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Business Cycle Fluctuations and Excess Sensitivity of Private Consumption

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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.

INTRODUCTION

Under very strict assumptions, the permanent income hypothesis implies that aggregate private consumption follows a random walk (Hall 1978); maximizing 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 liberalization 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 could 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 e.g. 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 rationalize 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 e.g. Stiglitz and Weiss 1981; Bernanke and Gertler 1989).[1] The deterioration of households' balance sheets in a recession decreases internal financing possibilities (i.e. 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', i.e. 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) emphasizes 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 estimate simultaneously a consumption growth equation and a multivariate stochastic process for the ES parameter. This approach differs 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 e.g. 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; i.e. it 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.

The paper is structured as follows. In Section I we present the theoretical framework. In Section II we present the empirical specification and discuss the estimation methodology. Section III presents the estimation results, while Section IV concludes.

I. Theoretical Framework

Suppose a representative consumer maximizes 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

$$(1) \qquad \Delta c_t = \alpha_t + \varepsilon_t,$$

where Δc_t is the growth rate of real per capita consumption, α_t encompasses the difference between the interest rate and the rate of time preference and the conditional variance of consumption growth, and 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

$$(2) \qquad \Delta c_t = \alpha_t + \beta_t \Delta y_t + \varepsilon_t,$$

where Δy_t is the growth rate of real per capita disposable income, and β_t is the excess sensitivity parameter ($0 \leqslant \beta_t \leqslant 1$). 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 emphasize 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; 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 liberalization), we also investigate the impact of business cycle fluctuations on β_t. We discuss our empirical specification for β_t in the next section.

II. Empirical Specification and Estimation Methodology

Empirical specification

We consider the following empirical specification:

$$\Delta c_t = \alpha_t + \beta_t \Delta y_t + \varepsilon_t + \theta \varepsilon_{t-1}, \tag{3}$$

$$\alpha_t = \alpha_{t-1} + \varepsilon_t^{\alpha}, \tag{4}$$

$$\beta_t = \beta_0 + \beta_1 lf_t + \beta_2 bc_t + \varepsilon_t^{\beta}. \tag{5}$$

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 $-1 \leqslant \theta \leqslant 1$. The reasons that we allow for an $MA(1)$ error in consumption growth are potential time aggregation (Working 1960), problems related to the 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 (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 (I) and 1 from 1982 (I) onward.[4] We proxy bc_t by the change in the unemployment rate Δu_t. As can be seen in Figure 1, this variable is highly correlated with the turning points of the business cycle as calculated by the National Bureau of Economic Research (i.e. 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 $\sigma_{\varepsilon}^2, \sigma_{\varepsilon^{\alpha}}^2$ and $\sigma_{\varepsilon^{\beta}}^2$, respectively).

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

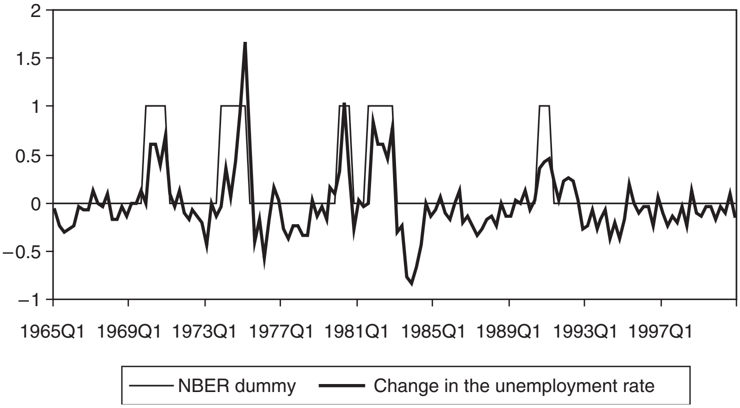

FIGURE 1. NBER turning points and the change in the unemployment rate: US data, 1965 (I)–2000 (IV).

resolved first (see Hamilton 1994, chapter 13). Both Δy_t and bc_t—which is proxied by Δu_t—are endogenous; i.e. 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 which 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), i.e. lagged disposable income growth, lagged consumption growth, lagged changes in the short-term nominal interest rate and a lagged error correction term, i.e. 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 (i.e. the difference between the short-term and the long-term interest rate), and lagged values of the corporate spread (i.e. 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 2–5 for all instruments except for the error correction term (only lag 2). The reason for starting with lag 2 in the construction of Δy_t^f and Δu_y^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 1 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 (e.g. the inclusion of an

TABLE 1
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

Note: The F-statistic tests the null hypothesis that all the coefficients in the first-stage regression are zero (except for the constant).

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, i.e. the state β_t. With the Kalman filter, the sample log likelihood function can be constructed which is maximized numerically with respect to the unknown parameters in the system (i.e. 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. 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, p. 34).

To estimate the system, we use quarterly data for the United States over the period 1965(I)–2000(IV) (i.e. we have 144 observations). The effective sample size is 139, since 5 observations are lost as a result of lagging. Data are seasonally adjusted where necessary. For c_t we use the log of real per capita expenditures on nondurables 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 nondurables and services (excluding shoes and clothing) with base year $1982 = 100$. Expenditures on nondurables and services, disposable income and the deflator are taken from the National Product and Income 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^f_t and Δu^f_t, we note that for the short-run interest rate we use the nominal 3-month Treasury bill rate, for the long-run interest rate we use the 10-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.

III. Results

In Table 2 we present the results from the estimation of the system over the period 1965(I)–2000(IV) (effective sample period 1966(II)–2000(IV)). 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 2 we report the results of estimating the system given in equations (3)–(5) with the fitted variables Δy^f_t and Δu^f_t 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 2, 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 2 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

TABLE 2
MAXIMUM LIKELIHOOD ESTIMATION OF EQUATIONS (3)–(5), US DATA, EFFECTIVE SAMPLE PERIOD 1966 (II)–2000 (IV)

	(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
$\bar{\beta}_t^{rec}$	–	0.3676	0.3724
$\bar{\beta}_t^{exp}$	–	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; $\bar{\beta}_t^{rec}$ ($\bar{\beta}_t^{exp}$) denotes the average value of the excess sensitivity parameter during recession (expansions) as defined by the NBER turning points.

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 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 2 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 2 (column 2), are presented in Figure 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.

IV. 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

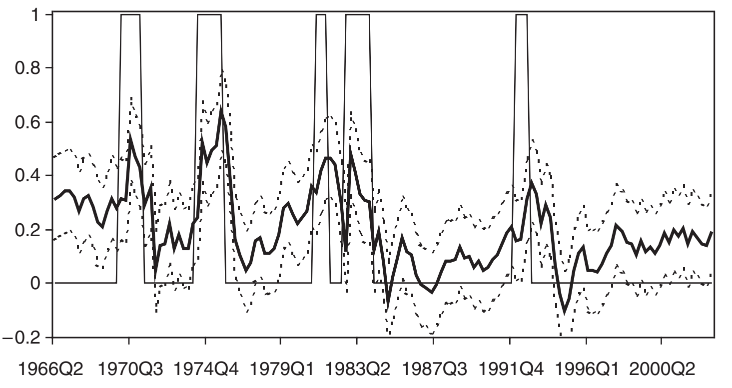

FIGURE 2. Time-varying excess sensitivity parameter β_t with 95% confidence bands and NBER turning points: US data, 1996 (II)–2000 (IV), result for specification (2) in Table 2.

using quarterly US data over the period 1965–2000. Our results suggest that ES is positively affected by the change in the unemployment rate; i.e. 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 maximizes $E_t \sum_{j=t}^{\infty} (1+\rho)^{-(j-t)} u(C_j)$ 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, i.e. $u(C_j) = (1-\gamma)^{-1} C_j^{1-\gamma}$ 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 $\mathbf{S}_t$.

(A1) $\Delta c_t = \mathbf{H}_t' \mathbf{S}_t,$

(A2) $\mathbf{S}_t = \mathbf{F}\mathbf{S}_{t-1} + \mathbf{D}\mathbf{Z}_t + \mathbf{v}_t,$

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 initialize the filter we use a diffuse prior; i.e. 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 maximized 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 $t = 1, \ldots, T$), the associated mean squared error matrices $\mathbf{P}_{t|t}$ (for $t = 1, \ldots, T$) 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 (i.e. we calculate Hessian-based standard errors). We refer to Hamilton (1994, chapter 13) for details.

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NOTES

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 (e.g. 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.
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.
4. This date is qualified by Kaminsky and Schmukler (2003) as the point in time where the domestic financial sector in the United States 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, e.g. 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.
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, p. 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 nonlinear. In a two-step approach, however, most of the parameters are estimated by linear OLS in a first step and the second step nonlinear maximum likelihood estimation contains only a small number of parameters.

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