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ABSTRACT

Consumption and Aggregate Constraints: Evidence from US States and Canadian Provinces*

State-level consumption exhibits excess sensitivity to lagged income to the same extent as US aggregate data, but state-specific (idiosyncratic) consumption exhibits substantially less sensitivity to lagged state-specific income – a result that also holds for Canadian Provinces. We propose the following interpretation: borrowing and lending in response to changes in consumer demand is easier for an individual US state than it is for the US as a whole. The PIH may thus be a good model for describing the reaction of consumption to idiosyncratic disposable income shocks even if it fails at the aggregate US level. Further analysis, centred on the persistence of income shocks and on the consumption/income ratio, is consistent with this interpretation but suggests that the PIH still require qualification. We contrast our results with tests of full interstate risk sharing.

JEL Classification: E21
Keywords: Canadian provinces, consumption, excess sensitivity, excess smoothness, permanent income, regional macroeconomics, risk sharing and US states

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1 Introduction

Personal consumption in the United States amounts to seventy percent of Gross Domestic Product and the modeling of consumer behavior is still a challenge to the profession in spite of much serious research. Hall's (1978) version of the Permanent Income Hypothesis (PIH) implies that current consumption is independent of lagged disposable income conditional on lagged consumption. Micro evidence for this proposition is mixed while macro evidence overwhelmingly rejects it, resulting in an empirical stylized fact—the excess sensitivity of consumption to lagged income.[1] The model also has implications for the relative volatility of consumption and income. In most empirical studies, consumption does not respond sufficiently strongly to income innovations, resulting in a second stylized fact—the excess smoothness of consumption.[2]

Hansen and Singleton (1982, 1983) developed and tested the empirical implications of the PIH when asset returns and, in particular, interest rates are time-varying and stochastic, but their tests have not been successful in fitting the model to US macroeconomic time-series.[3] Micro studies allowing for a time-varying interest rate have been more favorable to the PIH. For example, Altonji and Siow (1987), who include time dummies in their regressions, and Mariger and Shaw (1993), who allow for time varying coefficients, find little or no excess sensitivity of consumption to lagged income.

One reason why the PIH performs less well with aggregate time-series data may be that credit market and trade frictions are smaller within the US so that borrowing and lending in response to changes in consumer demand is easier for an individual US state than it is for the US as a whole. This interpretation is consistent with the following "general equilibrium effect" suggested by Michener (1984).[4] When consumers in a fairly closed economy such as the US wish to increase the share of National Income devoted to consumption there is increased

[1] The basic reference for analyses of excess sensitivity based on Hall's (1978) version of the PIH and using national level aggregate data is Flavin (1981), while Hall and Mishkin (1982) is the basic reference for micro studies.

[2] The implications of the PIH model for the relative volatility of consumption and income were derived by Hansen, Roberds, and Sargent (1991) who argue that the high persistence of income implies that consumption should respond strongly to income innovations. Yet in reality, as shown by Deaton (1987), West (1988), and Campbell and Deaton (1989), consumption is excessively smooth. There is by now a large literature explaining these seeming deviations from optimal consumer behavior. See Deaton (1992) for a comprehensive survey, and the discussion in Section 4.

[3] Other tests of the time-varying interest rate PIH model include Mankiw (1981) and Shapiro (1984) who also reject the model. Hansen and Singleton's (1982, 1983) analysis allows for more general utility functions, and typically involves estimation of non-linear first order conditions. In that tradition, excess sensitivity tests are commonly referred to as tests of orthogonality conditions.

[4] Christiano (1987) makes a similar point.

competition for scarce resources, since aggregate consumption cannot adjust immediately.[5] This creates upward pressure on the US-wide interest rate depressing the demand for consumption. Therefore, in the aggregate, we should indeed expect substantial deviations from Hall's *constant interest rate* benchmark PIH model.

This intuition also suggests that if movements in the interest rate (and asset returns in general) are explicitly controlled for in empirical tests of the PIH, the model should perform better. Yet, the literature following Hansen and Singleton (1982, 1983) has not been successful in fitting the variable interest rate PIH model to macroeconomic data. This may be due to heterogeneity in the population (different people face different interest rates) or to an inherent difficulty in correctly measuring the relevant "market clearing" interest rate (and the relevant equilibrium asset returns more generally).

We suggest a way around this problem. If the failure of the PIH model is partly due to the "general equilibrium effect" described above and if asset returns do not capture this effect well, we may still expect the model to perform better with *idiosyncratic* (state-specific) consumption and income than with US-wide aggregate consumption and income. The logic is simple: it is easy for a state to borrow and lend in the event of a state-specific change in consumption demand, even if it may be hard for the US as a whole to borrow and lend in response to a US-wide change in consumption demand. Moreover, changes in state-specific consumption demand do not affect the US-wide equilibrium interest rate since state-specific shocks sum to zero by definition in any given year.

To corroborate this conjecture empirically, we examine implications of the PIH using data on personal disposable income and consumption for US states and Canadian provinces. Regional data at the sub-national level are much underutilized for the study of consumer behavior. Such data are sufficiently aggregated to be regarded as macroeconomic data, yet exhibit considerable cross-sectional variation that can be exploited in empirical analysis. Endogeneity of state-specific income is not likely to be a major problem and measurement error is less serious than in micro data.

We perform standard tests using state-level consumption and disposable income and find considerable excess sensitivity, similar to that found using aggregate US data.[6] We then remove the aggregate US-wide component in the data and find that state-specific consumption exhibits substantially less sensitivity to lagged state-specific disposable income. Thus, once movements

[5]For example, because it takes time to increase the quantity of goods imported. In Section 3, we discuss potential mechanisms that may cause aggregate consumption to adjust slowly.

[6]In most of the discussion, we will refer only to US states. The empirical results for Canadian provinces are very similar.

in aggregate income and consumption are controlled for, the deviation from optimal consumption behavior in macroeconomic data is not as large. We conclude that it may be premature to discount the "general equilibrium PIH model," since the model describes the reaction of idiosyncratic consumption to idiosyncratic disposable income reasonably well.

The persistence of income shocks differs across states and we find that excess sensitivity of consumption increases with the persistence of income shocks. We also check whether the lagged ratio of consumption to disposable income predicts current consumption growth.[7] Finally, we briefly explore the parallel between tests of the PIH and tests of optimal risk sharing.[8]

The next section is devoted to a description of the statistical properties of US state-level disposable income and consumption series. In Section 3, we study excess sensitivity of consumption and in Section 4, we perform further tests. Section 5 is devoted to the analysis of Canadian province-level data and Section 6 concludes.

2 US State-Level Data

We use annual data for 1963–1995. Disposable personal income data are from the Bureau of Economic Analysis (BEA). We approximate state-level private non-durable consumption by state-level retail sales of non-durable goods.[9] We transform the data series to per capita terms using population data from the BEA.[10]

Let $y_{it} = Y_{it} - Y_t$ denote state i's period t idiosyncratic (state-specific) disposable log-income per capita, where Y_{it} is its period t (total) disposable log-income per capita, and Y_t is the period t aggregate (US-wide) disposable log-income per capita. We assume that $\mathrm{Cov}(y_{it}, Y_t) = 0$ and $\Sigma_i y_{it} = 0$.[11]

[7] This type of regression was used by Cochrane (1994) in the framework of a bivariate system to predict US aggregate income (more precisely, GNP) and to identify the transitory component in income.

[8] For empirical tests of full risk sharing (and deviations thereof) see, e.g., Cochrane (1991), Mace (1991), Townsend (1994), Obstfeld (1994), Asdrubali, Sørensen, and Yosha (1996), and Sørensen and Yosha (1998).

[9] Retail sales by state are published in the Survey of Buying Power in Sales Management (after 1976, Sales & Marketing Management). These data are proprietary and we thank the publishers of Sales & Marketing Management for permission to use the series. We thank Marco del Negro for providing us these data, sub-divided into consumption categories, in electronic-readable form. Retail sales is a somewhat noisy proxy for state private consumption (e.g. travel expenses are not included in retail sales) but, to our knowledge, it is the best available. The correlation between annual percentage increments of aggregate US non-durable retail sales and aggregate US non-durable private consumption (in the NIPA data), both measured in real (cpi deflated) terms, is 0.68.

[10] We will often refer to state per capita personal disposable income as "income" or "disposable income."

[11] We can think of the idiosyncratic components of income as representing time-varying income shares. For instance, we can write the level of period t (total) state disposable income as $Y^*_{it} = y^*_{it} Y^*_t$ where Y^*_t is aggregate (US-wide) per capita disposable income and y^*_{it} are idiosyncratic (state-specific) shares such that $\Sigma_i y^*_{it} = 1$. Taking logs, this yields $y_{it} = Y_{it} - Y_t$ with $\Sigma_i y_{it} \approx 0$. In reality, the dynamics of state-level disposable income may be more complex. In particular, a common shock to the US economy (e.g., an oil price shock) might have

It is widely accepted that US aggregate income series are non-stationary. By contrast, the statistical properties of the *idiosyncratic* components of US state-level data have not been studied. Exploiting the panel structure of our data, we perform the Im, Pesaran, and Shin (1997) (IPS) test for a unit root in y_{it}. The null hypothesis of non-stationarity is not rejected with 1, 2, and 3 lags (with P-values of 0.13, 0.45, and 0.34 respectively). The IPS test is valid for independent observations and since the idiosyncratic components of income are unlikely to be fully independent, the critical values of the test-statistics must be taken as approximations.[12] State-by-state Augmented Dickey-Fuller (ADF) tests for unit roots reject the unit root null hypothesis for only a few states, at conventional levels of significance. ADF tests provide somewhat weak evidence, since they have low power for samples as short as ours. The overall impression is, nevertheless, that the idiosyncratic component of US state-level disposable income is well described as an integrated process.

We define idiosyncratic state-level per capita consumption in the same manner: $c_{it} = C_{it} - C_t$ (where the variables are expressed in logs). The aggregate per capita non-durable retail sales series, C_t, is clearly non-stationary according to standard ADF tests. As for the income series, the IPS test may only be approximately valid, but together with ADF tests it can provide a reasonable guide to specification. We find that the null hypothesis of non-stationarity is not rejected for a specification with 1 lag (with a P-value of 0.10) but is rejected with 2 and 3 lags (with P-values of 0.01 and 0.02 respectively). State-by-state ADF tests rarely reject the hypothesis of non-stationarity, so we conclude that the idiosyncratic component of consumption is best regarded as non-stationary.

Most models of consumption imply that consumption tracks income in the long run. An interpretation of this is that the process $(c_{it} - y_{it})$ is stationary; i.e., that consumption and income are cointegrated series with a coefficient of unity. We can test this hypothesis simply by performing the IPS test. Such a test consistently rejects the null of a unit root, for various lag lengths and whether a drift term is allowed or not. We, therefore, feel confident treating $(c_{it} - y_{it})$ as a stationary process.[13]

Since both y_{it} and c_{it} are best regarded as non-stationary, we carry out the empirical analysis using first-differenced series. We estimated AR(2) models for these series and found the coeffi-

different dynamic effects on income in different states (e.g., agricultural versus industrial states). Such dynamics would not be captured by our statistical model, but might be captured by more complicated factor analysis models like the one suggested by Forni and Reichlin (1997).

[12] When the aggregate component is not subtracted, the income of different states typically displays a clear positive correlation and the IPS test is not valid. (Applying the test to state-level disposable income indeed produces meaningless results, rejecting the null of a unit root in favor of an explosive alternative.)

[13] A similar result holds for $C_{it} - Y_{it}$, when a trend is allowed for.

cients of the twice lagged variables very small and typically insignificant.[14] Moreover, a formal test of the hypothesis that the AR(2) coefficients are all zero provided no evidence against the null, so a simple AR(1) model in log-differences seems appropriate. Table I displays AR(1) models for state-level total and idiosyncratic disposable income and consumption. Each model is estimated using feasible GLS allowing for cross-correlations of the disturbances between states.[15] For state-level disposable log-income per capita, ΔY_{it}, the estimated average of the AR(1) coefficients is 0.16 with the absolute value of the t-statistics averaging at 2.46, and the hypothesis that the AR(1) coefficients are equal across states is strongly rejected. State-level idiosyncratic disposable log-income per capita, $\Delta Y_{it} - \Delta Y_t$, exhibits similar properties with an average AR(1) coefficient of 0.05, a 2.65 average for the absolute value of the t-statistics, and rejection of the hypothesis that the coefficients are equal across states. By contrast, the AR(1) coefficients for state-level consumption and idiosyncratic consumption are negative, –0.06 and –0.08 respectively, although not strongly significant (the average absolute values of the t-statistics are 1.47 and 1.88). The negative autocorrelation may be due to "classical" measurement error in the retail sales series.[16] In all four regressions displayed in Table I, the null hypothesis $\phi_i = \phi$ for all i is strongly rejected with P-values of 0.00.

To get a sense of how "wild" the variation in these series might be, we display the idiosyncratic series graphically in Figure 1 for five states and the aggregate series for the US.[17] The retail sales series show more variation than the disposable income series, with average (across the 50 states) standard deviations $\sigma_{\Delta c_{it}} = 4.19$ and $\sigma_{\Delta y_{it}} = 2.31$. For the series ΔC_{it} and ΔY_{it}, we have $\sigma_{\Delta C_{it}} = 3.12$ and $\sigma_{\Delta Y_{it}} = 5.01$. The high variance of the non-durable retail sales series, and the negative autocorrelation reported previously, are both consistent with measurement error. There is, however, nothing that indicates that the amount of measurement error in the income series varies much between the state-level data and the idiosyncratic state-level data.[18]

[14]We do not provide detailed tables for the AR(2) estimations.

[15]We estimated an unrestricted variance-covariance matrix for the 50 states based on the residuals from an initial panel data OLS estimation. Since we have less than 50 time-series observations for each state, this estimated variance-covariance matrix is singular and in order to perform the second stage GLS estimation we modified the estimated variance-covariance matrix by boosting the diagonal elements by 10 percent. The estimated coefficients were very similar to those obtained from the OLS regressions. The estimated standard errors depend on the procedure used, but we verified empirically that the qualitative conclusions of the present paper hold even if covariances across states are set equal to 0, as long as variances are allowed to differ across states.

[16]By which we mean measurement error that is independently distributed over time and across individuals.

[17]We chose five states at random (states 10, 20, 30, 40, 50 in alphabetical order).

[18]Greater measurement error in the idiosyncratic income data would invalidate our conclusions from the empirical findings reported in the next section.

3 Sensitivity of State- and Province-Level Consumption to Lagged Income

We turn to our central empirical question: is excess sensitivity lower when aggregate fluctuations are controlled for? In the regressions, we control for such fluctuations in three different manners: (1) by regressing the idiosyncratic consumption growth, Δc_{it}, on lagged idiosyncratic income growth, $\Delta y_{i,t-1}$; (2) by including time dummy variables (time fixed effects) in the regressions of ΔC_{it} on $\Delta Y_{i,t-1}$; and (3) by including instead aggregate consumption as a regressor. In all three cases we expect a lower excess sensitivity coefficient compared to the coefficient in the regression of total consumption, ΔC_{it}, on total lagged income, $\Delta Y_{i,t-1}$, with no time fixed effects.

The empirical results displayed in Table II confirm this prediction. Without controlling for aggregate fluctuations the coefficient of lagged disposable income is 0.23 and is highly significant. When aggregate fluctuations are controlled for by regressing $\Delta C_{it} - \Delta C_t$ on $\Delta Y_{i,t-1} - \Delta Y_{t-1}$, the coefficient of lagged income falls drastically by about half. This result is very robust to alternative ways of "controlling for the aggregate": the coefficient to lagged income is virtually unchanged if instead of subtracting aggregate variables we include time fixed effects, or include aggregate consumption as a regressor (with or without state-specific coefficients for this regressor).[19]

In Section 5, we report similar results for Canadian provinces using province-level national accounts data. Similar results (details not tabulated) hold for US states using first-differenced levels (rather than logs) of disposable income and consumption, as well as using total (durable and non-durable) retail sales data. A clear and robust empirical regularity emerges: controlling for aggregate fluctuations and focusing on the reaction of idiosyncratic consumption to idiosyncratic disposable income dramatically reduces the excess sensitivity of consumption.

Interpretation and discussion

Our preferred interpretation of this finding relies on the "closedness" of the US economy. There are (at least) two ways of thinking about "closedness." The first stresses frictions and imperfections in international capital and credit markets rendering international borrowing and lending difficult and preventing rapid adjustment of aggregate consumption to US-wide changes in consumption demand. By contrast, individual states are relatively open in the sense that

[19] We further regressed ΔC_{it} on $\Delta Y_{i,t-1}$ with a separate coefficient for each state and obtained a similar drop in the average value of these coefficients when time fixed effects are included. (The hypothesis that these coefficients are the same for all states is clearly rejected both with and without time fixed effects.) The detailed results of the regression are not reported in the tables. In all the regressions we include a dummy variable for each state (a state fixed effect). The results are not affected substantially when these dummy variables are omitted.

they can more easily borrow and lend among themselves. Thus, the adjustment of state-specific consumption to changes in state-specific consumption demand should be faster.

An alternative manner of thinking about the "closedness" of the US economy is centered on the slow adjustment of US net imports in response to fluctuations in US consumption demand. In a fully integrated and frictionless world, aggregate net imports would immediately increase in response to higher consumption demand. In reality, it may take time to adjust aggregate imports (not to speak of exports). For example, an increased demand for Toyota cars in the US will typically be reflected in higher prices (no "dealer incentives") and less attractive financing opportunities, since adjustment of Japanese exports can not be done instantaneously. By contrast, net imports of a state within the US can adjust much more rapidly. If, in some year, Massachusetts residents have a large idiosyncratic demand for consumption, this demand may be satisfied relatively quickly by moving goods from other states where idiosyncratic demand is low (recall that the state-specific components of income and consumption sum to zero each year, by definition).

These economic mechanisms may be independent or complementary (e.g., imports adjust slowly *because* international credit markets are imperfect) and we do not have adequate data to disentangle them. Our empirical results strongly suggest that such mechanisms are part of the explanation for the seeming deviations from optimal consumer behavior in macroeconomic data.[20]

The log-linear consumption function provides a simple framework for summarizing this discussion. As is well known, if utility is isoelastic with intertemporal elasticity of substitution given by a parameter ρ then, subject to a log-linear approximation and assuming that r_t (the country-wide interest rate on savings made at t) is known with certainty at t, state i's log-consumption follows the process $\Delta C_{it} = \alpha_i + \rho r_{t-1} + e_{it}$ where $E_{t-1}\, e_{it} = 0$.[21]

If the country is a "small open economy," in the sense that it can trade freely with other countries as well as borrow and lend internationally without frictions taking world prices as given, r_t is best interpreted as the world interest rate which is independent of shocks to the income of the country or any of its "states." Therefore, for any state i within the country, the covariance of r_{t-1} and $\Delta Y_{i,t-1}$ is zero, and a regression of ΔC_{it} on $\Delta Y_{i,t-1}$ will give a coefficient of zero.[22]

If the country is not a "small open economy" in the above sense (the US in our case),

[20] Attfield, Demery, and Duck (1992) demonstrate that adjustment costs in consumption may explain deviations from PIH behavior.

[21] For details consult Deaton (1992), p.64.

[22] This is, of course, the excess sensitivity test proposed by Hall (1978).

adjustment of the aggregate consumption in response to shocks to aggregate income may take time. This is so because a positive aggregate shock to the country's income in period $t-1$ will typically induce all states to increase their demand for current (period $t-1$) consumption by more than the rise in current income.[23] In a country which does not satisfy the assumptions of a "small open economy," this demand for consumption may not be satisfied fully due to the aggregate constraint on consumption. This aggregate constraint will exert upward pressure on the country-wide interest rate which will depress consumption and restore equilibrium. This will result in a positive correlation between consumption growth, ΔC_{it}, and lagged income growth, $\Delta Y_{i,t-1}$. If this effect is controlled for in regressions of ΔC_{it} on $\Delta Y_{i,t-1}$ the coefficient should be zero (no excess sensitivity).

Controlling for the aggregate constraint in the framework of the log-linear consumption model can be done as follows. Suppose that aggregate log-consumption follows the process $\Delta C_t = \alpha + \rho r_{t-1} + u_t$.[24] Subtracting this equation from that of state i's log-consumption will result in the regression $\Delta C_{it} - \Delta C_t = (\alpha_i - \alpha) + (e_{it} - u_t)$, namely, the country-wide interest rate washes out and a regression of $\Delta C_{it} - \Delta C_t$ on $\Delta Y_{i,t-1} - \Delta Y_{t-1}$ should yield a zero coefficient.

This regression allows us to circumvent the problem of how to measure the prevailing equilibrium interest rate. In practice, measured interest rates are affected by many factors such as monetary or fiscal policy that are typically not incorporated in theoretical models and in empirical analyses of consumption. Moreover, consumers are often unable to obtain credit at posted interest rates. Our strategy is to avoid directly addressing these (important) issues and to focus on one key point: controlling for aggregate constraints improves the empirical performance of the PIH. The results in Table II indeed indicate that doing so substantially reduces excess sensitivity, but does not eliminate it. The remaining excess sensitivity may be due to other frictions that have been extensively researched.[25]

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[23] According to the point estimates for the AR(1) model for income reported in Table 1 a positive income shock is likely followed by another (smaller) positive income shock (the size of the unit root in disposable income is larger than one in the language of Cochrane 1988). This implies that permanent income and, hence, consumption demand will increase more than income; see Deaton (1992), p.108.

[24] This is an approximation, but for illustrative purposes this formulation is sufficient. In our regressions we alternatively subtracted average consumption growth (this corresponds to regressions with time fixed effects).

[25] Hall and Mishkin (1982) and Zeldes (1989) stress credit rationing while Campbell and Mankiw (1990) emphasize the presence of rule-of-thumb consumers. Heaton (1993) emphasizes intertemporal non-separabilities such as durability of consumption or habit persistence in preferences. Gali (1990) and Clarida (1991) suggest that aggregation over individuals with finite horizons (due to retirement and finite lifetimes) may explain excess sensitivity even if all individuals satisfy the life cycle model. Pischke (1995) argues that deviations from the PIH may be due to consumers not separating between transitory idiosyncratic and permanent aggregate income shocks, while Attanasio and Weber (1995) emphasize aggregation across households and failure to control for demographic and labor supply variables in macro studies (as well as non-separabilities in consumption). Kuznits (2000) explores the implications of direct utility from wealth for optimal consumer behavior with very encouraging results. Others

Excess sensitivity of consumption and the persistence in disposable income

So far, we have concentrated on average (across US states) excess sensitivity ignoring potential heterogeneity in state-level patterns of income and consumption. The results in Table I, however, indicate that the AR(1) processes for state-level disposable income growth are not identical. A larger AR(1) coefficient in the income process implies a larger size of the unit root, to which we will (loosely) refer as more persistent income shocks. We address the possibility that excess sensitivity differs with income persistence. We measure persistence by the coefficient ϕ_i in the regression $\Delta Y_{it} - \Delta Y_t = \alpha_i + \phi_i(\Delta Y_{i,t-1} - \Delta Y_{t-1}) + \epsilon_{it}$, and sort states into three groups according to this measure.[26]

Table III reveals a potentially important stylized fact. Excess sensitivity is systematically larger the higher the persistence of disposable income. Several models are consistent with this result. For example, Keynesian models where consumers consume a fixed fraction of their current income or "rule-of-thumb" models (Hall and Mishkin 1982 and Campbell and Mankiw 1990) where a fixed fraction of consumers consume all their income are consistent with this pattern, since current income growth is more correlated with lagged income growth the higher the persistence (i.e., the higher the AR(1) coefficient). Therefore, a "mis-specified" consumption function that uses lagged (rather than current) income is closer to the "true" specification the more persistent is income. It may also be that states themselves do not satisfy the assumptions of the "small open economy model" perfectly. In this case, states with more persistent income shocks (where permanent income differs relatively more from actual income) may face larger problems satisfying consumption demand.

Models of precautionary saving stress the importance of the variance of income as a determinant of consumption patterns. We verified that in our data, persistence is not just a proxy for variance (actually income persistence and variance are slightly negatively correlated in our sample). We will not pursue further the implications of heterogeneity across states, but utilizing state-level differences in order to sort through the many models of consumer behavior seems a fruitful area for research.

Turning to the bottom panel of Table III, when aggregate fluctuations are controlled for—either by removing the US-wide component of consumption and income from the data, or by

emphasize methodological aspects of empirical research. Quah (1990), for example, argues that excess smoothness may be an artifact of consumers separating temporary from persistent shocks but the econometrician does not have enough information to do so, while Christiano, Eichenbaum, and Marshall (1991) stress time aggregation biases. Of related interest are papers that estimate the intertemporal elasticity of substitution in consumption: Hall (1988) and Campbell and Mankiw (1989) use aggregate US data while Beaudry and van Wincoop (1996), using a panel of US state-level data.

[26] The Campbell and Mankiw (1987) measure of persistence for this model is simply $1/(1-\phi_i)$.

including time fixed effects in the regression—excess sensitivity of consumption decreases within each persistence group.

4 Further Results

Richer dynamics

Our basic result in Table II is robust to the inclusion of a second lag of disposable income in the excess sensitivity regression, as well as an "error correction term" as in Cochrane (1994); see Table IV. In both cases, the coefficient of one-year lagged income is smaller when aggregate fluctuations are controlled for. In the specification with an error correction term, the coefficient of two-year lagged income is also smaller if aggregate fluctuations are controlled for.

The coefficient of the error correction term is negative and significant suggesting that, indeed, this specification captures important dynamics in the data. The negative coefficient is considerably larger and more strongly significant when aggregate fluctuations are controlled for. The high t-statistic associated with this coefficient (especially in the regression that controls for aggregate fluctuations) implies that the lagged income/consumption ratio helps forecast consumption growth contrary to the predictions of the PIH. This result may be another piece of evidence that the "general equilibrium (aggregate constraints)" PIH model still requires qualification; but it might also be a result of measurement error in the consumption data.

Excess smoothness of consumption

We want to briefly address the issue of whether controlling for aggregate fluctuations also helps explain the excess smoothness of consumption; see Campbell and Deaton (1989), Hansen, Roberds, and Sargent (1991), and Gali (1991).[27] A commonly used method is to compare the ratio of the standard deviations of the consumption and income time-series to the corresponding ratio predicted by the underlying theory for given stochastic properties of the time-series for (disposable) income; see, e.g., Pischke (1995). Such an approach is not valid if any of the series are subject to measurement error, as those will drive up the estimated standard deviation. In our case, as previously reported, the standard deviation of state-specific US retail sales is higher than the standard deviation of state-specific disposable income (i.e., there is no evidence of excess smoothness in these data), but this is likely due to higher measurement error in the retail sales data. Moreover, for Canadian province-level data (where measurement error in the

[27] Excess smoothness and excess sensitivity are closely related so any explanation of excess sensitivity is likely to help explain excess smoothness.

consumption data is probably smaller) the standard deviation of province-level consumption is lower than the standard deviation of province-level disposable income. No conclusions can thus be drawn regarding excess smoothness from a comparison of the standard deviation of the income and consumption series.

Even if we disregard measurement error, tests for excess smoothness are more delicate than tests for excess sensitivity, since it is necessary to correctly identity consumers' information sets in order to determine the amount of new information about future income that agents obtain in each period. Simple tests for excess smoothness assume that agents estimate their permanent income based on current and past income; see Deaton (1992), chapter 4. We will outline the simplest version of such a test using our data. Assume that state-level idiosyncratic disposable income is a random walk. (This approximation is not far from the actual statistical properties in the data; see Table I.) Further assume that the innovation to the univariate state-level idiosyncratic disposable income process corresponds to the innovation to the information of the representative agent of the state. Given the random walk assumption, this is also the innovation to the state's permanent income. Thus, idiosyncratic consumption should move one-to-one with movements in idiosyncratic state-level disposable income, that is, the regression $\Delta C_{it} - \Delta C_t = \alpha_i + b\,(\Delta Y_{it} - \Delta Y_t) + \epsilon_{it}$ should yield a coefficient $b = 1$. We estimated this regression obtaining a coefficient of 0.2 which is well below unity. (We do not provide details in tables.) This result is robust to whether state specific fixed effects are included and to the manner aggregate fluctuations are controlled for. We will not explore this issue further since progress will require us to face the more daunting task of identifying the information available to the representative agent of each state,[28] but the results of the simple regression we performed appear to be consistent with excess smoothness found in previous studies.

Relation to the literature on risk sharing

The simple regressions interpreted in the previous sub-section as tests for excess smoothness (regressions of current idiosyncratic consumption on current idiosyncratic income) have a very different interpretation if we assume that consumers can *insure* their consumption ex-ante, before shocks occur. (This assumption is clearly stronger than the assumption underlying the PIH, namely that consumers only need to have full access to a credit market where they can borrow and lend ex-post, after shocks occur.) Under commonly used assumptions—symmetric information, no transaction costs, CRRA utility, identical rate of time preference for all agents—

[28]The savings behavior of agents is informative about their private information (see Campbell and Deaton 1989), but savings data are not available by state.

full (Pareto efficient) risk sharing within a group implies that $\Delta C_{it} = \Delta C_t$. That is, under full risk sharing, the growth rate of each agent's consumption will equal the growth rate of the group's aggregate consumption and not depend on any idiosyncratic characteristic of the agent (in particular, income). This holds for every period and for every realization of the state of the world. An implication is that a regression of consumption on contemporaneous income should yield a coefficient of zero if aggregate fluctuations are controlled for.

Tests of full risk sharing differ slightly in the way they control for aggregate fluctuations. The simplest method is that used by Cochrane (1991) who constructs a cross-sectional micro data set (i.e., without aggregate fluctuations over time) and regresses individual-level consumption on individual variables such as job loss, duration of job search, and duration of illness—variables that are likely to be exogenous determinants of individual-level income. Mace (1991) also uses micro data and regresses, for a panel of consumers, individual-level consumption on individual-level income controlling for aggregate fluctuations by including aggregate consumption as a regressor. If perfect risk sharing holds, then such a regression should give a coefficient of unity on aggregate consumption and zero on current income.[29]

Table V reports results of two such regressions where aggregate fluctuations are controlled for in two manners: (1) by subtracting the average consumption and disposable income from the state-by-state data;[30] and (2) by including time fixed effects.[31] In both regressions, if there is full risk sharing the coefficient δ should be zero.

In Table V, the regressor is disposable personal income which incorporates net income obtained from other states through inter-state ownership of productive assets (e.g., dividend and rental income) *and* net income from the federal government (e.g., social security benefits and net taxes). In other words, the effect of income smoothing (insurance) through diversification of income sources and income smoothing provided by the federal tax-transfer system are already reflected in the regressor. Therefore, $\delta = 0$ tests whether the full risk sharing allocation is

[29]See also Townsend (1994) who regresses individual-level consumption on individual-level income for a sample of three Indian villages, and Obstfeld (1994) who performs a similar analysis using country-level macroeconomic data. Asdrubali, Sørensen, and Yosha (1996) nest this type of regression within a decomposition of the cross-sectional variance of US state-level gross product. They introduce the notion of "levels of income and consumption smoothing" and estimate, within a system of equations, the fractions of idiosyncratic shocks to gross state product that are absorbed through each of the following mechanisms—diversification of income sources, the federal tax-transfer system, and saving behavior. The last regression in their system of equations is similar to that used by Cochrane (1991), Mace (1991), and Townsend (1994), but uses gross state product as a regressor.

[30]This specification is similar to that used by Mace (1991) although we impose a unit coefficient on aggregate consumption by subtracting it from state-level consumption.

[31]This specification is similar to that used by Cochrane (1991), since the estimated coefficient from a panel regression with time fixed effects is equivalent to a weighted average of period-by-period cross-sectional regressions; see Asdrubali, Sørensen, and Yosha (1996), footnote 5.

achieved through borrowing and lending *and* through informal insurance, e.g., among friends and within families. Like the overwhelming majority of such tests in the literature, we too reject the hypothesis of full risk sharing.[32]

Recall that with no informal risk sharing, and assuming that innovations to permanent income are the same as innovations to the univariate series for disposable income, PIH implies $\delta = 1$. Our estimate $\delta = 0.2$ is thus "closer" to 0, the predicted value under full risk sharing. A full blown interpretation is still a challenge and is left for future research, but we want to make the point—seemingly ignored in the literature—that tests of excess smoothness typically assume that there is no informal risk sharing. If there is such risk sharing, more explicit modeling is needed to formulate what "no excess smoothness" means. By contrast, excess sensitivity tests (Hall 1978, Hansen and Singleton 1982, 1983) are not affected by whether or not there is informal risk sharing.

Canadian province-level disposable income and consumption

To verify the robustness of our empirical findings, we perform a similar analysis for Canadian provinces.[33] Data are available from the CANSIM database maintained by Statistics Canada. We use the series personal disposable income, non-durable consumption (defined as the sum of non-durables, semi-durables, and services), population, and aggregate consumer prices. The sample period is 1961–1996.

The statistical properties of the province-level income and consumption series are similar to their US state-level counterparts. The IPS test for a unit root in y_{it} rejects the null hypothesis of non-stationarity for an AR-model with 1 lag, but not for models with 2 or 3 lags (P-values of 0.00, 0.07, and 0.07, respectively).[34] For c_{it}, the IPS test easily accepts non-stationarity. We estimated AR(2) models for the differenced series obtaining coefficients on twice lagged variables that are typically very small, typically insignificant, and jointly not different from zero. AR(1) models for province-level total and idiosyncratic disposable income and consumption are similar

[32]There are many potential explanations for less than full risk sharing; see, e.g., Kocherlakota (1996) who stresses limited enforceability and commitment. Others (see, e.g., Heaton and Lucas 1996 and Constantinides and Duffie 1996) study the conditions that ensure that the full risk sharing allocation is approximated (or even achieved) among heterogeneous agents in the absence of insurance opportunities when only inter-temporal smoothing is present. Whether our estimated coefficient, $\delta = 0.2$, can be interpreted as "full risk sharing being closely approximated" is not evident, although it certainly indicates that the degree of inter-state consumption insurance is substantial—in terms of the method suggested by Asdrubali, Sørensen, and Yosha (1996), $\delta = 0.2$ means that eighty percent of idiosyncratic shocks are absorbed on average through inter-state consumption insurance.

[33]We report only excess sensitivity tests in a table. The other results can be obtained in table format from the authors upon request.

[34]Province-by-province ADF tests provide no evidence against unit roots in these series.

to their counterparts for US states:[35] ΔY_t and ΔC_t are positively autocorrelated with average AR(1) coefficients of 0.14 and 0.31; for $\Delta Y_{it} - \Delta Y_t$, the average AR(1) coefficient is –0.08, and for $\Delta C_{it} - \Delta C_t$ the average AR(1) coefficient is 0.15.[36]

In Table VI, we replicate our main findings for the US displayed in Table II. The results show patterns that are qualitatively very similar to those found for US states. Without controlling for aggregate fluctuations, namely by regressing ΔC_{it} on $\Delta Y_{i,t-1}$, the coefficient of lagged disposable income is 0.11 and is highly significant. When aggregate fluctuations are controlled for by regressing $\Delta C_{it} - \Delta C_t$ on $\Delta Y_{i,t-1} - \Delta Y_t$, or by including time fixed effects, the coefficient of lagged income falls drastically to 0.03. If we include aggregate consumption as a regressor, the excess sensitivity coefficient falls to 0.04. In none of the regressions that control for aggregate constraints is the PIH model rejected![37]

We further performed regressions similar to those reported for US states in Table IV ("Richer Dynamics"). We do not report the details, but the broad findings are very similar to the findings for the US.[38]

5 Concluding Remark

We have argued that the PIH may be a good model for describing the reaction of idiosyncratic consumption to idiosyncratic disposable income shocks, and found empirically that this is true for both the US and Canada. However, in states where disposable income shocks are highly persistent, idiosyncratic consumption still exhibits considerable excess sensitivity. Moreover, the lagged consumption to income ratio predicts consumption (regardless of persistence), which constitutes yet another deviation from the PIH. Our conclusion is that aggregate constraints are important for reconciling the PIH with observed patterns of consumption, but that more remains to be done.

[35] We allow for cross-correlations of the disturbances across provinces, as in the estimation for US states; see footnote 15.

[36] Since province-level consumption is part of province-level "National Accounts," measurement error is likely to be less severe than in the US state-level retail sales data. The lower autocorrelation in idiosyncratic province-level consumption, relative to the autocorrelation in province-level consumption, is in and of itself *prima facie* evidence that the PIH model fits province-specific data better than aggregate data.

[37] It is also interesting that the excess sensitivity coefficient using total (non-idiosyncratic) data is lower for Canada—a more open country—than for the US (0.11 versus 0.23; see Table II). This is consistent with our interpretation that the empirical failure of the PIH model for aggregate data is partly due to the closedness of the aggregate economies.

[38] Curiously, the error-correction term has a positive coefficient for Canada.

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Table I

US State-Level Disposable Income and Consumption Processes

Model: $\Delta Y_{it} = \alpha_i + \phi_i \, \Delta Y_{i,t-1} + \epsilon_{it}$

	Mean	Range
ϕ_i :	0.16	[–0.30 , 0.37]
Abs. value of t-statistics:	2.46	[0.14 , 4.35]

Model: $\Delta Y_{it} - \Delta Y_t = \alpha_i + \phi_i \, (\Delta Y_{i,t-1} - \Delta Y_{t-1}) + \epsilon_{it}$

	Mean	Range
ϕ_i :	0.05	[–0.48 , 0.46]
Abs. value of t-statistics:	2.65	[0.01 , 6.96]

Model: $\Delta C_{it} = \alpha_i + \phi_i \, \Delta C_{i,t-1} + \epsilon_{it}$

	Mean	Range
ϕ_i :	–0.06	[–0.46 , 0.20]
Abs. value of t-statistics:	1.47	[0.18 , 5.83]

Model: $\Delta C_{it} - \Delta C_t = \alpha_i + \phi_i \, (\Delta C_{i,t-1} - \Delta C_{t-1}) + \epsilon_{it}$

	Mean	Range
ϕ_i :	–0.08	[–0.49 , 0.35]
Abs. value of t-statistics:	1.88	[0.03 , 6.26]

Notes. ΔY_{it} is the period t log-difference of state i's (total) per capita disposable income. $\Delta Y_{it} - \Delta Y_t$ is the period t log-difference of state i's idiosyncratic (state-specific) per capita disposable income, where ΔY_t is the period t log-difference of aggregate (US-wide) per capita disposable income. Similarly for the consumption series. State-level consumption is proxied by non-durable retail sales. In all four regressions, the null hypothesis $\phi_i = \phi$ for all i is strongly rejected with P-values of 0.00. Sample period: 1964–1995.

Table II

Sensitivity of US State-Level Consumption to Lagged Income

	Estimate	t-statistic
Model: $\Delta C_{it} = \alpha_i + b\,\Delta Y_{i,t-1} + \epsilon_{it}$		
b :	0.23	8.69
Model: $\Delta C_{it} - \Delta C_t) = \alpha_i + b\,(\Delta Y_{i,t-1} - \Delta Y_{t-1}) + \epsilon_{it}$		
b :	0.12	4.95
Model: $\Delta C_{it} = \alpha_t + \alpha_i + b\,\Delta Y_{i,t-1} + \epsilon_{it}$		
b :	0.11	4.63
Model: $\Delta C_{it} = \alpha_i + \gamma\,\Delta C_t + b\,(\Delta Y_{i,t-1} - \Delta Y_{t-1}) + \epsilon_{it}$		
γ :	1.0	86.84
b :	0.12	4.94
Model: $\Delta C_{it} = \alpha_i + \gamma_i\,\Delta C_t + b\,(\Delta Y_{i,t-1} - \Delta Y_{t-1}) + \epsilon_{it}$		
γ_i (mean):	1.00	
Range:	[0.17 , 1.76]	[0.46 , 8.27]
b :	0.12	4.90

Notes. ΔY_{it} is the period t log-difference of state i's (total) per capita disposable income. $\Delta Y_{it} - \Delta Y_t$ is the period t log-difference of state i's idiosyncratic (state-specific) per capita disposable income, where ΔY_t is the period t log-difference of aggregate (US-wide) per capita disposable income. Similarly for the consumption series. State-level consumption is proxied by non-durable retail sales. Sample period: 1964–1995.

Table III

Sensitivity of Consumption to Lagged Income:

High versus Low Persistence in Income

	Low Persistence	Medium Persistence	High Persistence
Average ϕ_i	–0.18	0.11	0.37
Model: $\Delta C_{it} = \alpha_i + b\,\Delta Y_{i,t-1} + \epsilon_{it}$			
b :	0.07	0.25	0.42
t-statistic:	1.49	3.12	6.31
Model: $\Delta C_{it} - \Delta C_t = \alpha_i + b\,(\Delta Y_{i,t-1} - \Delta Y_{t-1}) + \epsilon_{it}$			
b :	–0.03	0.14	0.37
t-statistic:	–0.81	1.77	5.65
Model: $\Delta C_{it} = \alpha_i + \alpha_t + b\,\Delta Y_{i,t-1} + \epsilon_{it}$			
b :	–0.04	0.09	0.36
t-statistic:	–1.02	1.20	5.79

Notes. ΔY_{it} is the period t log-difference of state i's (total) per capita disposable income. $\Delta Y_{it} - \Delta Y_t$ is the period t log-difference of state i's idiosyncratic (state-specific) per capita disposable income, where ΔY_t is the period t log-difference of aggregate (US-wide) per capita disposable income. Similarly for the consumption series. State-level consumption is proxied by non-durable retail sales. Sample period: 1964–1995. States are classified according to the persistence of the state-specific component of disposable income, as measured by the coefficient ϕ_i in the regression $\Delta(Y_{it} - Y_t) = \alpha_i + \phi_i \Delta(Y_{i,t-1} - Y_{t-1}) + \epsilon_{it}$, estimated for each state i separately. "Average ϕ_i" is the average of the ϕ_i coefficients over the states in the group.

Table IV

Sensitivity of US State-Level Consumption to Lagged Income: Richer Dynamics

	Estimate	t-statistic
Model: $\Delta C_{it} = \alpha_i + b_1\, \Delta Y_{i,t-1} + b_2\, \Delta Y_{i,t-2} + \epsilon_{it}$		
b_1 :	0.21	7.93
b_2 :	0.06	2.43
Model: $\Delta C_{it} = \alpha_i + b_1\, \Delta Y_{i,t-1} + b_2\, \Delta Y_{i,t-2} + b_3(C_{i,t-1} - Y_{i,t-1}) + \epsilon_{it}$		
b_1 :	0.19	6.69
b_2 :	0.07	2.48
b_3 :	–0.09	–7.21
Model: $\Delta c_{it} = \alpha_i + b_1\, \Delta y_{i,t-1} + b_2\, \Delta y_{i,t-2} + \epsilon_{it}$		
b_1 :	0.12	4.86
b_2 :	0.07	2.78
Model: $\Delta c_{it} = \alpha_i + b_1\, \Delta y_{i,t-1} + b_2\, \Delta y_{i,t-2} + b_3(c_{i,t-1} - y_{i,t-1}) + \epsilon_{it}$		
b_1 :	0.00	–0.14
b_2 :	–0.03	–1.16
b_3 :	–0.19	–12.68

Notes. ΔY_{it} is the period t log-difference of state i's (total) per capita disposable income. $\Delta y_{it} = \Delta Y_{it} - \Delta Y_t$ is the period t log-difference of state i's idiosyncratic (state-specific) per capita disposable income, where ΔY_t is the period t log-difference of aggregate (US-wide) per capita disposable income. Similarly for the consumption series. State-level consumption is proxied by non-durable retail sales. Sample period: 1965–1995.

Table V

Testing for Full Inter-State Risk Sharing

Model: $\Delta C_{it} - \Delta C_t = \alpha + \delta\,\Delta(Y_{it} - \Delta Y_t) + \epsilon_{it}$

δ :	0.23
t-statistic:	9.61

Model: $\Delta C_{it} = \nu_t + \ \delta\,\Delta Y_{it} + \epsilon_{it}$

δ :	0.22
t-statistic:	9.11

Notes. ΔY_{it} is the period t log-difference of state i's (total) per capita disposable income. $\Delta Y_{it} - \Delta Y_t$ is the period t log-difference of state i's idiosyncratic (state-specific) per capita disposable income, where ΔY_t is the period t log-difference of aggregate (US-wide) per capita disposable income. Similarly for the consumption series. State-level consumption is proxied by non-durable retail sales. Sample period: 1964–1995.

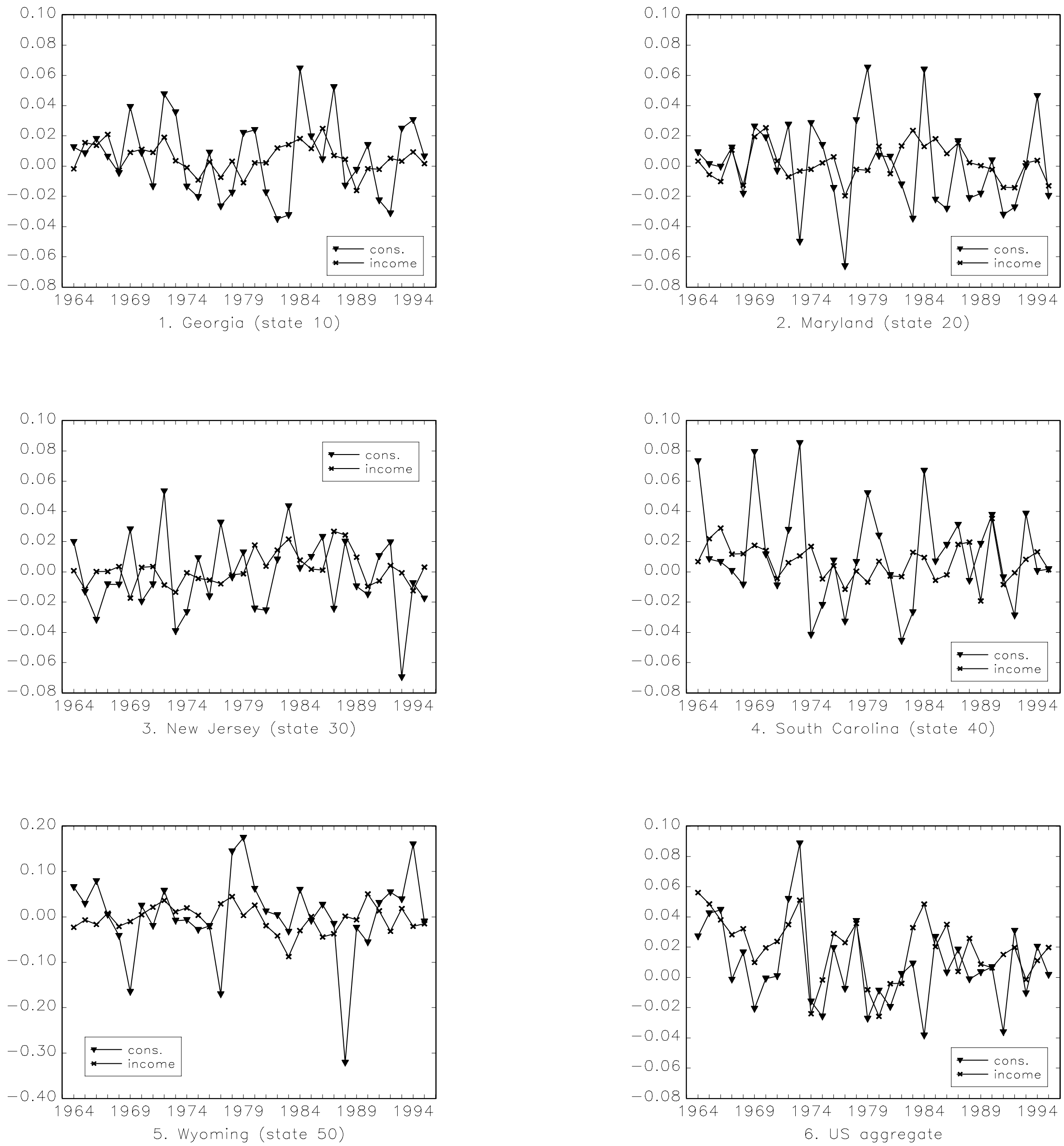
Table VI

Sensitivity of Canadian Province-Level Consumption to Lagged Income

	Estimate	t-statistic
	Model: $\Delta C_{it} = \alpha_i + b\,\Delta Y_{i,t-1} + \epsilon_{it}$	
b :	0.11	4.32
	Model: $\Delta C_{it} - \Delta C_t = \alpha_i + b\,(\Delta Y_{i,t-1} - \Delta Y_{t-1}) + \epsilon_{it}$	
b :	0.03	1.62
	Model: $\Delta C_{it} = \alpha_t + \alpha_i + b\,\Delta Y_{i,t-1} + \epsilon_{it}$	
b :	0.03	1.15
	Model: $\Delta C_{it} = \alpha_i + \gamma\,\Delta C_t + b\,(\Delta Y_{i,t-1} - \Delta Y_{t-1}) + \epsilon_{it}$	
γ :	1.00	142.78
b :	0.04	1.68
	Model: $\Delta C_{it} = \alpha_i + \gamma_i\,\Delta C_t + b\,\Delta(Y_{i,t-1} - \Delta Y_{t-1}) + \epsilon_{it}$	
γ_i (mean):	0.97	
Range:	[0.81 , 1.09]	[4.44 , 16.21]
b :	0.04	1.66

Notes. ΔY_{it} is the period t log-difference of province i's per capita disposable income. $\Delta Y_{it} - \Delta Y_t$ is the period t log-difference of province i's idiosyncratic (province-specific) per capita disposable income, where ΔY_t is the period t log-difference of aggregate (Canada-wide) per capita disposable income. Similarly for the consumption series. Sample period: 1962–1996.

Figure 1: US State Idiosyncratic Consumption and Income



Notes. Panels 1–5: Income is the log-difference of state idiosyncratic (state-specific) per capita disposable income. Similar for the consumption series. Consumption is proxied by non-durable retail sales. Panel 6: Income is the log-differences of aggregate (US-wide) per capita disposable income. Consumption is the log-difference of aggregate (US-wide) per capita non-durable retail sales. Sample period: 1964–1995.