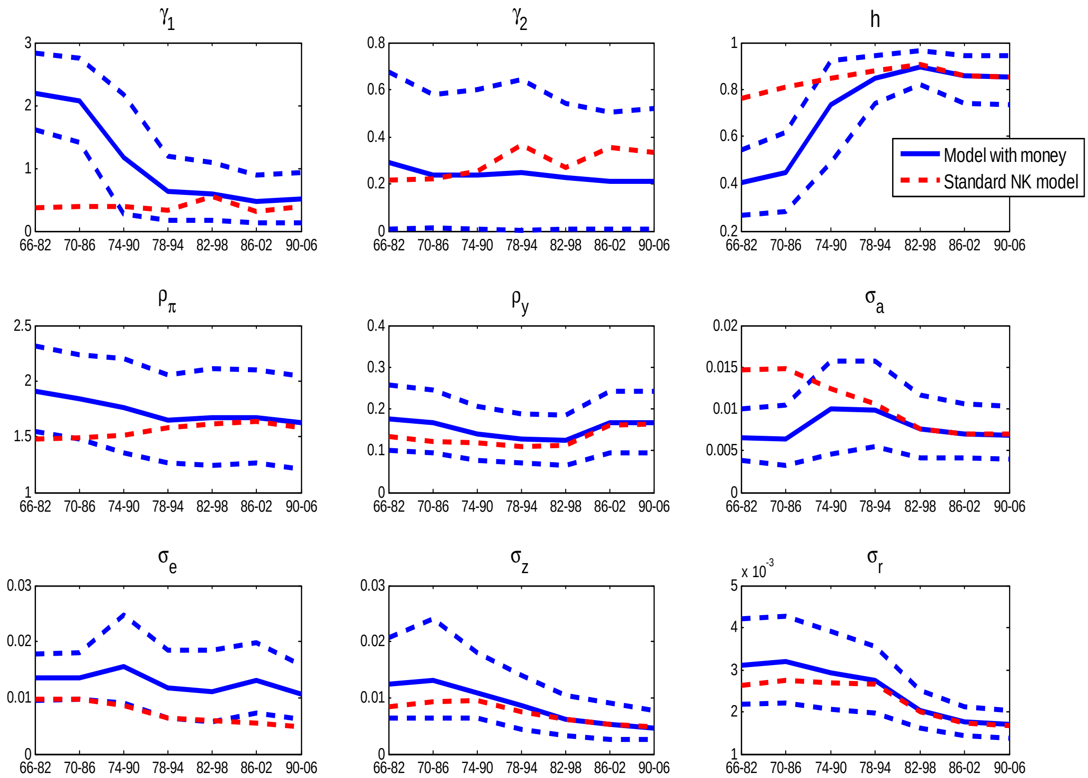

Figure 2: **Evolution of Structural Parameters over Time: Model Comparison.** Solid line: Posterior medians. Dotted line: 5th and 95th posterior percentiles. Evolution of the parameters constructed by employing seven rolling windows of 16-year constant length. Windows: [1966:I-1982:IV, 1970:I-1986:IV, ..., 1990:I-2006:IV].

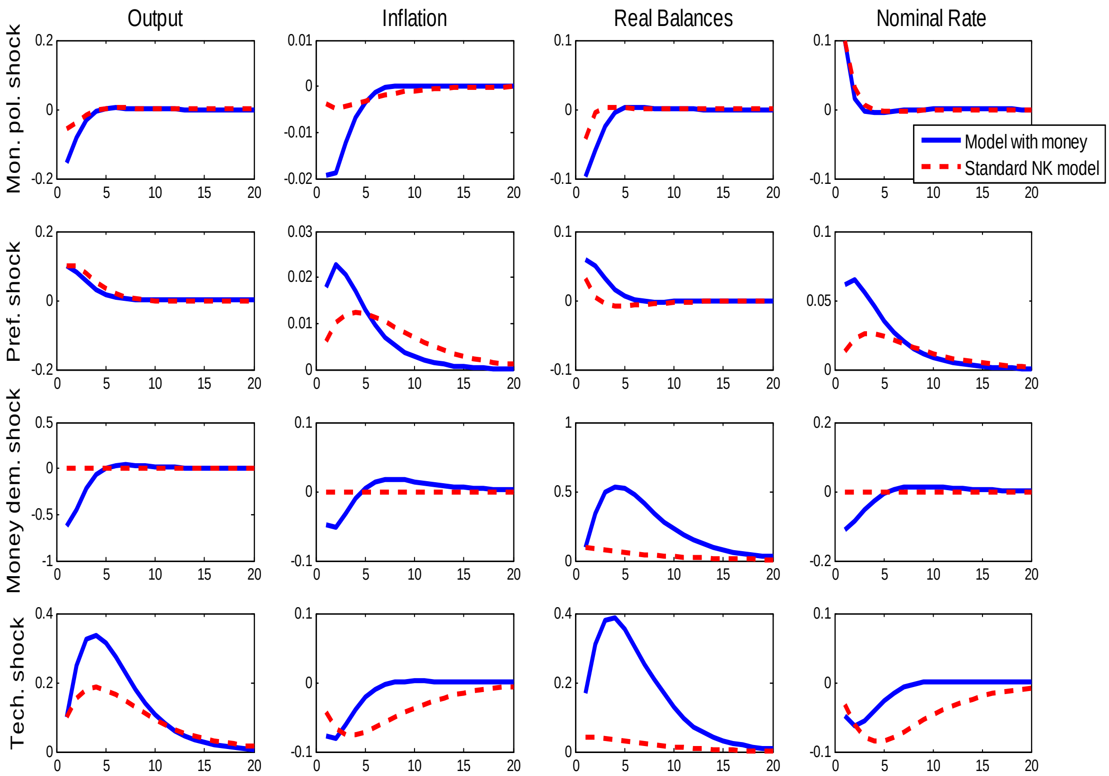

Figure 3: **Responses to Shocks: 1966:I-1982:IV.** Monetary policy / preference / money demand / technology shock normalized to induce a 0.1 impact reaction of, respectively, the nominal interest rate / output / real balances / output.

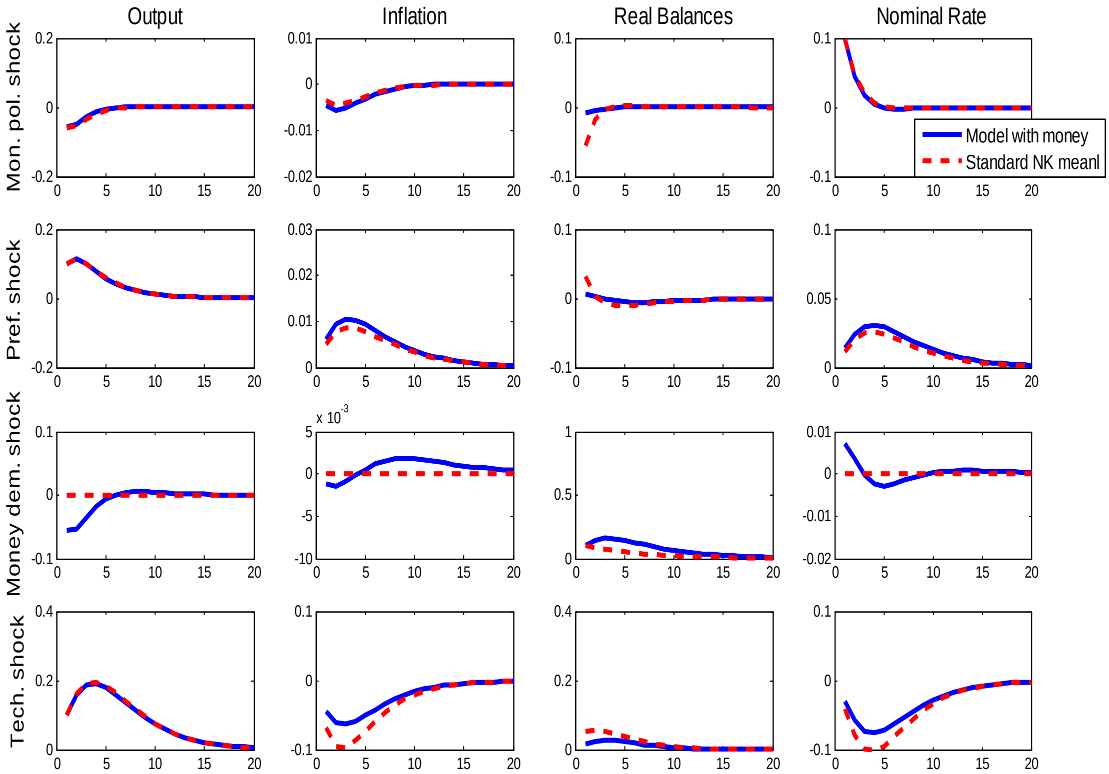

Figure 4: **Responses to Shocks: 1990:I-2006:IV.** Monetary policy / preference / money demand / technology shock normalized to induce a 0.1 impact reaction of, respectively, the nominal interest rate / output / real balances / output.