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Abstract

We estimate 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. In several countries, the sticky-consumption-growth and Campbell–Mankiw models work about equally well.

Keywords: Sticky Expectations, Consumption Dynamics, Habit Formation

JEL classification: E21, F41

Non-technical Summary

A large literature ranging across macroeconomics, finance, and international economics has argued that 'habit formation' can explain many empirical facts related to consumption dynamics, including stickiness in aggregate consumption growth. Other explanations for the persistence of aggregate spending growth, or 'excess smoothness' (in Campbell and Deaton (1989)'s terminology), include imperfect attentiveness to macroeconomic news on the part of consumers (Carroll and Slacalek (2006); Reis (2006)), or consumers' inability to distinguish micro- from macro-economic shocks (Pischke (1995)). Further explanations could undoubtedly be imagined.

But a full consensus has not yet 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 well known papers by Lettau and Ludvigson (2001, 2004)). Also, in the presence of uncertainty, the precautionary saving 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) rather than actual spending smoothness. 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 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 do not take a stand here on whether such smoothness reflects habits, inattention, or other factors.)

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. Although there is some variation across countries in the 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. Furthermore, wherever there is a clear distinction between the two non-random-walk models, the consumption stickiness 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.)

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, it is about 0.35 in U.S. data). 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 reflects temporary variation that 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).

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.

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, including stickiness in aggregate consumption growth.[1] Other explanations for the persistence of aggregate spending growth, or 'excess smoothness' (in Campbell and Deaton (1989)'s terminology), include imperfect attentiveness to macroeconomic news on the part of consumers (Carroll and Slacalek (2006); Reis (2006)), or consumers' inability to distinguish micro- from macro-economic shocks (Pischke (1995)). Further explanations could undoubtedly be imagined.

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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 systematic estimates of three simple canonical models of con-

[1]Facts that have been interpreted using habit formation models include the equity premium puzzle (Constantinides (1990) and Campbell 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)).

sumption 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 do not take a stand here on whether such smoothness reflects habits, inattention, or other factors.)

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 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. Furthermore, wherever there is a clear distinction between the two non-random-walk models, the consumption stickiness 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, it is about 0.35 in U.S. data). 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 reflects temporary variation that 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

[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.

at the quarterly frequency, cf. Sommer (2007)).[4]

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 (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

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

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

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

where β is the discount factor, C is the consumption level, B is beginning-of-period net assets, R is the constant interest factor, and Y is noncapital 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

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

[5]Carroll and Slacalek (2006) argue that the models can be distinguished using microeconomic data.

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.

Muellbauer (1988) and Dynan (2000) have shown that for a habit-forming consumer with a Constant Relative Risk Aversion (CRRA) outer utility $\mathbf{u}(X) = X^{1-\rho}/(1-\rho)$ and $R\beta = 1$, optimal consumption growth approximately follows an AR(1) process:

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

where ϵ_t denotes innovations to lifetime resources. Hence, in contrast to the standard intertemporally separable utility specification, some of the period t consumption growth is predetermined at time $t-1$. The strength of habits χ can be estimated as the autocorrelation coefficient in the equation for aggregate consumption growth.

2.2 Sticky Expectations

Carroll and Slacalek (2006) present an alternative model of consumer behavior that also generates sticky aggregate consumption growth, but without departing from the standard intertemporally separable utility specification. The key assumption is that consumers are mildly inattentive to macro developments—for example, they do not immediately and fully take into account information contained in aggregate macroeconomic indicators such as productivity growth or the unemployment rate.[6]

Assume for a moment that consumers maximize the discounted sum of time separable quadratic utility streams $-\sum_{t=s}^{\infty} \beta^{t-s}(C_t - \bar{C})^2$ subject to the budget constraint (2). In the standard Hall (1978) model, in which households use all available information, the optimal consumption level follows a random walk and consumption growth is a white noise: $\Delta C_t = \epsilon_t$.

Suppose now instead that the economy consists of a continuum of *inattentive* 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 the consumer last updated his information set and independent of his income or wealth. The model therefore is similar to the Calvo (1983) model of price setting frequently used in the monetary economics literature. Carroll and Slacalek (2006) show that the change in aggregate consumption, ΔC_t, approximately follows an AR(1) process, whose autocorrelation roughly equals the share of consumers $(1-\Pi)$ who

[6] For evidence on this kind of inattention in the context of inflation expectations, see Carroll (2003).

do not have up-to-date information about macroeconomic developments. If utility is of the CRRA form, consumption growth is well approximated by:[7]

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

In addition, in the spirit of Akerlof and Yellen (1985) and Cochrane (1991), Carroll and Slacalek (2006) show that the utility loss from the infrequent updating of expectations is very small.

As noted above, since the two theories of stickiness generate identical implications for aggregate consumption growth dynamics, evidence of a positive serial correlation coefficient will be consistent with either theory. (The theories can be distinguished in other ways, e.g. using microeconomic data; see Carroll and Slacalek for details and evidence.)

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 adopted in the excess sensitivity literature,[8] which has been expanded here to include a term capturing stickiness of consumption growth:

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

where Y denotes household income and A denotes household (net) assets. 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 the capital-to-income ratio A).[9]

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

[8]Early contributions include Flavin (1981), Campbell and Deaton (1989), and Campbell and Mankiw (1989).

[9]By including the assets in the estimated equation, we follow the literature on precautionary savings and liquidity constraints. Calibrated theoretical models imply that the

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 unseasonal weather (or even, for some smaller countries, one-time events such as the hosting of the Olympics). None of the theoretical models include 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.[10]

We develop these points using the United States as an example. The clearest source of measurement error in *quarterly* aggregate consumption data is the services sector, because many components of the quarterly services data are calculated by interpolating from the underlying annual or biennial data sources (Bureau of Economic Analysis (2006)). However, sampling and non-sampling errors introduce significant measurement error even into other categories of consumption data (see Sommer (2007) for details). Sommer also argues that weather-related events can significantly influence aggregate consumption expenditures. For example, 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. 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. All in all, Sommer estimates that measurement error and transitory consumption together account for about 50 percent of the quarterly U.S. nondurables and services consumption volatility, consistent with his empirical finding that the IV estimates of consumption persistence are about twice as high as the OLS estimates.

To address these three estimation issues (measurement error, transitory consumption, and time aggregation) in quarterly consumption data, we use

relationship between A_t and $\mathbf{E}_t[R_{t+1}]$ is almost linear. 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).

[10]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 (Muellbauer (1988), Sommer (2007)), but the MA(2) coefficient is generally small.

two econometric methods developed in Sommer (2007). The first technique attempts to correct for the estimation issues using instrumental variables regressions. As with any IV method, validity of the estimation results depends on the ability to find suitable instruments. As an alternative, the second technique therefore uses the Kalman filter to separate 'true' consumption growth from its transitory components and measurement error.[11] In this case, the usual caveat applies: The validity of this maximum likelihood method hinges on the assumed structure of the stochastic processes for measurement error and 'true' consumption dynamics.

3.1 Sticky Consumption Growth in IV Regressions

3.1.1 Dataset

Equation (5) is estimated using aggregate quarterly data from thirteen advanced economies ranging roughly over the past forty years (table 6 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.[12] Finally, Y and A are measured as household disposable income and the ratio of financial wealth to disposable income, respectively.

[11] 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.

[12] 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. Japan is not included in our sample as creating a quarterly dataset with consumption data 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)). For the U.S., it is possible to perform similar experiments using data on purely nondurable goods spending and on retail sales spending, with results similar to those reported here for PCE excluding durables.

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

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

where the λs are complicated functions of χ. In addition, Sommer verifies that 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).[13]

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.[14] 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.

[13]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.

[14]Consumer 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 widely 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: $V_t^P = \sigma_{t-4,t}^P / \mu_{t-4,t}^P$, where $\sigma_{t-4,t}^P$ is the standard deviation of price level between quarters $t - 4$ and t and $\mu_{t-4,t}^P$ denotes the mean of price level between quarters $t - 4$ and t.

3.1.3 Estimation Results

Table 1 summarizes the baseline estimation results for four alternative econometric specifications nested in equation (5).[15] 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 extends Sommer's (2007) findings for the United States to the international data: the IV estimates of consumption persistence χ are for all countries much higher than would be the OLS estimates and are highly statistically significant. 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 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 four cases insignificant. 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 third 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, where, in addition, the coefficient α has 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 data.[16]

[15] 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. Our parameters could thus be used as an input for prior distributions (as in Del Negro and Schorfheide (2004)).

[16] Microeconomic evidence suggests that the precautionary saving motive may be an important determinant of household-level consumption decisions, see for example Carroll and

The fourth column 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.[17,18]

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 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.[19] 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. The estimation results are broadly consistent with our baseline: (i) the coefficient on lagged consumption growth in univariate re-

Samwick (1997), Gourinchas and Parker (2002), and Fuchs-Schündeln and Schündeln (2005).

[17] 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. 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.

[18] 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.

[19] Germany tends to be an outlier in all our IV regressions (reported and unreported). This may reflect data difficulties associated with comparing pre- and post-reunification German data, or the unpredictable movements in consumption growth during the years following reunification.

gressions 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 (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 follows an MA(1) process.[20] 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 6). 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 coefficients estimated below 0.4. For the United States, the estimated consumption persistence is about 0.7, which is consistent with previous studies.

[20] 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. See Sommer (2007) for a more detailed discussion.

It is encouraging that the Kalman filter estimates of consumption persistence tend to be close to the IV estimates. This indicates that even if instruments such as the lags of consumer sentiment and interest rates happened to be contaminated with some measurement error from the published consumption data, the practical impact on the IV estimates reported in the previous subsection is likely not large. 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), as a result of the combination of measurement error and transitory components.[21] The interesting question is whether the Campbell–Mankiw predicted income term carries any information about 'true' consumption growth beyond the information already contained in consumption persistence. In other words, the question is whether the sticky expectations model is a better model of consumption growth than the rule-of-thumb model after measurement error and transitory consumption have been Kalman-filtered from the data.[22]

Table 5 presents the second-stage IV regression results. "True" consumption growth (as measured by the Kalman smoother) displays little correlation with predicted income, especially when the regressions include a term reflecting sticky expectations (or habit formation). All point estimates are much smaller than the 0.5 estimated by Campbell and Mankiw (1989) and other authors. The estimates of η are mostly statistically insignificant, and none exceeds 0.15. Interestingly, the coefficient on lagged consumption growth changes very little after adding predicted income and wealth to the univariate regression with only consumption growth (compare the left and right panels of the table).

[21] There is an interesting link between the signal-to-noise ratio from the estimated Kalman filter models, $\mathrm{var}(\Delta \log C_t^*)/\,\mathrm{var}(\Delta \log C_t)$, 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.

[22] The figures are interesting in ways that are independent of our analysis in this paper; measurement errors seem to be systematically much larger in some countries than in others, and for several countries appear to have declined markedly over time. These would be interesting topics for future research, but are beyond the scope of our analysis.

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 (three-equation) 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 the classical techniques, including the maximum likelihood estimator. From one point of view, this approach allows the data to have a complete control over the estimates of χ. From another perspective, our estimates of χ could be used as an extra (out-of-sample) information to calibrate priors about consumption sluggishness (or habit persistence) in larger-scale Bayesian DSGE models. In fact, our Kalman filter estimation could also be seen as a special case of the Bayesian framework with uninformative priors.

4 Conclusions

Hall (1978) provided macroeconomists with a clean theoretical benchmark against 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 households (in the range of 10–20 percent) may simply consume their current income every quarter.

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.

Table 1: Consumption Dynamics—All Countries

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

	Estimation with one regressor only				Estimation with all three regressors			
Country	χ	η	α	$\bar{R}_c^2$	χ	η	α	OID
G7 Countries								
Canada†	0.72***	0.32***	0.33	0.17	0.64***	0.05	0.11	0.85
France†	0.61***	0.29***	0.04	0.03	0.44	0.19	−0.04	0.94
Germany†	0.40*	0.72***	−0.36	0.05	0.16	0.66***	−0.17	0.81
Italy†	0.65***	0.20**	−0.05	0.05	0.53**	0.13	−0.02	0.65
United Kingdom†	0.83***	0.10	0.27*	0.12	1.00***	−0.17	0.01	0.95
United States†	0.83***	0.54***	0.26*	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***	0.12	0.10	0.11	0.51**	0.03	0.01	0.63
Belgium‡	0.64***	0.34**	0.11	0.09	0.56**	0.12	0.01	0.72
Denmark‡	0.86***	0.43	−0.34	0.08	0.78***	0.27	−0.32	0.57
Finland‡	0.90***	0.61**	0.53	0.26	0.86***	0.07	−0.13	0.71
Netherlands‡	0.70***	0.09	0.21	0.00	0.53	−0.14	0.10	0.82
Spain‡	0.94***	0.79***	0.83***	0.38	0.71**	0.04	0.21	0.38
Sweden‡	0.83***	0.37**	0.58***	0.16	0.88***	0.32*	−0.25	0.39
Mean Other	0.77***	0.39	0.29	–	0.69***	0.10	−0.05	–

Instruments: `L(2/4).un L(2/4).lr L(2/4).pceinfvol L(2/4).sent`

Notes: Left Panel: Regressions were estimated with one regressor only. Right Panel: Regressions were estimated with all three regressors. Consumption variable: †: nondurables, semidurables and services consumption , ‡: total personal consumption expenditure, A: ratio of household financial wealth to income. $\{^*, ^{**}, ^{***}\}$ = Statistical significance at $\{10, 5, 1\}$ percent (using robust standard errors). $\bar{R}_c^2$: 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)

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

	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)

Instruments: `L(2/4).un L(2/4).lr L(2/4).pceinfvol L(2/4).sent`

Notes: 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.
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—Alternative Instrument Set

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

	Estimation with one regressor only				Estimation with all three regressors			
Country	χ	η	α	$\bar{R}^2_c$	χ	η	α	OID
G7 Countries								
Canada†	0.69***	0.33***	0.91**	0.16	0.40	0.16	0.35	0.36
France†	0.03	0.23*	0.08	0.16	−0.31	0.36	0.09	0.32
Germany†	0.02	0.88***	−0.29	0.07	−0.14	0.89***	−0.18	0.96
Italy†	0.62***	0.29*	−0.06	0.11	0.55***	0.10	−0.02	0.23
United Kingdom†	0.41**	0.07	0.24	0.23	0.58**	−0.20	0.13	0.67
United States†	0.74***	0.41***	0.23	0.17	0.53**	0.16	0.04	0.72
Mean G7	0.42**	0.37***	0.18	–	0.27	0.25	0.07	–
Other Countries								
Australia‡	0.71***	0.18	0.11	0.07	0.73***	−0.05	0.02	0.76
Belgium‡	0.71***	0.27*	0.09	0.09	0.77**	0.13	−0.09	0.69
Denmark‡	0.35	0.10	−1.26**	0.07	0.19	−0.00	−1.14*	0.53
Finland‡	0.88***	0.49***	2.86***	0.22	0.56*	0.15	0.82	0.66
Netherlands‡	0.75***	0.16	0.12	0.11	0.71***	0.14	0.02	0.82
Spain‡	0.94***	0.58***	0.97***	0.38	0.67**	0.18	0.15	0.86
Sweden‡	0.86***	0.05	0.81***	0.20	0.85**	−0.03	0.04	0.36
Mean Other	0.74***	0.26	0.53	–	0.64**	0.07	−0.03	–

Instruments: `L(2/4).un L(2/4).dinc L(2/4).irSpread L(2/4).sent`

Notes: Left Panel: Regressions were estimated with one regressor only. Right Panel: Regressions were estimated with all three regressors. Consumption variable: †: nondurables, semidurables and services consumption , ‡: 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

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

	Parameter Estimates				
Country	χ	θ	$\log \sigma_u^2$	$\log \sigma_v^2$	$\frac{\text{var}(\Delta \log \mathbf{C}_t^*)}{\text{var}(\Delta \log \mathbf{C}_t)}$
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: †: nondurables, semidurables and services consumption, ‡: total personal consumption expenditure. {*, **, ***} = Statistical significance at {10, 5, 1} percent. Samples for the Kalman filter estimation are as follows: Canada Q2:1961–Q4:2006, France Q2:1978–Q4:2006, Germany Q2:1970–Q4:2006, Italy Q2:1981–Q3:2006, United Kingdom Q2:1964–Q3:2006, United States Q2:1960–Q4:2006, Australia Q3:1965–Q4:2003, Belgium, Denmark, Finland and Sweden Q2:1961–Q2:2003, Netherlands Q2:1961–Q2:2004, and Spain Q3:1962–Q2:2003.

Table 5: Consumption Dynamics—Second-Stage Kalman Filter Estimates

$$\Delta \log C_t^* = \varsigma + \chi \mathbf{E}_{t-2}[\Delta \log C_{t-1}^*] + \eta \mathbf{E}_{t-2}[\Delta \log Y_t] + \alpha \mathbf{E}_{t-2}[A_{t-1}]$$

	Estimation with one regressor only				Estimation with all three regressors			
Country	χ	η	α	$\bar{R}_c^2$	χ	η	α	OID
G7 Countries								
Canada†	0.92***	0.32***	0.28	0.56	0.91***	0.00	0.02	0.50
France†	0.91***	0.21***	0.08	0.53	0.84***	0.05**	−0.00	0.58
Germany†	0.92***	0.34***	−0.33	0.46	0.81***	0.12***	−0.00	0.74
Italy†	0.84***	0.16*	−0.04	0.19	0.82***	0.01	−0.01	0.38
United Kingdom†	0.89***	0.14*	0.24*	0.30	1.00***	−0.08	−0.00	0.88
United States†	0.89***	0.43***	0.24**	0.31	0.72***	0.15**	0.01	0.60
Mean G7	0.90***	0.27***	0.08	–	0.85***	0.04	0.00	–
Other Countries								
Australia‡	0.76***	0.11*	0.11*	0.34	0.73***	0.03	0.00	0.32
Belgium‡	0.80***	0.35***	0.12	0.31	0.82***	−0.02	−0.00	0.46
Denmark‡	0.97***	0.31	−0.29	0.29	0.92***	0.20	−0.32*	0.69
Finland‡	0.91***	0.56***	0.47	0.49	0.96***	−0.04	−0.10	0.73
Netherlands‡	0.93***	0.11	0.22**	0.28	0.92***	0.01	0.00	0.22
Spain‡	0.99***	0.77***	0.79***	0.48	0.97***	−0.03	0.05	0.32
Sweden‡	0.89***	0.26*	0.51***	0.25	0.96***	0.14	−0.17	0.20
Mean Other	0.89***	0.35**	0.27*	–	0.90***	0.04	−0.08	–

Instruments: `L(2/4).un L(2/4).lr L(2/4).pceinfvol L(2/4).sent`

Notes: Left Panel: Regressions were estimated with one regressor only. Right Panel: Regressions were estimated with all three regressors. Consumption variable: †: true nondurables, semidurables and services consumption from the Kalman smoother, ‡: true total personal consumption expenditure from the Kalman smoother, A: ratio of household financial wealth to income. $\{^*, ^{**}, ^{***}\}$ = Statistical significance at $\{10, 5, 1\}$ percent (using robust standard errors). $\bar{R}_c^2$: Adjusted R^2 from the first-stage regression of consumption growth on instruments. OID: p-value from the Hansen's J statistic for overidentification.

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

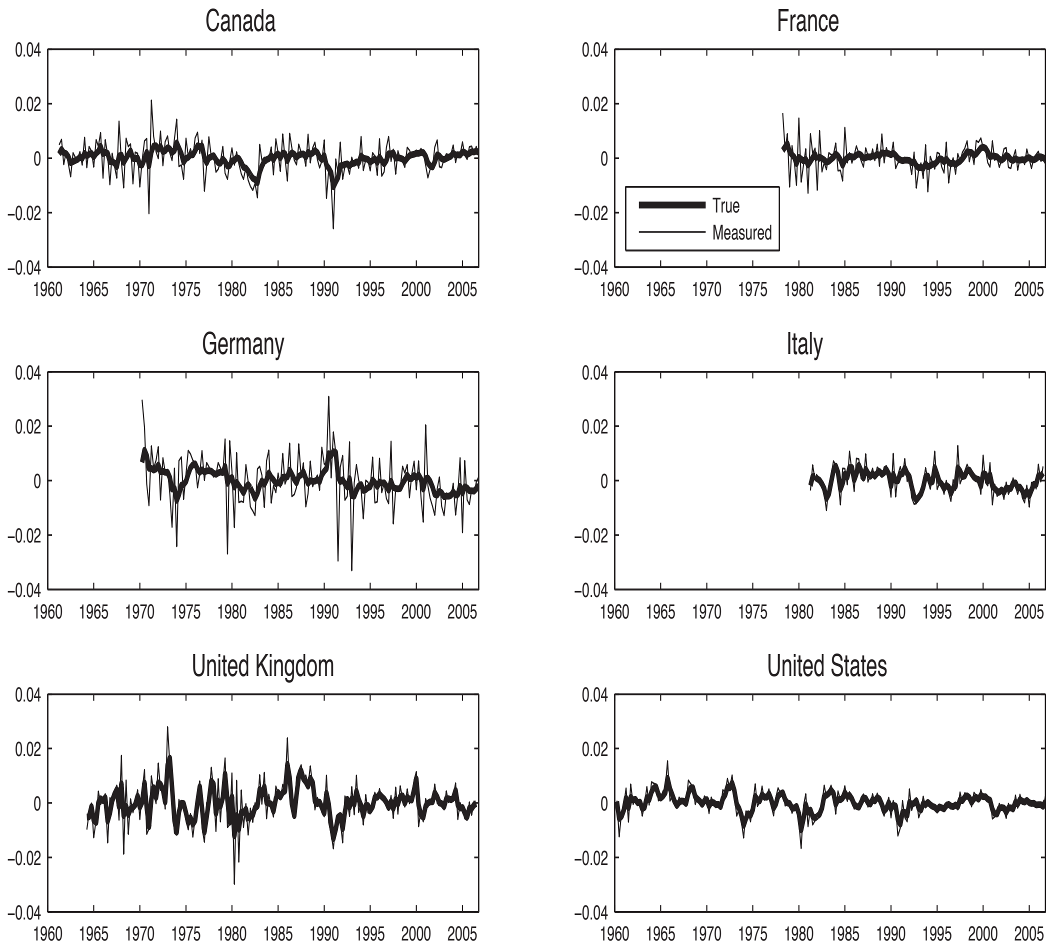
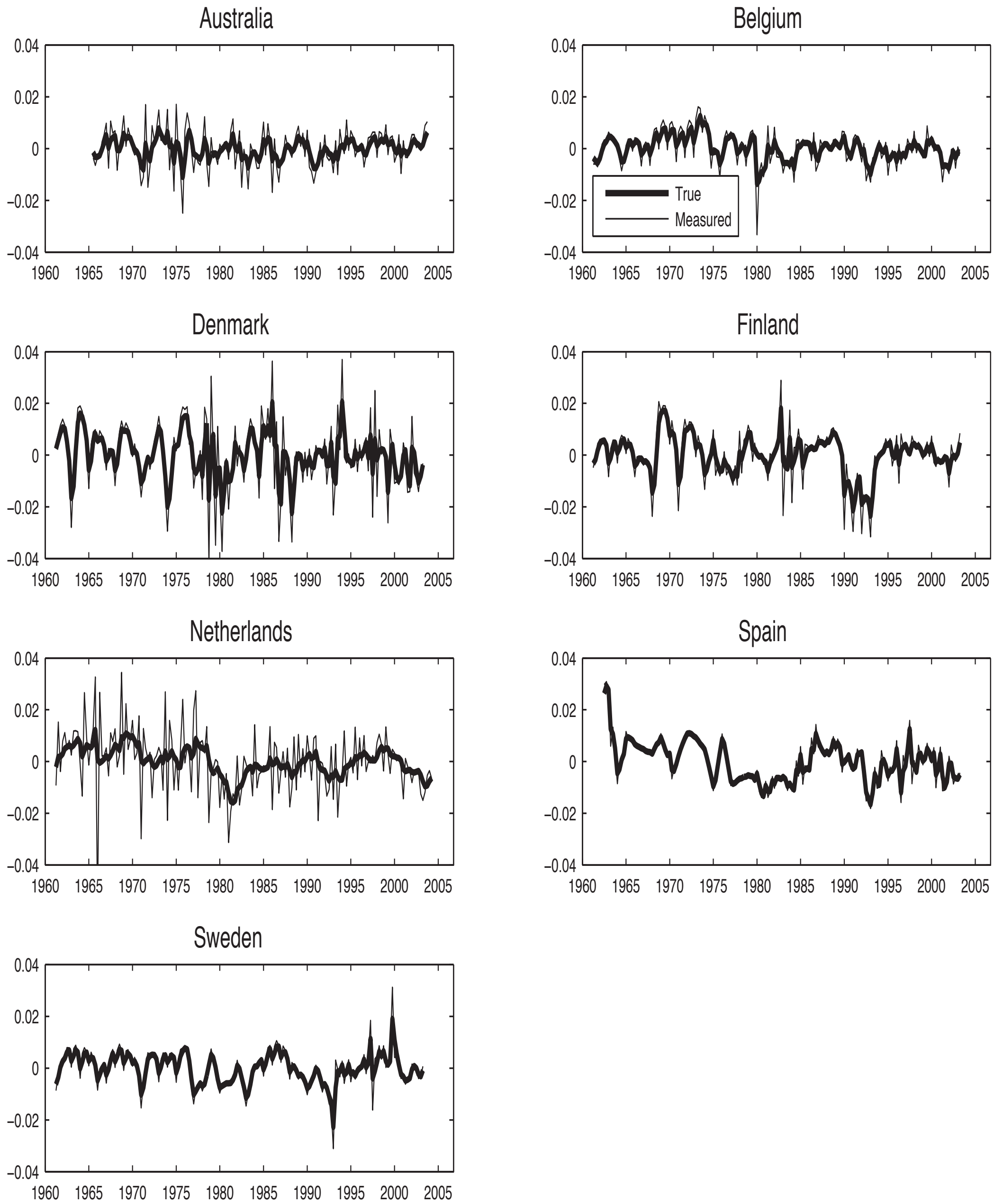


Figure 2: Measured and "True" Consumption Growth—Other Countries



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 6: Consumption Data, Its Sources, and Samples for IV Regressions

Country	Time Frame	Consumption Series	Source
G7 Countries			
Canada	Q4:1970–Q3:2002	Nondurables + Services	Haver Analytics
France	Q1:1985–Q4:2003	Nondurables + Services	Haver Analytics
Germany†	Q4:1975–Q4:2002	Nondurables + Services	Haver Analytics
Italy	Q1:1981–Q4:2003	Nondurables + Services	Haver Analytics
United Kingdom	Q1:1974–Q4:2003	Nondurables + Services	Haver Analytics
United States	Q3:1962–Q2:2004	Nondurables + Services	Haver Analytics
Other Countries			
Australia	Q4:1975–Q4:1999	Total PCE	Haver Analytics
Belgium	Q2:1980–Q4:2002	Total PCE	NiGEM/MEI
Denmark	Q1:1977–Q2:2003	Total PCE	NiGEM/MEI
Finland	Q3:1973–Q2:2003	Total PCE	NiGEM/MEI
Netherlands	Q1:1975–Q4:2002	Total PCE	NiGEM/MEI
Spain	Q1:1978–Q4:1999	Total PCE	NiGEM/MEI
Sweden	Q1:1977–Q4:2002	Total PCE	NiGEM/MEI

Notes: †: Regressions for Germany were estimated with a reunification dummy in Q1:1991; Source: 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 6. The coefficients λ_1 and λ_2 are not free parameters but instead depend on the consumption persistence coefficient χ: $\lambda_1 = f(\chi), \lambda_2 = g(\chi)$ as detailed in the appendix to Sommer (2007). 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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