International Evidence On Sticky Consumption Growth

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Abstract

This paper estimates the degree of 'stickiness' in aggregate consumption growth (sometimes interpreted as reflecting consumption habits) for thirteen advanced economies. We find that, after controlling for measurement error, consumption growth has a high degree of autocorrelation, with a stickiness parameter of about 0.7 on average across countries. The sticky-consumption-growth model outperforms the random walk model of Hall (1978), and typically fits the data better than the popular Campbell and Mankiw (1989) model, though in a few countries the sticky-consumption-growth and Campbell–Mankiw models work about equally well.

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(Contains data and estimation software producing paper's results)

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1 Introduction

A large literature ranging across macroeconomics, finance, and international economics has argued that 'habit formation' can explain many empirical facts related to consumption dynamics.[1] The core empirical pattern driving all these findings appears to be that aggregate consumption growth is too 'sticky' to be explained with standard models. Other explanations for the persistence of aggregate spending growth, or 'excess smoothness' (in Campbell and Deaton (1989)'s terminology), include consumers' inattentiveness to macroeconomic news (Sims (2003); Reis (2006); Carroll and Slacalek (2007)), or their inability to distinguish micro- from macroeconomic shocks (Pischke (1995)). Further explanations could undoubtedly be imagined.

But a full consensus has not emerged on whether empirical data are irreconcilable with Hall (1978)'s benchmark random walk model of consumption. Hall's model implies that consumption growth is unpredictable (excess smoothness is zero). However, standard extensions of the Hall model can generate some degree of stickiness in consumption growth. For example, excess smoothness might merely reflect the fact that spending decisions are made more frequently than consumption data are measured (Working (1960); this viewpoint has recently been advocated in papers by Ludvigson and Lettau (2001); Lettau and Ludvigson (2004)). Also, in the presence of uncertainty, the precautionary motive slows down consumers' response to shocks, which could also explain part (though not all) of the excess smoothness (Ludvigson and Michaelides (2001)). Another possibility, not often mentioned but nevertheless worth serious consideration, is that the smoothness of measured spending reflects data construction methods (e.g. for components of spending for which quarterly observations are imputed using annual data sources). Finally, many of the papers in the habit formation literature have not carefully examined the possibility that their results might reflect the presence of some 'rule-of-thumb' consumers, who simply set consumption equal to income in each period, as proposed in influential papers by Campbell and Mankiw (1989, 1991).

Motivated by this debate and by the fact that much of the empirical evidence on excess smoothness has come from a single country (the U.S.), this paper provides

[1] Facts that have been interpreted using habit formation models include the equity premium puzzle (Constantinides (1990) and Campbell, John Y. and Cochrane (1999)), Granger causality from growth rates to saving rates (Carroll, Overland, and Weil (2000)), the hump-shaped response of consumption to income shocks (Fuhrer (2000)), the dynamic effects of fiscal policy (Ljungqvist and Uhlig (2000)), persistence in current account balances (Gruber (2004)), and the home bias puzzle (Shore and White (2006)). (We do not distinguish here between 'internal' and 'external' habits models because in our view they are empirically indistinguishable using macroeconomic data; see Carroll, Overland, and Weil (1997) for the argument.)

Golinelli and Robert Metz for help in constructing the dataset. Data and econometric programs that generated all of the results in the paper are available from the first author's web page. The views presented in this paper are those of the authors, and should not be attributed to the International Monetary Fund, its Executive Board, or management, or to the European Central Bank.

systematic estimates of three simple canonical models of consumption dynamics using data for all advanced economies for which we were able to construct appropriate datasets (thirteen countries in all). We compare the random walk model of Hall (1978) with two alternatives: the Campbell and Mankiw (1989) model, and a model that permits (but does not require) excess smoothness. We remain deliberately agnostic (in this paper) about whether such smoothness reflects habits, inattention, or other factors; our aim is simply to document the key stylized facts that should be matched by any model of aggregate consumption dynamics.

Using both instrumental variables (IV) (section 3.1) and Kalman filter structural (section 3.2) estimation methods, we find strong evidence of excess smoothness ('stickiness') in consumption growth in every country in our sample.[2] Although there is some variation across countries in the estimated degree of stickiness, in every country we can reject the hypothesis that the stickiness coefficient is zero (the random walk theory), while in no country can we reject the hypothesis that it is 0.7 in quarterly data. Furthermore, wherever there is a clear distinction between the two non-random-walk models, the sticky consumption growth model outperforms the rule-of-thumb model, usually by a decisive statistical margin. (In a few cases, the two non-random-walk models are not statistically distinguishable from each other.)[3]

The large size of our estimated stickiness parameter may come as a surprise to some readers, because the serial correlation coefficient for spending growth in the raw data is much lower than 0.7 (for instance, in U.S. data the OLS estimate of the AR(1) coefficient for nondurables and services consumption growth is about 0.35). The discrepancy reflects our use of econometric methods that are robust to the presence of measurement error. Consistent with Sommer (2007)'s findings for the United States, our estimates suggest that in most countries at least half of the quarterly variation in consumption growth can be interpreted either as measurement error or as truly transitory spending disturbances unrelated to the theoretical consumption model (caused, for example, by unseasonal weather, which can have a nontrivial effect at the quarterly frequency in most countries).[4]

[2]Section 3.2.1 shows how our Kalman filter technique can be interpreted as a particularly simple example of structural estimation of a DSGE model. Embedding our framework in a larger macroeconomic structure would be relatively straightforward.

[3]To our knowledge, the only comparable paper is Braun, Constantinides, and Ferson (1993) (henceforth BCF), who estimate a habit formation model using data on total personal consumption expenditures for six countries. BCF find evidence for stickiness in aggregate consumption growth data in most countries. Their estimates of the habit persistence coefficient range between 0.57 and 0.93, but are often insignificant. Their paper also does not test the assumption of habit formation against alternative models of consumption dynamics, such as the Campbell–Mankiw model. Ferson and Constantinides (1991) report in a framework closely related to BCF that the evidence for habit formation seems stronger in the U.S. data than in their international dataset. However, both papers use GMM to estimate a nonlinear Euler equation, a method which is not robust to the presence of substantial measurement error in consumption data.

[4]Interestingly, Friedman (1957)'s original statement of the permanent income hypothesis gave almost equal billing to transitory consumption shocks and transitory income shocks, but the subsequent literature has focused almost exclusively on income shocks.

The remainder of the paper is organized as follows. Section 2 outlines two theoretical frameworks that generate sticky consumption growth and provide the conceptual framework for our estimation strategy. Section 3 presents the main empirical results and Section 4 concludes.

2 Two Theories of Stickiness

This section sketches the two most popular theoretical frameworks—habit formation and sticky expectations—that can generate serial correlation in aggregate consumption growth. In the habit formation model, the serial correlation coefficient χ reflects the strength of habits (if $\chi = 0$, the model collapses to the Hall random walk model); in the sticky information model, χ is the fraction of aggregate expenditure by households that have not fully updated their information set about the latest macroeconomic developments (and again, $\chi = 0$ corresponds to the Hall model). Because the implications of the two frameworks are indistinguishable in aggregate data, our empirical evidence is consistent with either model.[5]

2.1 Habit Formation

Muellbauer (1988) proposed a simple model of habit persistence, in which the representative consumer maximizes time-nonseparable utility

$$\max \mathbb{E} \sum_{t=s}^{\infty} \beta^{t-s} \mathbf{u}(C_t - \chi C_{t-1}) \tag{1}$$

subject to the usual transversality condition and the dynamic budget constraint:

$$M_{t+1} = (M_t - C_t)R + Y_{t+1}, \tag{2}$$

where β is the discount factor, C is the consumption level, M is market resources (net worth plus current income), R is the constant interest factor, and Y is non-capital income. C_{t-1} in (1) represents the 'habit stock,' i.e., the reference level of consumption to which the consumer compares the current consumption level. The parameter χ captures the strength of habits. After rewriting the utility function as $\mathbf{u}(C_t - \chi C_{t-1}) = \mathbf{u}\big((1-\chi)C_t + \chi \Delta C_t\big)$, one can see that, for $\chi \in (0, 1)$, the consumer derives utility from both the level and the change in consumption.

Dynan (2000) shows that for a habit-forming consumer with Constant Relative Risk Aversion (CRRA) outer utility $\mathbf{u}(Z) = Z^{1-\rho}/(1-\rho)$ and $R\beta = 1$, a first order

[5] For forecasting and some other purposes, it may not matter which theory is closer to the truth. For other purposes (like welfare analysis) the two models could yield quite different conclusions. Carroll and Slacalek (2007) argue that the models can be distinguished using microeconomic data, which suggest the sticky expectations model is closer to the truth.

approximation to the Euler equation leads to consumption dynamics that satisfy:

$$\Delta \log C_t \approx \chi \Delta \log C_{t-1} + \epsilon_t, \tag{3}$$

where ϵ_t mainly reflects innovations to lifetime resources.[6] Hence, in contrast to the standard intertemporally separable utility specification, some of period t's consumption growth is predictable at time $t-1$, and the strength of habits χ can be measured directly by estimating an AR(1) regression like (3) on aggregate consumption data.

2.2 Sticky Expectations

Carroll and Slacalek (2007) present an alternative model that also generates sticky aggregate consumption growth, but without departing from the conventional intertemporally separable utility specification. The key assumption is that consumers are mildly inattentive to macro developments—for example, some households do not immediately notice shocks to aggregate macroeconomic indicators such as productivity growth or the unemployment rate.[7]

Assume that consumers maximize the discounted sum of time separable utility $\sum_{t=s}^{\infty} \beta^{t-s} \mathbf{u}(C_t)$ subject to the budget constraint (2). In a Hall (1978) model with quadratic utility, in which households use all available information, the optimal consumption level follows a random walk: $\Delta C_t = \epsilon_t$. Numerical simulations in Carroll and Slacalek (2007) show that when quadratic utility is replaced with CRRA utility and the model is solved with realistic calibrations of idiosyncratic and aggregate uncertainty, the *log* of aggregate consumption is close to a random walk with drift (the drift reflects the precautionary motive and the attendant nonlinearities): $\Delta \log C_t = \mu + \epsilon_t$.

Suppose now that the economy consists of a continuum of *inattentive* but otherwise-standard CRRA-utility consumers, each of whom updates the information about his permanent income with probability Π in each period. For each consumer, this probability is assumed to be independent of the date when he last updated his information set (and independent of his income, wealth, or other characteristics). This assumption resembles firm behavior in Calvo (1983)'s model of price setting,

[6] We neglect an uninteresting constant term. ϵ_t will also include any higher-order terms that are discarded in the process of the log-linearization, including terms that reflect the precautionary motive. Note, however, that the excess smoothness of aggregate consumption *cannot* be explained by a precautionary saving motive in a model without habits (Ludvigson and Michaelides (2001)). See Michaelides (2002) for a careful numerical examination of a model with both habits and a precautionary motive. Unfortunately, that paper does not examine the accuracy of the approximation (3) in the presence of uncertainty, and we are not aware of any other paper that does so. But Carroll and Slacalek (2007) show that the random walk implication of the model without habits survives largely intact for simulated aggregate data for an economy populated by households facing both idiosyncratic and aggregate risk; this suggests that the log-linearized approximation is likely to be plausible.

[7] The possibility that households do not immediately perceive aggregate shocks is lent credibility by recent work of Aruoba (2008), who shows that advance and preliminary releases of national income accounts data are not unbiased predictors of the final revised data. If even the U.S. national statistical agency is not able to construct an unbiased estimate of the truth immediately, it seems hard to argue against the plausibility of the proposition that households are even slower on the uptake.

which is commonly used in the monetary economics literature. Carroll and Slacalek (2007) show that the change in the log of aggregate consumption, $\Delta \log C_t$, approximately follows an AR(1) process, whose autocorrelation coefficient approximates the share of consumers $(1 - \Pi)$ who do not have up-to-date information about macroeconomic developments. That is, consumption growth is well approximated by:[8]

$$\Delta \log C_t = \mu + \underbrace{(1 - \Pi)}_{\equiv \chi} \Delta \log C_{t-1} + \epsilon_t. \tag{4}$$

In addition, in the spirit of Akerlof and Yellen (1985) and Cochrane (1991), Carroll and Slacalek (2007) show that the utility loss from the infrequent updating of expectations is very small under standard calibrations of the model with $\Pi = 0.25$ per quarter.[9]

3 Empirical Results

This section tests the model of sticky consumption growth (3) and (4) against the alternatives of rule-of-thumb behavior and the random walk hypothesis. The organizing framework for our empirical analysis is a specification for consumption growth from the excess sensitivity literature,[10] which has been expanded here to include a term capturing stickiness of consumption growth:

$$\Delta \log C_t = \varsigma + \chi \, \mathbb{E}_{t-2}[\Delta \log C_{t-1}] + \eta \, \mathbb{E}_{t-2}[\Delta \log Y_t] + \alpha \, \mathbb{E}_{t-2}[a_{t-1}] + \epsilon_t, \tag{5}$$

where Y is household income and a denotes the ratio of household (net) assets to permanent income. The first two right-hand side regressors correspond to two of the tested theories of consumption behavior: inattentiveness or habit formation ($\Delta \log C_{t-1}$) and rule-of-thumb consumers ($\Delta \log Y_t$). Under the third tested theory—the random walk hypothesis—the coefficients χ and η should both be zero. The third term in the equation above (a_{t-1}) is included as a control—any of the three theories allow for some direct effect of asset holdings on consumption growth, either due to effects related to uncertainty (which induces a precautionary saving motive) or due to time variation in interest rates (which we assume is captured by time variation in a).[11]

[8] Sluggish dynamics of aggregate consumption growth are also implied by the 'rational inattention' models of Reis (2006) and Sims (2003).

[9] Carroll (2003) estimates that the probability that a household updates their inflation expectations is 0.27 per quarter, similar to the 0.25 rate assumed in Mankiw and Reis (2001).

[10] Early contributions include Flavin (1981), Campbell and Deaton (1989), and Campbell and Mankiw (1989); for more recent work see, e.g., Luengo-Prado and Sørensen (2008) and the citations therein.

[11] By including the assets in the estimated equation, we follow the literature on precautionary saving and liquidity constraints. The alternative justification for including a is as a proxy for expected interest rates R_{t+1}; calibrated general equilibrium models imply that the relationship between a_t and $\mathbb{E}_t[R_{t+1}]$ is very close to linear. If such

There are at least three reasons to expect the OLS estimates of coefficients in (5) to be biased and inconsistent. First, as argued by Wilcox (1992) and Sommer (2007), quarterly consumption data may be contaminated with substantial measurement error. Second is the undoubted existence of transitory spending disturbances such as those related to weather (or even, for some smaller countries, one-time events like the hosting of the Olympics). Standard theoretical models ignore these kinds of shocks, yet back-of-the-envelope calculations suggest their effects could be substantial in quarterly data. Our final reason for expecting OLS to be biased is the well-known problem of time aggregation.[12]

We develop the points about importance of measurement error and transitory spending fluctuations using the United States as an example. The Bureau of Economic Analysis (2006) describes the methodology by which aggregate expenditures on nondurable goods are estimated using data on retail sales at a sample of retail outlets; since only a subset of retail stores are surveyed, the retail sales figures must contain sampling error. As an example of a "transitory disturbance," under some plausible assumptions, Hurricane Katrina may have reduced quarterly personal consumption expenditure (PCE) growth by about 1 percentage point on an annualized basis in Q3:2005.[13] However, even a much more benign event such as mild winter can reduce annualized quarterly consumption growth significantly—for instance, by about 1/4 percentage point in the United States in Q1:2006—through lower outlays on energy.

To address these three estimation issues (measurement error, transitory consumption, and time aggregation) in quarterly consumption data, we use two econometric methods. The first technique attempts to overcome these problems using instrumental variables estimation. As with any IV method, validity of the results depends on our ability to find suitable instruments (though the extensive literature on the predictability of consumption growth provides good candidates). As an alternative for those who dislike IV regressions, our second technique uses the Kalman filter and structural modeling assumptions to separate 'true' consumption growth from its transitory components and measurement error.[14] In this case, the usual caveat applies: The validity of this maximum likelihood method hinges on the assumed

models are a good way of interpreting the data, the a term should therefore capture the interest rate effects implied by the theory. However, empirical estimates of Euler equations using macro data generally produce insignificant (or even implausible) coefficients on expected interest rates (see, e.g., Hall (1988) and table 3 of Campbell and Mankiw (1991); and Vissing-Jørgensen (2002) for evidence in micro consumption data).

[12] Working (1960)'s analysis shows that if consumers with time separable preferences make purchase decisions more often than consumption data are observed, time aggregation generates an MA(1) process in observed consumption growth even when preferences are otherwise standard as in Hall (1978). In a simple habit formation or sticky information model of the type presented in this paper, time aggregation generates an MA(2) process in consumption growth, but the MA(2) coefficient is generally small.

[13] See Sommer (2007) for details.

[14] Aficionados of Bayesian estimation of DSGE models may wish to reinterpret our estimates as a maximum likelihood estimator of a particularly simple structural model with measurement error and a weak prior. See Section 3.2 for details.

structure of the stochastic processes for measurement error and 'true' consumption dynamics.[15] We view the similarity between the results obtained from these two different methods, along with the coherence of our results with the large literature on habit formation in macroeconomics, as persuasive evidence that stickiness in consumption growth is a robust phenomenon.

3.1 Sticky Consumption Growth in IV Regressions

3.1.1 Dataset

Equation (5) is estimated using aggregate quarterly data for thirteen advanced economies ranging roughly over the past forty years (table 5 provides data details). Our preferred measure of consumption is the sum of expenditures on nondurable goods and services. However, this measure is available only for six countries in our sample (Canada, France, Germany, Italy, the U.K. and the U.S.); total personal consumption expenditures are therefore used for the other sample countries.[16] Finally, Y and a are measured as household disposable income and the ratio of financial wealth to disposable income, respectively.[17]

3.1.2 Instruments

The main advantage of IV estimation is that with appropriate instruments, there is no need to make assumptions about the stochastic structure of measurement error and other transitory fluctuations in quarterly consumption growth. The only requirements are that the instruments are uncorrelated with measurement error and temporary consumption fluctuations, but correlated with the instrumented variables.

Under habit formation or sticky expectations, Sommer (2007) shows that time aggregation makes "true" consumption growth $\Delta \log C_t^*$ (i.e., consumption growth

[15]In principle, the 'habit formation' model could be estimated by GMM, if we were willing to assert that the representative agent model with habits is a perfect description of aggregate consumption choice. One problem with nonlinear GMM estimation is that it is often difficult to be sure how much identification is coming from the higher derivatives of the Euler equation; in a context where there is an unknown amount of measurement error with an unknown distribution, this is worrisome. Also, nonlinear GMM is not really applicable for the sticky expectations model, whose full and precise implication is exactly the linear equation we estimate.

[16]For the six countries for which nondurables and services data are readily available, regression results using total PCE are similar to those reported in the paper for nondurables and services. Since durable consumption growth is generally mildly negatively autocorrelated (Mankiw (1982)), the estimates of consumption persistence χ for the other countries for which we use data on the total PCE (see the bottom panel of table 1) may be biased *downward*, making our evidence in favor of strong consumption stickiness likely to be conservative. For the U.S., it is possible to perform similar estimations using data on purely nondurable goods spending and on retail sales spending, with results similar to those reported here for PCE excluding durables. Japan is not included in our sample for reasons explained in Appendix A.

[17]In the denominator of the wealth–income ratio, we have also experimented with using permanent component of income extracted from the random walk model with transitory noise. Doing so does not practically change the results because, as previous literature has found, essentially all aggregate shocks to the log-level of income are permanent and consequently, in aggregate data permanent income essentially equals actual income.

without measurement error and transitory consumption) follow an ARMA(1,2) process:

$$\Delta \log C_t^* = c_0 + \chi \Delta \log C_{t-1}^* + v_t + \lambda_1(\chi) v_{t-1} + \lambda_2(\chi) v_{t-2}, \tag{6}$$

where the λs are complicated functions of χ. In addition, the MA(2) coefficient λ_2 is close to zero for all reasonable values of $\chi \in (0,1)$, so that $\Delta \log C_t^*$ is approximately ARMA(1,1). Given these considerations, equation (5) can be estimated using the IV estimator with instruments lagged at least twice (e.g., dated as of time $t-2$ and earlier).[18]

The baseline instrument set for the IV regressions consists of variables that are strongly correlated with consumption growth and yet unlikely to be correlated with measurement error: the unemployment rate, a long-term interest rate, and an index of price volatility.[19] Consumer sentiment is also used as an instrument whenever available (the G-7 countries and Australia), as in Carroll, Fuhrer, and Wilcox (1994) and others.

3.1.3 Estimation Results

Table 1 summarizes the baseline estimation results for four alternative econometric specifications nested in equation (5).[20] The left panel reports the results from univariate regressions in which each right-hand side variable enters the estimated specification as the only regressor. The first column presents the IV estimates of consumption persistence χ, which are for all countries much higher than the (unreported) OLS estimates and are always highly statistically significant.[21] The IV estimates of consumption persistence in table 1 are on average about 0.7—a strong rejection of the random walk proposition which implies a coefficient of zero. The second column reports p values of the null hypothesis $\chi = 0$ implied by the heteroscedasticity and autocorrelation robust version of the conditional likelihood

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[18] Ideally, it would be desirable to use instruments dated $t-3$ or earlier, but for some countries the $t-3$ instruments did not have sufficient predictive power for the instrumented variables.

[19] Price volatility is robustly negatively correlated with real consumption growth in all sample countries—this relationship is known among business cycle forecasters as the 'Katona Effect'; see, e.g., Okun (1981), p. 216. In economic terms, periods of above-average price volatility tend to be associated with shocks that may also have an impact on permanent income. This instrument is attractive because it can be readily calculated for any country and it is unlikely to be correlated with measurement error in consumption growth. The variable appears to be used in the professional forecasting community but is not as common in academic work. Price volatility at time t, V_t^P, is calculated as the coefficient of variation over the past four quarters: $V_t^P = \sigma_{t-3,t}^P / \mu_{t-3,t}^P$, where $\sigma_{t-3,t}^P = \sqrt{1/4 \times \sum_{i=0}^{3} (P_{t-i} - \mu_{t-3,t}^P)^2}$ is the standard deviation of price level between quarters $t-3$ and t and $\mu_{t-3,t}^P = 1/4 \times \sum_{i=0}^{3} P_{t-i}$ denotes the mean of price level P between quarters $t-3$ and t. To calculate price volatility we use quarterly data on consumption (PCE) deflator.

[20] An advantage of our reduced-form estimates of the consumption function over the estimated dynamic stochastic general equilibrium models (DSGE, which started with the influential work of Smets and Wouters (2003); see An and Schorfheide (2007) for a review) is that we do not use informative priors.

[21] The OLS estimates, and many further results, can be obtained by downloading the replication archive available at the first author's website.

ratio (HAR-CLR) test of Andrews, Moreira, and Stock (2004). The test is robust to potentially weak instruments and is effectively uniformly most powerful among tests invariant to rotations of the instruments. The p values indicate that the zero restriction on χ is soundly rejected in almost all countries.

The third column estimates the Campbell–Mankiw model. Our results are broadly consistent with the evidence presented in Campbell and Mankiw (1991): Rule-of-thumb consumers (for whom, by assumption, consumption equals current income) are on average estimated to earn about $\eta \approx 0.4$ of aggregate income. Interestingly, the estimates of η in the left panel are often less significant than those of consumption persistence χ and are in three or four cases insignificant (depending on whether the standard or HAR-CLR p values are used). This means that—aside from the question of how the Campbell–Mankiw model stands up against the alternative of habit formation or sticky expectations—rule-of-thumb spending behavior cannot be reliably detected in about a third of our sample countries.

The fifth column investigates the relative importance of wealth (expressed as the ratio of net financial assets to income) in aggregate consumption dynamics. The coefficient on the wealth–to–income ratio, α, turns out to be statistically significant only in four countries, although the HAR-CLR p values suggest more often that α is not zero. In addition, the coefficient α has in most countries the opposite sign to that predicted by either precautionary saving theory or intertemporal substitution as channelled through the interest rate. This is unsurprising for at least two reasons. First, the overwhelming significance of consumption (and also income) in the previous regressions implies a severe omitted-variable bias problem with the univariate regression that only includes wealth. Second, the previous literature generally finds little evidence of interest rate or precautionary saving effects in aggregate consumption growth data.[22]

The last column of the left panel displays the adjusted R^2s from the first-stage regressions of consumption growth on instruments (denoted $\bar{R}_c^2$). This measure of the strength of instruments ranges between 0.1 and 0.2 for most countries.[23,24]

The right panel of table 1 reports estimation results when all three regressors are included in equation (5). The results strongly suggest that past consumption growth is by far the strongest predictor of current consumption growth. The average persistence parameter in the country regressions falls only very slightly compared

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[22] Microeconomic evidence suggests that the precautionary saving motive may be an important determinant of household-level consumption decisions, see for example Carroll and Samwick (1997), Gourinchas and Parker (2002), and Fuchs-Schündeln and Schündeln (2005).

[23] Ideally, one would prefer first stage $\bar{R}_c^2$ coefficients larger than those generated by our instrument set for some countries. For each individual country it is possible to find a country-specific instrument set that performs considerably better than our universal instrument set. However, we preferred to run the well-understood risks of weak instruments (coefficients biased toward the OLS value) rather than the much more difficult to quantify risks associated with cherry picking a different instrument set for each country.

[24] The adjusted R^2s from the first-stage regressions on income growth are comparable with the R^2s from the first-stage regressions on consumption growth. The R^2s are much higher for the wealth–to–income ratio, about 0.8.

with the average estimates from univariate regressions reported in the left panel (from $\chi \approx 0.7$ to $\chi \approx 0.6$) and remains statistically significant at the five percent level in ten of our thirteen countries. The predicted income growth term dominates the lagged consumption term only in one country, Germany.[25] The last column of the right panel reports the p-values of the Hansen's overidentification test—results of which imply that the null of instrument exogeneity cannot be rejected.

Table 2 averages the coefficient estimates from table 1 across various country groups. As in table 1, while the average consumption persistence χ falls relatively little after income and wealth are added to the estimated equations (compare the right and left panels of the table), the income and wealth coefficients become essentially zero. The result holds for all five groups of countries reported in the table which suggests considerable homogeneity in χ among advanced economies, a fact already apparent in the previous table with the results for individual countries.

Table 3, whose format is identical to table 1, estimates aggregate consumption dynamics with an alternative instrument set, in which long-run interest rates and price volatility have been replaced with income growth and the interest-rate spread.[26]

The estimation results are broadly consistent with our baseline: (i) the coefficient on lagged consumption growth in univariate regressions is large and significant for ten countries, (ii) in the regressions that include all three regressors, the coefficients on instrumented income growth and wealth tend to be small and less often statistically significant compared with univariate regressions, and (iii) lagged consumption growth beats lagged income in nine horse-race regressions (but gets badly beaten in German data).

3.2 Kalman Filter/Maximum Likelihood Evidence on Sticky Consumption Growth

As a more efficient alternative to IV, we also estimate the dynamics of consumption growth using the Kalman filter. To proceed, it is necessary to specify an assumption about the stochastic process of measurement error. We follow the methodology of Sommer (2007) and assume that measurement error in the log-level of consumption

[25] Germany tends to be an outlier in all our IV regressions (reported and unreported). This may reflect difficulties associated with comparing pre- and post-reunification German data. Prior to reunification in 1991, the German data reflect only West German economic growth. Subsequent to unification, they are for the united Germany. We include a dummy for the quarter of reunification, but it would be surprising if there were structural stability across such an extreme event. These problems are compounded by the highly erratic behavior of German consumption growth during the years immediately following reunification.

[26] The interest rate spread is a variable that has long been used in the consumption growth literature, having first been shown to have robust explanatory power for consumption growth in the literature testing Hall's random walk theory in the 1980s. Interest rates are measured essentially without error, and we can think of no reason the spread would be correlated with measurement error or transitory disturbances to consumption growth. Income growth has been perhaps the most intensively studied variable in this literature, dating back to the original work of Hall (1978) and Flavin (1981).

follows an MA(1) process.[27] Observed consumption growth, $\Delta \log C_t$, can be written as the sum of 'true' consumption growth, $\Delta \log C_t^*$, and a measurement error, u_t, as follows:

$$\Delta \log C_t = \Delta \log C_t^* + u_t + (\theta - 1)u_{t-1} - \theta u_{t-2}, \tag{7}$$

$$\Delta \log C_t^* = c_0 + \chi \Delta \log C_{t-1}^* + v_t + \lambda_1(\chi)v_{t-1} + \lambda_2(\chi)v_{t-2}. \tag{8}$$

As noted above, λs are not free parameters but are complicated functions of χ. The Kalman filter jointly estimates the sticky expectations coefficient χ and the degree of the first autocorrelation in measurement errors, θ. The filter also generates separate estimates of 'true' consumption growth, $\Delta \log C_t^*$, and the measurement error component, u_t. For the purposes of this subsection, we assume that the correlation structure of measurement error remains unchanged over the sample period.

The model described in equations (7) and (8) has been rewritten in a state-space form (see appendix B) and estimated using consumption data for the countries in our dataset (listed in table 5). Table 4 presents the estimation results. As in the case of the IV estimation, the coefficient reflecting consumption growth stickiness, χ, is large and highly statistically significant in almost all sample countries. The value of χ typically ranges between 0.6 and 0.8, with only Denmark and the United Kingdom having coefficients estimated below 0.4. For the United States, the estimated consumption persistence is about 0.7, which is consistent with previous studies (e.g. Fuhrer (2000)).

It is encouraging that the Kalman filter estimates of consumption persistence tend to be close to the IV estimates. This suggests that stickiness of consumption growth is a robust feature of the data that appears similarly even when viewed through quite different lenses.

The estimation results also suggest that measurement error in the level of consumption is positively and significantly autocorrelated in about half of our sample countries—a fact that is not surprising given the interpolation techniques that are often used by statistical agencies when constructing quarterly consumption data.

The Kalman filter's estimate of "true" consumption growth, $\Delta \log C_t^*$, is presented, along with the raw data, in figures 1 and 2. The Kalman filter estimation suggests that the share of transitory components in published quarterly consumption data is large (about 50 percent for the United States and even more for some countries).[28]

[27] Taking a classical approach with white noise measurement error in the level of consumption is *a priori* not justifiable because all three main measurement error types are likely to be serially correlated. The measurement error is therefore allowed to be serially correlated in our model but the impact of error on the serial correlation properties of the consumption data is limited.

[28] There is an interesting link between the signal-to-noise ratio from the estimated Kalman filter models, $\text{var}(\Delta \log C_t^*)/\,\text{var}(\Delta \log C_t)$, in table 4 and the first-stage R^2 for consumption growth from the IV regressions in table 1. The correlation between the two statistics is about 80 percent across countries, confirming that consumption growth can be predicted better in the countries with smaller measurement error and transitory fluctuations.

To see how the restrictions on λs imposed by the theoretical model with habits affect estimates of χ, we have also experimented with several versions of model (7)–(8) in which λs are free parameters (rather than known functions of χ). In such models, consumption sluggishness χ robustly turns out to be similar to the values shown in Table 4. However, the fact that in a few cases λs appear unrealistic (greater than one or smaller than minus one) suggests that imposing theoretical restrictions is helpful in identifying them (rather than χ).

3.2.1 Relationship with the Structural Estimation Literature

The state-space representation (7)–(8) fits nicely into the structural DSGE framework recently proposed by Ireland (2004), who estimates a small log-linearized model with the Kalman filter. Control variables $\mathbf{f}_t$ in his model can be solved in terms of state variables $\mathbf{s}_t$ and residuals $\mathbf{u}_t$:

$$\mathbf{f}_t = \mathbf{C}\mathbf{s}_t + \mathbf{u}_t. \tag{9}$$

Ireland, p. 1210 views the disturbances $\mathbf{u}_t$ as follows: "the residuals $[\mathbf{u}_t]$ may ... soak up both measurement errors, but they can be interpreted more liberally as capturing all of the movements and co-movements in the data that the real business cycle model, because of its elegance and simplicity, cannot explain." Once we plug our transition equation for consumption growth (8) into the measurement equation (7), the Kalman filter model we estimate above has exactly the structure (9) with $\mathbf{f}_t = \Delta \log C_t$, $\mathbf{s}_t = \Delta \log C^*_{t-1}$, $\mathbf{u}_t = u_t + (\theta - 1)u_{t-1} - \theta u_{t-2} + v_t + \lambda_1(\chi)v_{t-1} + \lambda_2(\chi)v_{t-2}$ and $\mathbf{C} = \chi$.

Thus the state-space representation (7)–(8) can be interpreted as a stripped-down version of Ireland's model with consumption habits in which measured consumption is affected by a combination of measurement errors u_t and shocks v_t to "true" consumption C^*_t. As our main goal is to estimate consumption stickiness χ, we do not take a stand on where the consumption shocks v_t come from (be it news about income, wealth, interest rates, fiscal policy or something else). Our model is simple enough to be estimable using classical techniques, including the maximum likelihood estimator, so that data have complete control over the estimates of χ, in contrast to larger-scale DSGE models, which are often inevitably estimated with Bayesian methods with informative priors.

4 Conclusions

Hall (1978) provided macroeconomists with a clean theoretical benchmark to which actual consumption data could be compared: Consumption growth should be essentially unpredictable. In contrast with this benchmark, we find that, when econometric techniques that account for measurement error are used, consumption growth exhibits a high degree of persistence or "momentum." The stickiness of aggregate consumption growth can be interpreted as reflecting the behavior of fully informed households with a strong consumption habit, or the behavior of an aggregate economy in which households are not always perfectly up to date in their knowledge of macroeconomic developments. Fitting the model to data from thirteen countries, we estimate that consumption growth persistence is always significantly above the random-walk benchmark of 0 and is never robustly different from about 0.7. Our analysis also suggests that, on balance, the model of sticky consumption growth describes aggregate consumption data better than the rule-of-thumb model of Campbell and Mankiw (1989), although our point estimates do typically indicate that a modest proportion of aggregate income (in the range of 10–20 percent) may be received by households who consume their current income every quarter.[29]

Our findings imply that the large literature claiming to find evidence of sticky consumption growth in the U.S. probably cannot be explained away as reflecting time aggregation problems or other mistreatment of the data, suggesting that many of the insights gleaned from that literature are likely applicable to other countries as well. (However, it is worth bearing in mind that analyses that rely heavily on the literal interpretation of the habits-in-the-utility-function framework, such as calculations of the welfare cost of aggregate fluctuations, may not hold up under alternative interpretations of consumption growth stickiness.)

Our analysis also strengthens a key policy message about the sluggish average response of consumption to monetary and fiscal policy innovations highlighted earlier in the context of the habit formation literature—an important policy consideration at the current cyclical juncture in many countries, including in the United States.

[29] If these households are poorer than the average, they may constitute a larger proportion of the population than they do of aggregate income.

Table 1 Consumption Dynamics—All Countries

$$\Delta \log C_t = \varsigma + \chi\mathbb{E}_{t-2}[\Delta \log C_{t-1}] + \eta\mathbb{E}_{t-2}[\Delta \log Y_t] + \alpha\mathbb{E}_{t-2}[a_{t-1}]$$

	Estimation with one regressor only							Estimation with all three regressors			
Country	χ	CLR p val$^{\diamond}$ $\chi = 0$	η	CLR p val$^{\diamond}$ $\eta = 0$	α	CLR p val$^{\diamond}$ $\alpha = 0$	$\bar{R}^2_c$	χ	η	α	OID
G7 Countries											
Canada*	0.72***	0.0000	0.32***	0.0020	0.31	0.0101	0.17	0.64***	0.05	0.11	0.85
France*	0.61***	0.0000	0.29***	0.0000	0.04	0.2595	0.03	0.44	0.19	−0.04	0.94
Germany*	0.40*	0.2932	0.72***	0.0033	−0.38	0.9567	0.05	0.16	0.66***	−0.17	0.81
Italy*	0.65***	0.0000	0.20**	0.0154	−0.06	0.3906	0.05	0.53**	0.13	−0.02	0.65
United Kingdom*	0.83***	0.0000	0.10	0.0000	0.32**	0.0008	0.12	1.00***	−0.17	0.01	0.95
United States*	0.83***	0.0000	0.54***	0.0056	0.27*	0.0191	0.17	0.55***	0.27*	0.02	0.95
Mean G7	0.67***	–	0.36***	–	0.08	–	–	0.55**	0.19	−0.01	–
Other Countries											
Australia‡	0.54***	0.0000	0.12	0.0802	0.10	0.0022	0.11	0.51**	0.03	0.01	0.63
Belgium‡	0.64***	0.0025	0.34**	0.0003	0.12	0.0200	0.09	0.56**	0.12	0.01	0.72
Denmark‡	0.86***	0.0000	0.43	0.0001	−0.28	0.0017	0.08	0.78***	0.27	−0.32	0.57
Finland‡	0.90***	0.0000	0.61**	0.0000	0.55	0.0233	0.26	0.86***	0.07	−0.13	0.71
Netherlands‡	0.70***	0.4322	0.09	0.5064	0.22	0.1872	0.00	0.53	−0.14	0.10	0.82
Spain‡	0.94***	0.0000	0.79***	0.0000	0.82***	0.0000	0.38	0.71**	0.04	0.21	0.38
Sweden‡	0.83***	0.0000	0.37**	0.3765	0.61***	0.0384	0.16	0.88***	0.32*	−0.25	0.39
Mean Other	0.77***	–	0.39	–	0.31	–	–	0.69***	0.10	−0.05	–

Notes: Left Panel: Regressions were estimated with one regressor only. Right Panel: Regressions were estimated with all three regressors. $\diamond$: p value of the null hypothesis that the parameter equals 0 tested using the HAC robust version of the conditional likelihood ratio (HAR-CLR) test of Andrews, Moreira, and Stock (2004), window: 4 lags. Consumption variable: $\star$: nondurables, semidurables and services consumption, $\ddagger$: total personal consumption expenditure, a: ratio of household financial wealth to income. $\{^{*}, ^{**}, ^{***}\}$ = Statistical significance at $\{10, 5, 1\}$ percent (using robust standard errors). $\bar{R}^2_c$: Adjusted R^2 from the first-stage regression of consumption growth on instruments. OID: p-value from the Hansen's J statistic for overidentification.

Table 2 Consumption Dynamics—Groups of Countries (Simple Averages)

$$\Delta \log \mathbf{C}_t = \varsigma + \chi \mathbb{E}_{t-2}[\Delta \log \mathbf{C}_{t-1}] + \eta \mathbb{E}_{t-2}[\Delta \log \mathbf{Y}_t] + \alpha \mathbb{E}_{t-2}[a_{t-1}]$$

	Estimation with one regressor only			Estimation with all three regressors		
Country	χ	η	α	χ	η	α
All Countries	0.73^{***}	0.38^{**}	0.19	0.63^{**}	0.14	-0.03
	(0.18)	(0.18)	(0.19)	(0.25)	(0.21)	(0.16)
G7 Countries	0.67^{***}	0.36^{***}	0.08	0.55^{**}	0.19	-0.01
	(0.18)	(0.11)	(0.19)	(0.23)	(0.14)	(0.12)
Anglo–Saxon	0.73^{***}	0.27^{**}	0.24	0.68^{***}	0.04	0.04
	(0.16)	(0.11)	(0.18)	(0.22)	(0.14)	(0.12)
Euro Area	0.69^{***}	0.43^{**}	0.19	0.54^{**}	0.15	-0.01
	(0.18)	(0.20)	(0.18)	(0.27)	(0.22)	(0.13)
European Union	0.73^{***}	0.39^{*}	0.18	0.65^{**}	0.15	-0.06
	(0.18)	(0.20)	(0.20)	(0.26)	(0.23)	(0.17)

Notes: Instruments: Lags $t-2, t-3$ and $t-4$ of the unemployment rate, long-run interest rate, price volatility and consumer sentiment. Left Panel: Regressions were estimated with one regressor only. Right Panel: Regressions were estimated with all three regressors. Robust standard errors are in parentheses. $\{^{*}, ^{**}, ^{***}\}$ = Statistical significance at $\{10, 5, 1\}$ percent. Standard errors are simple averages of individual countries in a given group.

All countries: Canada, France, Germany, Italy, the United Kingdom, the United States, Australia, Belgium, Denmark, Finland, the Netherlands, Spain, Sweden. G7 countries: Canada, France, Germany, Italy, the United Kingdom, the United States. Anglo–Saxon Countries: Australia, Canada, the United Kingdom, the United States. Euro Area Countries: France, Germany, Italy, Belgium, Finland, the Netherlands, Spain. European Union: France, Germany, Italy, the United Kingdom, Belgium, Denmark, Finland, the Netherlands, Spain, Sweden.

Table 3 Consumption Dynamics—All Countries

$$\Delta \log C_t = \varsigma + \chi \mathbb{E}_{t-2}[\Delta \log C_{t-1}] + \eta \mathbb{E}_{t-2}[\Delta \log Y_t] + \alpha \mathbb{E}_{t-2}[a_{t-1}]$$

	Estimation with one regressor only							Estimation with all three regressors			
Country	χ	CLR p val$^\diamond$ $\chi = 0$	η	CLR p val$^\diamond$ $\eta = 0$	α	CLR p val$^\diamond$ $\alpha = 0$	$\bar{R}^2_c$	χ	η	α	OID
G7 Countries											
Canada*	0.69***	0.0000	0.33***	0.0000	0.83**	0.0000	0.16	0.40	0.16	0.35	0.36
France*	0.03	0.0000	0.23*	0.0000	0.07	0.0000	0.16	−0.31	0.36	0.09	0.32
Germany*	0.02	0.0031	0.88***	0.0000	−0.33	0.0091	0.07	−0.14	0.89***	−0.18	0.96
Italy*	0.62***	0.0000	0.29*	0.0000	−0.05	0.1520	0.11	0.55***	0.10	−0.02	0.23
United Kingdom*	0.41**	0.0000	0.07	0.0008	0.29	0.0016	0.23	0.58**	−0.20	0.13	0.67
United States*	0.74***	0.0000	0.41***	0.0004	0.24	0.2905	0.17	0.53**	0.16	0.04	0.72
Mean G7	0.42**	–	0.37***	–	0.17	–	–	0.27	0.25	0.07	–
Other Countries											
Australia‡	0.71***	0.0000	0.18	0.0000	0.13	0.3919	0.07	0.73***	−0.05	0.02	0.76
Belgium‡	0.71***	0.0071	0.27*	0.1195	0.10	0.5938	0.09	0.77**	0.13	−0.09	0.69
Denmark‡	0.35	0.0041	0.10	0.2857	−1.42*	0.0230	0.07	0.19	−0.00	−1.14*	0.53
Finland‡	0.88***	0.0001	0.49***	0.0220	2.89***	0.0000	0.22	0.56*	0.15	0.82	0.66
Netherlands‡	0.75***	0.0000	0.16	0.1855	0.13	0.7037	0.11	0.71***	0.14	0.02	0.82
Spain‡	0.94***	0.0000	0.58***	0.0000	0.94***	0.0000	0.38	0.67**	0.18	0.15	0.86
Sweden‡	0.86***	0.0001	0.05	0.0128	0.88***	0.0217	0.20	0.85**	−0.03	0.04	0.36
Mean Other	0.74***	–	0.26	–	0.52	–	–	0.64**	0.07	−0.03	–

Notes: Left Panel: Regressions were estimated with one regressor only. Right Panel: Regressions were estimated with all three regressors. $\diamond$: p value of the null hypothesis that the parameter equals 0 tested using the HAC robust version of the conditional likelihood ratio (HAR-CLR) test of Andrews, Moreira, and Stock (2004), window: 4 lags. Consumption variable: $\star$: nondurables, semidurables and services consumption, $\ddagger$: total personal consumption expenditure, a: ratio of household financial wealth to income. $\{^{*}, ^{**}, ^{***}\}$ = Statistical significance at $\{10, 5, 1\}$ percent (using robust standard errors). $\bar{R}^2_c$: Adjusted R^2 from the first-stage regression of consumption growth on instruments. OID: p-value from the Hansen's J statistic for overidentification.

Table 4 Consumption Dynamics—First-Stage Kalman Filter Estimates

$$\begin{aligned}\Delta \log C_t &= \Delta \log C_t^* + u_t + (\theta - 1)u_{t-1} - \theta u_{t-2},\\ \Delta \log C_t^* &= c_0 + \chi \Delta \log C_{t-1}^* + v_t + \lambda_1(\chi)v_{t-1} + \lambda_2(\chi)v_{t-2}\end{aligned}$$

	Parameter Estimates				
Country	χ	θ	$\log \sigma_u^2$	$\log \sigma_v^2$	$\frac{\mathrm{var}(\Delta \log \mathbf{C}_t^*)}{\mathrm{var}(\Delta \log \mathbf{C}_t)}$
G7 Countries					
Canada*	0.78***	0.25**	−11.03***	−13.02***	0.18
France*	0.81***	−0.01	−11.42***	−14.00***	0.10
Germany*	0.83***	0.25*	−9.97***	−12.49***	0.14
Italy*	0.62***	−0.08	−12.04***	−12.26***	0.37
United Kingdom*	0.36***	−1.00	−12.21***	−10.79***	0.39
United States*	0.67***	0.30**	−12.26***	−12.58***	0.44
Other Countries					
Australia‡	0.49*	0.23	−10.78***	−11.50***	0.21
Belgium‡	0.70***	0.39***	−11.44***	−11.83***	0.45
Denmark‡	0.39*	−0.23	−10.38***	−9.85***	0.38
Finland‡	0.72***	0.20	−10.95***	−11.00***	0.55
Netherlands‡	0.90***	−0.08	−9.85***	−12.64***	0.18
Spain‡	0.84***	0.23	−12.08***	−11.39***	0.82
Sweden‡	0.67***	0.27*	−11.71***	−11.40***	0.60

Notes: Consumption variable: $\star$: nondurables, semidurables and services consumption, $\ddagger$: total personal consumption expenditure. $\{^*, ^{**}, ^{***}\}$ = Statistical significance at $\{10, 5, 1\}$ percent.

Figure 1 Measured and "True" Consumption Growth—G7 Countries

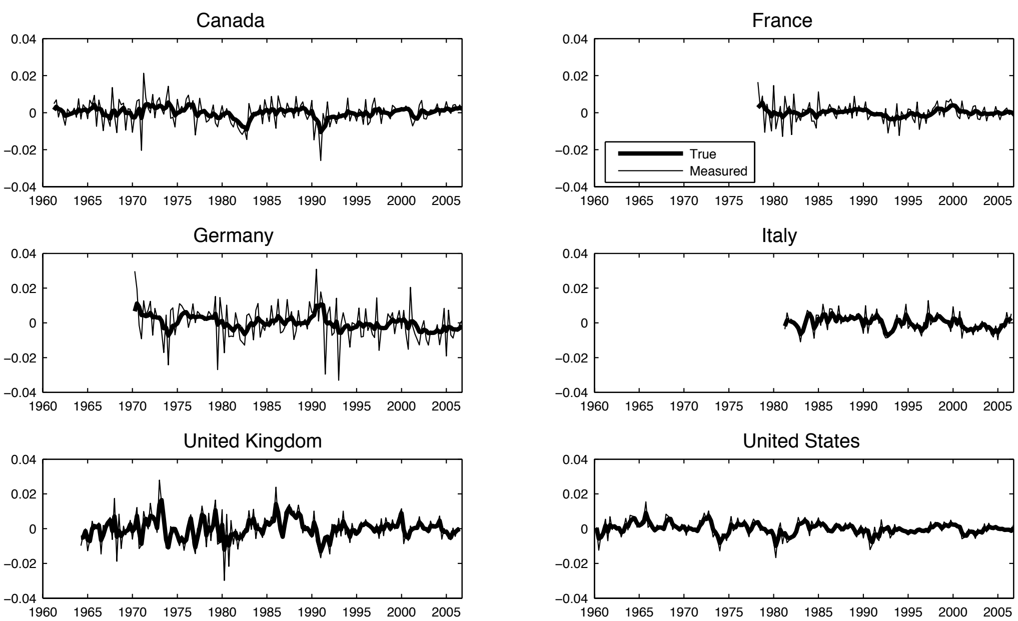
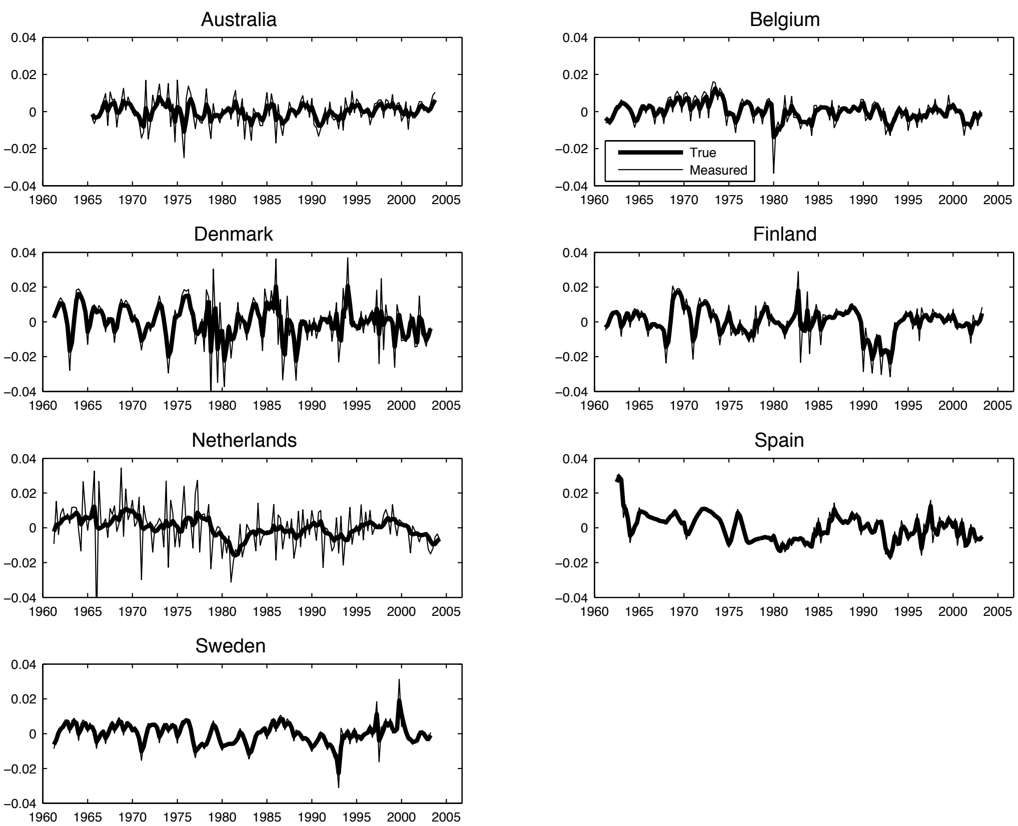


Figure 2 Measured and "True" Consumption Growth—Other Countries



Appendix A: Description of Data

Data for the G-7 economies are from the Haver Analytics database. Data for other countries are from the database of the NiGEM model of the NIESR Institute, London. The original sources for most of these data are OECD, Eurostat, national statistical offices and central banks. Income is measured as personal disposable income. Wealth is approximated using data on the net financial wealth. All series were deflated with consumption deflators and expressed in per capita terms. The population series are from DRI International and were interpolated from annual data to quarterly observations. Japan is not included in our sample as creating a quarterly dataset with consumption data going prior to 1980 would involve splicing consumption series based on three very different methodologies. Adjustments to the Japanese national accounts methodology in 2002 and 2004 have significantly improved the reliability of quarterly consumption series but the current-methodology data are only available since Q1:1994 (International Monetary Fund (2006)).

We thank Roberto Golinelli for consumer sentiment series for G7 countries and Australia used (and described in detail) in Golinelli and Parigi (2004). (We have not used consumer sentiment series for the remaining countries, because the data are not available before 1985.) We are grateful to Carol Bertaut and Nathalie Girouard for providing us with the data used in Bertaut (2002) and Catte, Girouard, Price, and Andre (2004), respectively. Ray Barrell, Amanda Choy and Robert Metz answered our questions about the NiGEM's database.

Appendix B: Details of the Kalman Filter Estimation

Following Sommer (2007), equations (7) and (8) can be rewritten in the state-space form with the measurement equation:

$$\Delta \log C_t = c_0 + \begin{bmatrix} 1 & 0 & 0 & 1 & 0 & 0 \end{bmatrix} \begin{bmatrix} \Delta \log C_t^* \\ u_t \\ -u_t + \theta \Delta u_t \\ \Delta u_t + \theta \Delta u_{t-1} \\ v_t \\ v_{t-1} \end{bmatrix} + 0,$$

Table 5 Consumption Data, Its Sources, and Samples for IV Regressions

Country	Time Frame	Consumption/Source	Income/Source	Wealth/Source
G7 Countries				
Canada	Q4:1970–Q3:2002	NDS/Haver	PDI/Haver	NFW/NiGEM
France	Q1:1985–Q4:2003	NDS/Haver	PDI/Haver	NFW/NiGEM
Germany‡	Q4:1975–Q4:2002	NDS/Haver	PDI/Haver	NFW/NiGEM
Italy	Q1:1981–Q4:2003	NDS/Haver	PDI/Haver	NFW/NiGEM
United Kingdom	Q1:1974–Q4:2003	NDS/Haver	PDI/Haver	NFW/NiGEM
United States	Q3:1962–Q2:2004	NDS/Haver	PDI/Haver	NFW/NiGEM
Other Countries				
Australia	Q4:1975–Q4:1999	PCE/Haver	PDI/Haver	NFW/NiGEM
Belgium	Q2:1980–Q4:2002	PCE/NiGEM&MEI	PDI/NiGEM&MEI	NFW/NiGEM
Denmark	Q1:1977–Q2:2003	PCE/NiGEM&MEI	PDI/NiGEM&MEI	NFW/NiGEM
Finland	Q3:1973–Q2:2003	PCE/NiGEM&MEI	PDI/NiGEM&MEI	NFW/NiGEM
Netherlands	Q1:1975–Q4:2002	PCE/NiGEM&MEI	PDI/NiGEM&MEI	NFW/NiGEM
Spain	Q1:1978–Q4:1999	PCE/NiGEM&MEI	PDI/NiGEM&MEI	NFW/NiGEM
Sweden	Q1:1977–Q4:2002	PCE/NiGEM&MEI	PDI/NiGEM&MEI	NFW/NiGEM

Notes: PCE = Total personal consumption expenditures, NDS = Nondurables and services, PDI = Personal disposable income, NFW = Net financial wealth, ‡: Regressions for Germany were estimated with a reunification dummy in Q1:1991; Source: Haver—Haver Analytics, NiGEM—Database of the NiGEM model of the NIESR Institute, London, MEI—Main Economic Indicators of OECD.

and the state-evolution equation:

$$\begin{bmatrix} \Delta\log C_t^* \\ u_t \\ -u_t+\theta\Delta u_t \\ \Delta u_t+\theta\Delta u_{t-1} \\ v_t \\ v_{t-1} \end{bmatrix} = \begin{bmatrix} \chi & 0 & 0 & 0 & \lambda_1 & \lambda_2 \\ 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & -\theta & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 \end{bmatrix} \begin{bmatrix} \Delta\log C_{t-1}^* \\ u_{t-1} \\ -u_{t-1}+\theta\Delta u_{t-1} \\ \Delta u_{t-1}+\theta\Delta u_{t-2} \\ v_{t-1} \\ v_{t-2} \end{bmatrix} + \begin{bmatrix} v_t \\ u_t \\ (\theta-1)u_t \\ u_t \\ v_t \\ 0 \end{bmatrix},$$

and with the associated covariance matrices $H = 0$ and

$$Q = \begin{bmatrix} \sigma_v^2 & 0 & 0 & 0 & \sigma_v^2 & 0 \\ 0 & \sigma_u^2 & (\theta-1)\sigma_u^2 & \sigma_u^2 & 0 & 0 \\ 0 & (\theta-1)\sigma_u^2 & (\theta-1)^2\sigma_u^2 & (\theta-1)\sigma_u^2 & 0 & 0 \\ 0 & \sigma_u^2 & (\theta-1)\sigma_u^2 & \sigma_u^2 & 0 & 0 \\ \sigma_v^2 & 0 & 0 & 0 & \sigma_v^2 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 \end{bmatrix},$$

respectively.

The state-space form is estimated with the Kalman filter using the consumption series described in table 5. The coefficients λ_1 and λ_2 are not free parameters but instead depend on the consumption persistence coefficient χ: $\lambda_1 = f(\chi), \lambda_2 = g(\chi)$. Our Kalman filter estimation incorporates this relationship between χ, λ_1, and λ_2.

Figures 1 and 2 display the measured consumption growth $\Delta\log C_t$ and true consumption $\Delta\log C_t^*$ estimated using the Kalman smoother based on the above state-space model.

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