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Contents

Abstract

We investigate whether business cycle fluctuations affect the degree of excess sensitivity of private consumption growth to disposable income growth. Using multivariate state space methods and quarterly US data for the period 1965-2000 we find that excess sensitivity is significantly higher during recessions.

Key words: Private consumption, excess sensitivity, business cycle, liquidity constraints, state space models.

JEL classification: E21.

Summary

When consumers can freely lend and borrow on capital markets, aggregate private consumption should only react to changes in permanent income. Previous scientific work, however, finds that total consumption growth in the economy is determined by the growth rate in total disposable income. An important interpretation of this observation is that a fraction of the consumers in the economy is having a hard time obtaining credit. We say that these consumers are liquidity constrained. Therefore, when confronted with a higher income, these consumers tend to spend the additional amount instead of saving it. Another part of the consumer population does not face difficulties obtaining a loan and is therefore able to consume as much as it can. When confronted with a higher income these consumers do not necessarily consume the additional amount: they save it.

In this paper we investigate whether the impact of disposable income growth on consumption growth is higher during recessions than during expansions, ie whether during recessions there is a higher number of consumers who spend their disposable income. We find that this is the case. Our finding is based on a data set for the US economy that covers the period 1965-2000.

From a policy point of view, our findings suggest that the impact of policy changes that affect disposable income is very likely to have greater effects during recessions than during expansions. Our study is motivated by theoretical results found in previous work where it is argued that during recessions liquidity constraints faced by consumers are more severe than in expansions. The reason is that the worsening of households' balance sheets in a recession decreases the possibility of consumers financing their expenditures through accumulated wealth. This raises the demand for credit. At the same time however the higher monitoring and contract enforcement costs faced by banks during recessions increases the cost of banks to give loans and therefore diminishes the credit supply. Our observation that consumption growth depends more heavily on disposable income growth during recessions thus supports previous theoretical results.

In our study we revisit an issue that was investigated in previous studies, namely the possibility that, over time, the fraction of liquidity-constrained consumers has decreased. In previous work it has been suggested that financial liberalisation and the development of credit markets that has occurred in the United States (especially during the 1980s) may have reduced the numbers of

consumers that are liquidity constrained. We test this hypothesis by looking at whether the impact of disposable income growth on private consumption growth has fallen over the period 1965-2000. We find that it has not, suggesting that the average number of consumers that are liquidity constrained has not decreased.

1 Introduction

Under very strict assumptions, the permanent income hypothesis implies that aggregate private consumption follows a random walk (Hall (1978)). Maximising forward-looking consumers lend and borrow freely on perfect capital markets to smooth consumption over time. In reality, however, private consumption growth is found to be excessively sensitive to current disposable income growth. This observed excess sensitivity (ES) can be explained theoretically by dropping Hall's assumptions. The most common interpretation of the observed ES is the prevalence of liquidity constraints (Campbell and Mankiw (1991); Bacchetta and Gerlach (1997)). More recent evidence by Ludvigson (1999) and Sarantis and Stewart (2003) reinforces this conclusion. Some theoretical models predict a correlation between consumption growth and income growth when consumers are liquidity constrained (Deaton (1991); Ludvigson (1999)). The second most often mentioned explanation is precautionary savings (Zeldes (1989); Caballero (1990); Carroll (1992, 1994); Ludvigson and Michaelides (2001)). In particular 'buffer stock' models of saving (Carroll (1992)) predict that consumers attribute a large weight to current income in their consumption decisions. While there is no consensus in the literature on the reasons for the observed ES, the assumption that the ES parameter is constant has been abandoned in recent studies in favour of time-varying specifications (Campbell and Mankiw (1991); McKiernan (1996); Bacchetta and Gerlach (1997); Pozzi *et al* (2004)). In particular, the impact of long-run driving factors of ES such as financial liberalisation and the development of credit markets has been documented extensively in previous studies (Campbell and Mankiw (1990, 1991); Bacchetta and Gerlach (1997)).

In this paper we investigate the impact of business cycle fluctuations on the degree of excess sensitivity of private consumption growth to disposable income growth by using quarterly US data over the period 1965-2000. The contribution of the paper is both empirical and methodological.

Empirically, the paper focuses on short-run factors that potentially affect the degree of excess sensitivity instead of long-run factors. While the potential impact of the business cycle on the excess sensitivity parameter has been sporadically hinted at (see eg Campbell and Mankiw (1991)) no focused investigation of this issue has yet been conducted. This is somewhat surprising since, from a theoretical perspective, both the liquidity constraints and the precautionary savings interpretation of ES can rationalise a role for the business cycle. With respect to liquidity constraints, there is a literature that suggests that liquidity constraints are more severe in recessions than in booms (see eg Stiglitz and Weiss (1981); Bernanke and Gertler (1989)).[1] The deterioration of households' balance sheets in a recession decreases internal financing possibilities (ie through income or accumulated wealth) thereby raising the demand for external finance. Higher monitoring and contract enforcement costs and information asymmetries may increase the risk for banks of giving loans in recessions and diminish the supply of credit. These factors may lead to a higher 'external finance premium', ie the difference between the cost of external and internal finance. As noted by Jappelli and Pagano (1989) a high 'external finance premium' may be the source of liquidity constraints and excess sensitivity.[2] With respect to precaution, Carroll (1992) emphasises that spells of unemployment may be the most important source of income uncertainty. If, as predicted by buffer stock models of consumption, uncertainty and precaution induce a correlation between consumption and current income growth then spells of unemployment occurring during recessions may reinforce this correlation.

Methodologically, we use state space methods to simultaneously estimate a consumption growth equation and a multivariate stochastic process for the ES parameter. This approach is different from the methods applied until now where, if a multivariate process for the ES parameter is considered, either a two-step approach is used (McKiernan (1996)) or the process for the ES parameter is, rather restrictively, assumed to be a deterministic function of the variables considered (see eg Evans and Karras (1998); Sarantis and Stewart (2003); Pozzi *et al* (2004)).

Our results suggest that ES is positively affected by the change in the unemployment rate, ie ES is significantly higher during recessions. This result can be reconciled with both the liquidity constraints and the precautionary savings interpretation of ES. We do not find a significant impact on ES of low frequency controls however as we find a negative but insignificant impact of both a dummy that allows for a different average ES parameter in the post-1982 period and a linear time trend. From a policy point of view, our findings suggest that the impact of policy changes that affect disposable income is very likely to have greater effects during recessions than during expansions.

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(1) So far, there is only empirical evidence on liquidity constraints and the business cycle for firms, not households. Gertler and Gilchrist (1994), Vermeulen (2002) and Peersman and Smets (2005) find that small firms are more liquidity constrained during downturns.

(2) Note that the possibility of a positive external finance premium (eg a wedge between lending and deposit rates) is a deviation from the standard permanent income hypothesis. The latter theorem is based on the assumption that the same interest rate applies to both lenders and borrowers.

The paper is structured as follows. In Section 2 we present the theoretical framework. In Section 3 we present the empirical specification and we discuss the estimation methodology. Section 4 presents the estimation results while Section 5 concludes.

2 Theoretical framework

Suppose a representative consumer maximises expected utility by choosing a consumption path over an infinite lifetime. If the instantaneous utility function of this consumer is of the constant relative risk aversion type, if the consumer lends and borrows against the same constant interest rate, and if the growth rate of private consumption is normally distributed, then we can write the first-order condition for this consumer as,

$$\Delta c_t = \alpha_t + \varepsilon_t \qquad \textbf{(1)}$$

where Δc_t is the growth rate of real per capita consumption, where α_t encompasses the difference between the interest rate and the rate of time preference and the conditional variance of consumption growth, and where ε_t is an innovation that is uncorrelated with lagged variables (for the derivation, see Appendix A).

A large literature has demonstrated that private consumption growth is typically excessively sensitive to the growth rate in disposable income (Campbell and Mankiw (1990, 1991)). Thus reality may be better approximated by

$$\Delta c_t = \alpha_t + \beta_t \Delta y_t + \varepsilon_t \qquad \textbf{(2)}$$

where Δy_t is the growth rate of real per capita disposable income, and β_t is the excess sensitivity parameter ($0 \leq \beta_t \leq 1$). The most common interpretation for $\beta_t > 0$ is that the representative agent solution does not hold because of liquidity constraints (see Campbell and Mankiw (1991); Bacchetta and Gerlach (1997)). The second most often mentioned explanation is precautionary savings (Zeldes (1989); Caballero (1990); Carroll (1992, 1994); Ludvigson and Michaelides (2001)). Liquidity constraints and precaution are the two explanations that we emphasise in the paper.[3] Note that while the early literature on excess sensitivity assumes a constant excess sensitivity parameter, we follow the approach undertaken in more recent studies which is to consider a time-varying degree of excess sensitivity (see Campbell and Mankiw (1991);

(3) Other explanations are myopia (see Flavin (1985) who dismisses this explanation in favour of a liquidity-constraints explanation) and imperfect information (see Pischke (1995)). Contrary to the liquidity-constraints and precaution hypotheses the latter two explanations offer no rationale however of why business cycle fluctuations would have an impact on excess sensitivity and therefore are less relevant in the present context.

McKiernan (1996); Bacchetta and Gerlach (1997); Pozzi *et al* (2004)). In particular, besides allowing only for low frequency movements in β_t as in Campbell and Mankiw (1991) and Bacchetta and Gerlach (1997) – they attribute low frequency time-variation in β_t to the development of credit markets and financial liberalisation – we also investigate the impact of business cycle fluctuations on β_t. We discuss our empirical specification for β_t in the next section.

3 Empirical specification and estimation methodology

3.1 Empirical specification

We consider the following empirical specification,

$$\Delta c_t = \alpha_t + \beta_t \Delta y_t + \varepsilon_t + \theta \varepsilon_{t-1} \quad \textbf{(3)}$$

$$\alpha_t = \alpha_{t-1} + \varepsilon_t^{\alpha} \quad \textbf{(4)}$$

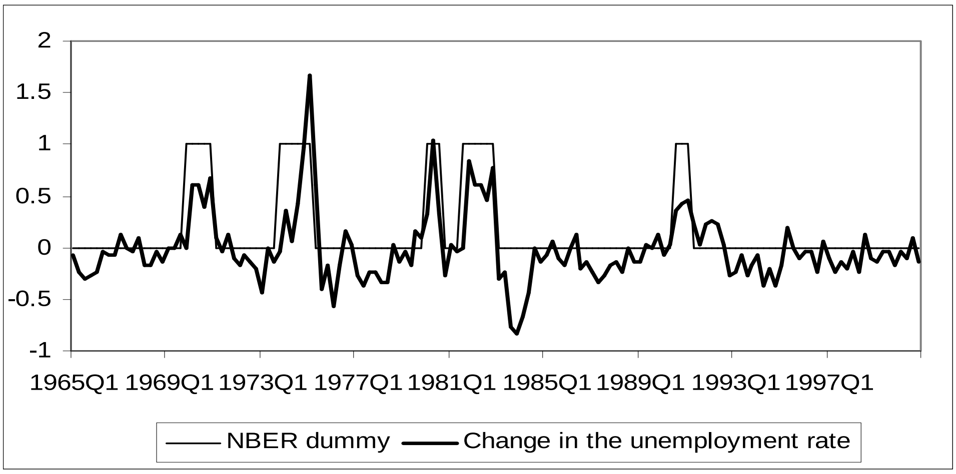
$$\beta_t = \beta_0 + \beta_1 lf_t + \beta_2 bc_t + \varepsilon_t^{\beta} \quad \textbf{(5)}$$

From equation **(3)** we note that the error term in consumption growth now has an $MA(1)$ structure where for the $MA(1)$ parameter θ we have $-1 \leq \theta \leq 1$. The reasons that we allow for an $MA(1)$ error in consumption growth are potential time aggregation (Working (1960)), problems related to the presence of durable components in our consumption measure (Mankiw (1982)), and potential transitory components in the log of consumption. Following Bacchetta and Gerlach (1997) we specify α_t as a random walk in equation **(4)**. Equation **(5)** is our specification for the time-varying excess sensitivity parameter β_t. We model β_t as a straightforward linear function of a low frequency control (lf_t), and a variable reflecting the state of the business cycle (bc_t). For lf_t we use both a linear time trend and a dummy variable that takes on the value 0 before 1982:01 and 1 from 1982:01 onward.[(4)] We proxy bc_t by the change in the unemployment rate Δu_t. As can be seen in Chart 1 this variable is highly correlated with the turning points of the business cycle as calculated by the National Bureau of Economic Research (ie the NBER recession dummy which takes on the value 1 in recessions).[(5)] Note, finally, that the error terms ε_t , ε_t^{α}, and ε_t^{β} are assumed to be independent Gaussian white noise terms (with variances σ_{ε}^2, $\sigma_{\varepsilon^{\alpha}}^2$, and $\sigma_{\varepsilon^{\beta}}^2$ respectively).

(4) This date is qualified by Kaminsky and Schmukler (2003) as the point in time where the domestic financial sector in the US can be considered 'fully liberalized' (to be interpreted as the date on which regulations like credit allocation control were fully lifted). However, it may also capture other events that may have had an impact on excess sensitivity, eg the Volcker disinflation.

(5) When estimating the system with the NBER recession dummy instead of the change in the unemployment rate we encountered numerical problems and our results were meaningless.

Chart 1: NBER turning points and the change in the unemployment rate (US data, 1965:01-2000:04)



3.2 Methodology

The system given by equations **(3)-(5)** can be written in state space form and Kalman filter estimates of the unknown states as well as maximum likelihood estimates of the parameters in the system can be obtained provided that the endogeneity issues are resolved first (see Hamilton (1994), Chapter 13). Both Δy_t and bc_t – which is proxied by Δu_t – are endogenous, ie they are correlated with the error terms ε_t , ε_t^{α}, and ε_t^{β}. To avoid inconsistent estimation we replace Δy_t and Δu_t by their fitted counterparts that are contemporaneously uncorrelated with the errors in the system. We construct the fitted disposable income growth series Δy_t^f as the fitted values of a regression of disposable income growth on a number of instruments suggested by Campbell and Mankiw (1990), ie lagged disposable income growth, lagged consumption growth, lagged changes in the short-term nominal interest rate and a lagged error correction term, ie log consumption minus log disposable income (see also Campbell (1987)). We construct the fitted change in the unemployment rate series Δu_t^f as the fitted values of a regression of the change in the unemployment rate on lagged changes in the unemployment rate, the lagged NBER dummy, lagged values of the term spread (ie the difference between the short-term and the long-term interest rate), and lagged values of the corporate spread (ie the difference between the interest rate on BAA bonds and the interest rate on AAA bonds). The term spread and the corporate spread are reported by Estrella and Mishkin (1998) as good predictors of US recessions. Note that we use lags 2 to 5 for all instruments except for the error correction term (only lag 2). The reason for

Table A: Statistics for the first-stage OLS regression of Δy_t and Δu_t on instruments

	Δy_t	Δu_t
R^2	0.1850	0.4684
R^2_{adj}	0.1077	0.4098
F	2.3929	7.9904
F (p-val)	0.0063	0.0000

Notes: The instruments used in the regressions are reported in Section 3.2. The F statistic tests the null hypothesis that all the coefficients in the first-stage regression are zero (except for the constant).

starting with lag 2 in the construction of Δy_t^f and Δu_t^f is the presence of an $MA(1)$ term in equation **(3)**. For Δy_t^f and Δu_t^f to be predetermined the instruments must be lagged at least twice. In Table A we report the (adjusted) R^2 and the F test statistic (and p-value) of the first-stage regressions conducted for Δy_t and Δu_t.(6) We note also that our results are robust to the use of alternative instrument sets (eg the inclusion of an additional lag). Results with alternative instrument sets are not reported but are available from the authors upon request.

In Appendix B we report the state space representation of the model. Application of the Kalman filter recursions (see Hamilton (1994), Chapter 13) to the system provides estimates and standard errors for the unobserved excess sensitivity parameter, ie the state β_t. With the Kalman filter the sample log likelihood function can be constructed which is maximised numerically with respect to the unknown parameters in the system (ie the parameters are β_0, β_1, β_2, θ, σ^2_{ε}, $\sigma^2_{\varepsilon^\alpha}$, and $\sigma^2_{\varepsilon^\beta}$). We report these maximum likelihood estimates of the parameters and associated standard errors based on the Hessian. We refer to Appendix B for more details. As a specification test we also calculate the Ljung Box statistic for autocorrelation. This statistic tests whether the so-called one-step ahead prediction errors of the state space system are autocorrelated (see Durbin and Koopman (2001, page 34)).

To estimate the system we use quarterly data for the United States over the period 1965:01-2000:04 (ie we have 144 observations). The effective sample size is 139 since five

(6) Our two-step procedure implies that we have a limited information maximum likelihood ($LIML$) procedure. If, instead, we were to add an equation for the change in the unemployment rate and an equation for the growth rate of disposable income to our state space system and estimate the full system in one step we would have a full information maximum likelihood ($FIML$) procedure. Reasons why the former method may be preferred over the latter are given in Greene (2003, page 509). The most important reason in our context is that the equations for the change in the unemployment rate and for the growth rate in disposable income contain a very large number of variables and therefore a very large number of parameters to estimate. A joint estimation of all parameters is numerically difficult since $FIML$ is non-linear. In a two-step approach, however, most of the parameters are estimated by linear OLS in a first step and the second step non-linear maximum likelihood estimation contains only a small number of parameters.

observations are lost due to lagging. Data are seasonally adjusted where necessary. For c_t we use the log of real per capita expenditures on non-durables and services (excluding shoes and clothing). For y_t we use the log of real per capita disposable income. Both are deflated by the deflator of non-durables and services (excluding shoes and clothing) with base year $1982 = 100$. Expenditures on non-durables and services, disposable income, and the deflator are taken from the National Income and Product Accounts (NIPA). Population data are taken from the US Census Bureau. The unemployment rate u_t is taken from the Bureau of Labor Statistics. With respect to the instruments used in the construction of Δy_t^f and Δu_t^f we note that for the short-run interest rate we use the nominal three-month Treasury Bill rate, for the long-run interest rate we use the ten-year government bond rate (both taken from OECD), and for the corporate spread we use the BAA corporate rate minus AAA corporate rate series as reported by the Federal Reserve Bank of St. Louis.

4 Results

In Table B we present the results from the estimation of the system over the period 1965:01-2000:04 (effective sample period 1966:02-2000:04). First, in column 1, we report the results of estimating the state space model under the restriction that α_t and β_t are constant. We find a value for the excess sensitivity parameter over the sample period of about 0.28 (significant at the 5% level). This value is close to the values of about 0.3 found by Bacchetta and Gerlach (1997) for the United States over the period 1970-95. The question is then whether this value hides important time-variation. In column 2 of Table B we report the results of estimating the system given in equations **(3)-(5)** with the fitted variables Δy_t^f and Δu_t^f used for Δy_t and bc_t and with the 1982 dummy used for lf_t. We find that there is a significant positive impact of changes in the unemployment rate on the excess sensitivity parameter. This suggests that excess sensitivity is significantly higher during recessions. As noted in Table B we find that the average value of β_t during recessions is about 0.37 while during expansions it is about 0.22. While it has the expected sign the estimate for the coefficient on the low frequency control is not significant. To allow for a less drastic shift in excess sensitivity, in column 3 of Table B, we use a deterministic linear time trend for fl_t. Again, we find a significant positive impact of bc_t on the excess sensitivity parameter. The coefficient on the low frequency control is negative but insignificant. We note further that in the time-varying cases we find estimates for the $MA(1)$ parameter θ of about 0.31. This value is close to the theoretical value of this parameter under time aggregation of the variables

Table B: Maximum likelihood estimation of equations (3)-(5), US data, effective sample period 1966:02-2000:04

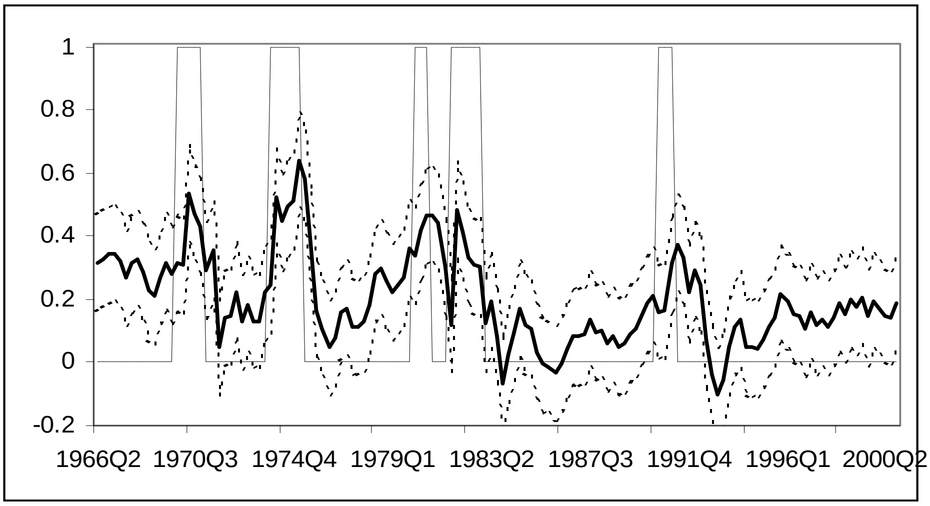
	(1) Time-invariant case	(2) Time-varying case dummy82 for lf_t	(3) Time-varying case linear time trend for lf_t
α_0	0.0042	-	-
	(0.0006)	-	-
β_0	0.2854	0.2459	0.2708
	(0.0921)	(0.1280)	(0.2037)
β_1	-	-0.0818	-0.0009
	-	(0.1898)	(0.0026)
β_2	-	0.5294	0.5392
	-	(0.2778)	(0.2740)
θ	0.3867	0.3133	0.3176
	(0.0786)	(0.1376)	(0.1331)
σ^2_{ε}	1.7E-5	1.3E-5	1.3E-5
	(2.1E-6)	(2.8E-6)	(2.7E-6)
$\sigma^2_{\varepsilon^\alpha}$	-	1.1E-6	1.1E-6
	-	(1.1E-6)	(1.1E-6)
$\sigma^2_{\varepsilon\beta}$	-	0.0087	0.0045
	-	(0.0319)	(0.0311)
$LB(4)$	0.0670	0.2210	0.2080
$LB(8)$	0.2550	0.4970	0.4880
$\overline{\beta}_t^{rec}$	-	0.3676	0.3724
$\overline{\beta}_t^{\exp}$	-	0.2206	0.2173

Notes: Hessian based standard errors between brackets. In the time-invariant case we estimate the equation $\Delta c_t = \alpha_0 + \beta_0 \Delta y_t^f + \varepsilon_t + \theta\varepsilon_{t-1}$. LB(k) denotes the p-value of the Ljung-Box statistic with the null hypothesis of no autocorrelation in the system up to lag k. $\overline{\beta}_t^{rec}$ ($\overline{\beta}_t^{\exp}$) denotes the average value of the excess sensitivity parameter during recessions (expansions) as defined by the NBER turning points.

in the consumption function and continuous decision-making by consumers (see Hall (1988) or Karras (1994)). Finally, we mention that our time-varying specifications are well supported by our Ljung-Box test for autocorrelation. In fact, based on this test, the time-varying cases reported in columns 2 and 3 of Table B are preferred over the time-invariant case reported in column 1.

Graphs of the evolution of the *filtered* estimates for β_t, as implied by the estimations in Table B (column 2), are presented in Chart 2. In this figure the positive impact of recessions on β_t is clear. Also, β_t is slightly declining over time. This reflects the negative (though insignificant) value of the coefficient on the low frequency control lf_t. Finally, while in a few periods β_t is slightly negative, these negative values are never significant.

Chart 2: Time-varying excess sensitivity parameter β_t with 95% confidence bands and NBER turning points (US data, 1966:02-2000:04, result for specification (2) in Table B)



5 Conclusions

We have investigated the impact of business cycle fluctuations on the degree of excess sensitivity (ES) of private consumption growth to disposable income growth by using quarterly US data over the period 1965-2000. Our results suggest that ES is positively affected by the change in the unemployment rate, ie ES is significantly higher during recessions. This result can be reconciled with both the liquidity constraints and the precautionary savings interpretation of ES. We do not find a significant impact on ES of low frequency controls however. These results suggest that short-run factors should be given more weight in future ES studies, especially because the relevance of short-run factors is implied by the economic theories used to explain the observed ES.

Appendix A: Derivation of equation (1)

The representative consumer maximises $E_t \sum_{j=t}^{\infty}(1+\rho)^{-(j-t)}u(C_j)$ with $0 < \rho < 1$ subject to a standard budget constraint with constant interest rate r (C_j is real per capita consumption and E_t is the expectations operator conditional on information available up to period t). With an instantaneous utility function of the constant relative risk aversion type, ie $u(C_j) = (1-\gamma)^{-1}C_j^{1-\gamma}$ with $\gamma > 0$, the first-order condition is $E_{t-1}[X_t] = (1+\rho)(1+r)^{-1}$ with $X_t \equiv (C_t/C_{t-1})^{-\gamma}$. Set $c_t \equiv \ln C_t$, $x_t \equiv \ln X_t$, and $\Delta c_t \equiv \ln(C_t/C_{t-1})$ then $x_t = -\gamma\,\Delta c_t$. Under the assumption that Δc_t is normally distributed with mean $E_{t-1}\Delta c_t$ and variance $V_{t-1}\Delta c_t$ we know that x_t is also Gaussian with mean $-\gamma\,E_{t-1}\Delta c_t$ and variance $\gamma^2 V_{t-1}\Delta c_t$. From the lognormal property we then have that $E_{t-1}(\exp(x_t)) = E_{t-1}[X_t] = \exp(-\gamma\,E_{t-1}\Delta c_t + 0.5\gamma^2 V_{t-1}\Delta c_t)$. After substituting the last expression into the first-order condition, taking logs, and rearranging we obtain $\Delta c_t = \alpha_t + \varepsilon_t$ where $\alpha_t = (r-\rho)\gamma^{-1} + 0.5\gamma\,V_{t-1}\Delta c_t$ and where $\varepsilon_t = \Delta c_t - E_{t-1}\Delta c_t$.

Appendix B: State space representation of the model

We report the state space representation of equations **(3)-(5)** with Δy_t and bc_t replaced by Δy_t^f and Δu_t^f. The state vector is $\mathbf{S}_t$.

$$\Delta c_t = \mathbf{H}_t' \mathbf{S}_t \qquad \textbf{(B-1)}$$

$$\mathbf{S}_t = \mathbf{F}\mathbf{S}_{t-1} + \mathbf{D}\mathbf{Z}_t + \mathbf{v}_t \qquad \textbf{(B-2)}$$

where $\mathbf{H}_t = \begin{bmatrix} 1 \\ \Delta y_t^f \\ 1 \\ \theta \end{bmatrix}$, $\mathbf{D} = \begin{bmatrix} 0 & 0 & 0 \\ \beta_0 & \beta_1 & \beta_2 \\ 0 & 0 & 0 \\ 0 & 0 & 0 \end{bmatrix}$, $\mathbf{S}_t = \begin{bmatrix} \alpha_t \\ \beta_t \\ \varepsilon_t \\ \varepsilon_{t-1} \end{bmatrix}$, $\mathbf{F} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 \end{bmatrix}$, $\mathbf{Z}_t = \begin{bmatrix} 1 \\ lf_t \\ \Delta u_t^f \end{bmatrix}$,

$\mathbf{v}_t = \begin{bmatrix} \varepsilon_t^\alpha \\ \varepsilon_t^\beta \\ \varepsilon_t \\ 0 \end{bmatrix}$ where $\mathbf{v}_t \sim N(\mathbf{O}, \mathbf{Q})$ with $\mathbf{Q} = E(\mathbf{v}_t \mathbf{v}_t') = \begin{bmatrix} \sigma_{\varepsilon^\alpha}^2 & 0 & 0 & 0 \\ 0 & \sigma_{\varepsilon^\beta}^2 & 0 & 0 \\ 0 & 0 & \sigma_\varepsilon^2 & 0 \\ 0 & 0 & 0 & 0 \end{bmatrix}$.

Given that the variables in $\mathbf{H}_t$ and $\mathbf{Z}_t$ are either exogenous or predetermined, the Kalman filter equations (see Hamilton (1994), Chapter 13) can be applied to the system. To initialise the filter we use a diffuse prior, ie we assume that the initial state vector $\mathbf{S}_0$ is random with covariance matrix $\kappa\mathbf{I}$ where $\kappa \to \infty$ and where $\mathbf{I}$ is an identity matrix. We use the Kalman filter to construct the sample log likelihood function which is then maximised numerically with respect to the unknown parameters in $\mathbf{H}_t$, $\mathbf{D}$ and $\mathbf{Q}$. This procedure provides the filtered states $\mathbf{S}_{t|t}$ (for $t = 1, ..., T$), the associated mean squared error matrices $\mathbf{P}_{t|t}$ (for $t = 1, ..., T$) used to construct confidence bounds for the states, and the maximum likelihood estimates of the parameters in $\mathbf{H}_t$, $\mathbf{D}$ and $\mathbf{Q}$. The asymptotic standard errors of the maximum likelihood estimates are calculated from the matrix of second derivatives of the log likelihood function (ie we calculate Hessian based standard errors). We refer to Hamilton (1994, Chapter 13) for details.

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