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Consumption and Aggregate Constraints: International Evidence*

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Abstract

This paper documents that region-level consumption exhibits excess sensitivity to lagged region-level income in Italy, Japan, Spain, the UK and West Germany. However, *region-specific* consumption exhibits substantially less sensitivity to lagged region-specific income. Moreover, excess sensitivity is inversely related to standard measures of openness and credit market integration and for most countries, it has decreased over time. These findings are consistent with the results reported by Ostergaard *et al.* [*Journal of Political Economy* (2002) Vol. 110, pp. 634–645] for US states and Canadian provinces, and provide empirical support for the hypothesis that closed-economy constraints may partly be responsible for the excess sensitivity phenomenon in aggregate data.

I. Introduction

A key implication of the rational expectations version of the Permanent Income Hypothesis (PIH) is that consumption should follow a martingale process. Hall (1978), who first discussed this implication, found that variables known in the previous period were generally insignificant in explaining the

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change in current consumption. By contrast, many subsequent studies found consumption to be 'too sensitive' to lagged information on income (e.g. Flavin, 1981; Hayashi, 1982; Nelson, 1987; Hansen and Singleton, 1983), and those have spawned an extensive literature that seeks to explain this *excess sensitivity* phenomenon.

In a recent paper, Ostergaard, Sørensen and Yosha (2002) use regional data on personal disposable income and consumption for the US and Canada to test the PIH. Consistent with previous findings in the literature, they report that region-level consumption exhibits excess sensitivity to lagged region-level income. However, when aggregate (nation-wide) fluctuations are controlled for, they find that region-specific consumption exhibits substantially less sensitivity to lagged region-specific income. They conjecture that this result may be explained by closed-economy constraints, namely, frictions in international credit markets and/or the slow adjustment of net imports in response to fluctuations in aggregate consumption demand. Indeed, at the country level, it is conceivable that it may take time to borrow in international credit markets and adjust the quantity of goods imported. Regions within countries, on the contrary, are relatively more open in the sense that they can easily borrow and import goods among themselves. Thus, even if the PIH model fails at the country level, the model should perform better with region-specific (idiosyncratic) consumption and income.

In this paper, we provide international evidence on how ignoring closed-economy constraints can lead to rejection of the PIH in aggregate data. We show that the result obtained by Ostergaard *et al.* (2002) is very robust and that excess sensitivity is more severe in less open economies, providing the first empirical support for the underlying economic mechanism that Ostergaard *et al.* (2002) only conjecture. Using regional data from Italy, Japan, Spain, the UK, West Germany, Canada and the US, we find considerable excess sensitivity of consumption to lagged income, in line with results of previous studies using country-level data (e.g. Japelli and Pagano, 1989; Bacchetta and Gerlach, 1997). Next, we control for aggregate fluctuations in consumption and income, and show that region-specific consumption exhibits much less sensitivity to lagged region-specific income. We also perform the same analysis for a sample of 22 Organization for Economic Co-operation and Development (OECD) countries to determine if our findings are a result of the methodology used. In this case, we treat each country as a region and the OECD as the aggregate. The results indicate that controlling for aggregate effects leads only to a slight reduction in the excess sensitivity coefficient. Restricting the sample to European Union members, on the contrary, further reduces the excess sensitivity coefficient, suggesting that closed-economy constraints may be less binding within this group. Finally and perhaps more importantly, we document that the magnitude of the excess sensitivity coefficient is negatively correlated with standard measures of openness and financial market integration and that for most countries, excess sensitivity has decreased over time.

Overall, these results ratify the hypothesis that excess sensitivity in macroeconomic data can be partly explained by closed-economy constraints or the lack of integration in credit/goods markets across countries. Thus, our findings suggest that national borders do matter. The remainder of the paper is organized as follows. Section II presents the empirical methodology. Section III discusses the data, while section IV presents the estimation results. Section V explores the relationship between excess sensitivity, openness and financial market integration. Section VI concludes.

II. Empirical methodology

For a given country, C_{it} and Y_{it} denote region i's real per capita consumption and real per capita income in period t respectively, and C_t and Y_t denote aggregate (nation-wide) per capita consumption and income. A standard test for excess sensitivity is to regress current consumption changes on lagged income changes, that is,

$$\text{Model 1:}\quad \Delta \log C_{it} = \alpha_i + \beta\Delta \log Y_{i,t-1} + \varepsilon_{it}, \tag{1}$$

where α_i are region-fixed effects, β represents the excess sensitivity parameter, and ε_{it} is a zero-mean random disturbance term.

Under the PIH, consumption follows a martingale process. Thus, the parameter β should be equal to zero (i.e. changes in current consumption should be uncorrelated with lagged changes in income). On the contrary, if β is significantly different from zero, consumption is said to be excessively sensitive to lagged income. With regional data, excess sensitivity of consumption could appear because: (i) credit/goods markets are not well integrated across countries, and/or (ii) regions within a country are not fully integrated.[1]

One way of disentangling these two effects is to remove the aggregate component in regional data, and estimate model 1 using region-specific income and consumption data. If regions within countries are relatively well integrated, we should observe less (or no) excess sensitivity after controlling

[1]This test of the PIH relies on a constant interest rate. When a country is not well integrated in the world credit market, an increase in aggregate consumption demand would imply more competition for domestic funds creating an upward pressure in interest rates. Tests of the PIH in aggregate data which allow for time-varying interest rates (Mankiw, 1981), and time-varying stochastic interest rates (Hansen and Singleton, 1982, 1983) also fail. Ostergaard *et al.* (2002) and Sørensen and Yosha (2000) conjecture that measured interest rates may not fully capture close-economy constraints, so we follow their approach of controlling for aggregate effects more generally instead.

for aggregate effects. We use three alternative approaches to account for aggregate fluctuations. First, aggregate consumption and income are subtracted from their corresponding regional components to obtain region-specific (idiosyncratic) consumption and income, i.e. we estimate the following regression equation:

$$\text{Model 2:} \quad \Delta(\log C_{it} - \log C_t) = \alpha_i + \beta\Delta(\log Y_{i,t-1} - \log Y_{t-1}) + \varepsilon_{it}. \quad (2)$$

Secondly, we allow for time-fixed effects, υ_t, that capture common shocks to all regions in the country.

$$\text{Model 3:} \quad \Delta \log C_{it} = \alpha_i + \upsilon_t + \beta\Delta \log Y_{i,t-1} + \varepsilon_{it}. \quad (3)$$

Finally, aggregate effects are taken into account by including the change in aggregate consumption in the regression. In essence, shocks that affect the entire economy may already be reflected in aggregate consumption movements.

$$\text{Model 4:} \quad \Delta \log C_{it} = \alpha_i + \beta\Delta(\log Y_{i,t-1} - \log Y_{t-1}) + \gamma\Delta \log C_t + \varepsilon_{it}. \quad (4)$$

If frictions in international credit/goods markets are important, then the estimate of β in model 1 should be significantly different from zero. However, if regions within countries are relatively well integrated, region-specific consumption should exhibit little or no sensitivity to lagged region-specific income, such that the estimate of β in models 2–4 should be insignificantly different from zero. Note that C_{it} should measure consumption of region *i*'s residents, as opposed to consumption in region *i*. Otherwise, at least part of the decrease in β would be due to cross-border spending. Region *i* may experience an idiosyncratic negative income shock but some consumers from neighbouring regions with positive idiosyncratic shocks could increase spending in region *i*, making idiosyncratic consumption in region *i* fairly smoothed. Models 1–4 are estimated using feasible generalized least squares allowing for heteroscedasticity across regions (see Greene, 2003).[2]

III. Data

Table 1 summarizes the regional data availability and data sources for the countries considered: Italy, Japan, Spain, the UK, West Germany, the US and

[2]We do not allow for contemporaneous spatial correlation of the error terms across regions since most of our panels are relatively short. We perform a robustness check by estimating OLS coefficients and calculating panel-corrected standard errors, allowing for heteroscedasticity and contemporaneous correlation of the error terms across regions (see Beck and Katz, 1995). Results are fairly robust. The estimated coefficients are in the same order of magnitude and the inclusion of aggregate effects decreases the excess sensitivity parameters substantially. However, the standard errors are substantially higher. For the UK, Germany and Japan, the initial estimated excess sensitivity parameter is not significant.

TABLE 1

Regional data on consumption and income

	Country	No. regions	Years	Consumption	Disp. income	GRP
(1)	UK	11	1971–94	Total	Yes	No
(2)	West Germany	11	1970–97	Total	No	Yes
(3)	Italy	20	1980–95	Total, nondur.	No	Yes
(4)	Spain (regions)	18	1985–96	Total	Yes	Yes
(5)	Spain (provinces)	52	1967–95	Total	Yes	Yes
(6)	Japan	47	1975–90	Total	No	Yes
(7)	US	50	1963–95	Retail sales	Yes	No
(8)	Canada	10	1961–96	Total, nondur.	Yes	No
(9)	OECD	22	1960–97	Total	No	Yes
(10)	EU-15	15	1960–97	Total	No	Yes

Sources:
(1) UK data from *Regional Trends*, a yearly publication of the UK. Central Statistical Office.
(2) West Germany's data from '*Arbeitskreis Volkswirtschaftliche Gesamtrechnungen der Länder*', Statistisches Landesamt Baden-Wúrttemberg, Stuttgard.
(3) Italian Data from '*Conti Economici Territoriali: Conti Regionali*' published by ISTAT, the Italian Central Statistical Office.
(4) Spain's regions data from '*Contabilidad Regional de España*' published by the INE, the Spanish National Statistical Office.
(5) Spain's provinces data from '*Renta Nacional de España y su Distribución Provincial*' published by the BBV, Banco Bilbao Vizcaya. Data are biannual.
(6) Japan's data from *Annual Report of Prefectural Accounts*, published by the Economic Planning Agency of the Government of Japan.
(7) US data from the *Bureau of Economic Analysis* for disposable income and GDP and from the *Survey of Buying Power* for Retail Sales.
(8) Canadian data from CANSIM.
(9)–(10) OECD and EU-15 data from the *Penn World Tables*. Data for Germany starts in 1970.

Canada. Annual data are employed except for the Spanish provinces, which are biannual.[3] Regional population and national consumer price index (CPI) are utilized to express the consumption and income series in real per-capita terms.

It is well known that the PIH applies best to the relationship between non-durable consumption and disposable income. However, we consider alternative series whenever these variables are not available. Regional disposable income is not obtainable for all the countries in our sample and so, as in many studies, gross regional product (GRP) is used as a proxy variable. For those countries where disposable income is available, we report results using both GRP and regional disposable income. Moreover, only one country in our sample, Italy, has regional non-durable consumption so we use regional total consumption instead. We present data for the US and Canada to facilitate the

[3]Spain has a three-tier level of government: central, regional and local. At the regional level there are 17 'self-governing' (autonomous) communities plus the cities of Ceuta and Melilla. We refer to these communities as regions. At the local level, there are 50 provinces plus the cities of Ceuta and Melilla. We use data for both regions and provinces because they come from different sources and are available for different time periods. See Table 1 for details.

comparison with the results of Ostergaard *et al.* (2002). For the US, data on non-durable consumption are not available at the state level so retail sales are used as a proxy.[4]

In addition, we consider a sample of OECD countries. We use data on annual Purchasing Power Parity (PPP)-adjusted real GDP, real total consumption, and population over the period 1960–97 obtained from the Penn World Tables. Our OECD sample consists of the following 22 countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, Luxembourg, the Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, the UK, and the US.[5] Finally, in some regressions, we restrict the sample to the 15 European Union members (EU-15) before May 2004: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Portugal, Spain, Sweden, and the UK.

Before proceeding with the empirical analysis, we present some summary statistics of our regional data in Table 2. Column (1) reports the panel average of regional consumption and income growth. It is apparent that average consumption growth is very similar to income (or GRP) growth in most countries, with the exception of the US and Japan, where it is lower. For the US, this is not surprising as consumption is proxied by retail sales which do not include expenditures on services such as health and education that tend to increase as income increases. For Japan, on the contrary, the result may be explained by demographic factors (i.e. older population) or simply higher saving.

With respect to the dispersion of consumption and income growth, column (2) shows that income growth is more volatile than consumption growth except in the US, i.e. the well-known stylized fact that aggregate consumption is not as volatile as aggregate income, whether income is taken to be disposable income or GRP. For the US, the higher dispersion in consumption may reflect the larger measurement error in retail sales data.

Columns (3)–(5) report the minimum, maximum and standard deviation of the region-specific consumption and income growth rates to give an idea of regional differences within a country. Finally, column (6) shows the standard deviation of the time-specific averages to illustrate the importance of aggregate effects.

IV. Discussion of the results

As a first step in our empirical investigation, we perform augmented Dickey–Fuller tests for the presence of a unit root in the consumption and income

[4]Our regional consumption and disposable income data are residence based. Only the retail sales data for the US states are location based.

[5]Data for Germany start in 1970.

TABLE 2

Summary statistics

	mean, $\bar{x}_{..}$ (1)	$sd(x_{it})$ (2)	$max(\bar{x}_{i.})$ (3)	$min(\bar{x}_{i.})$ (4)	$sd(\bar{x}_{i.})$ (5)	$sd(\bar{x}_{.t})$ (6)
United Kingdom						
Δ log C_{it} (Total)	2.14	3.40	2.33	1.90	0.12	2.86
Δ log Y_{it} (GRP)	1.88	3.46	2.29	1.40	0.26	2.96
Δ log Y_{it} (Disp. inc.)	2.25	3.59	3.02	1.91	0.30	2.96
West Germany						
Δ log C_{it} (Total)	2.04	1.91	2.43	0.85	0.42	1.61
Δ log Y_{it} (GRP)	1.93	2.14	2.42	1.47	0.27	1.75
Italy						
Δ log C_{it} (Total)	1.87	1.87	2.32	1.37	0.28	1.64
Δ log C_{it} (Nondur.)	1.72	1.95	2.19	1.27	0.23	1.59
Δ log Y_{it} (GRP)	1.57	2.04	2.43	0.92	0.36	1.28
Spain (regions)						
Δ log C_{it} (Total)	2.98	2.70	3.73	2.55	0.32	1.95
Δ log Y_{it} (GRP)	3.02	2.94	3.94	1.54	0.54	2.13
Δ log Y_{it} (Disp. inc.)	2.99	3.00	3.81	2.39	0.38	2.25
Spain (provinces)						
Δ log C_{it} (Total)	2.76	3.56	4.38	1.46	0.52	3.02
Δ log Y_{it} (GRP)	2.84	2.84	3.95	1.58	0.56	2.33
Δ log Y_{it} (Disp. inc.)	2.73	3.62	4.31	1.47	0.51	3.13
Japan						
Δ log C_{it} (Total)	2.47	2.02	3.65	1.80	0.46	1.00
Δ log Y_{it} (GRP)	3.75	2.47	7.46	2.85	0.74	1.29
US						
Δ log C_{it} (Total retail sales)	1.21	5.15	2.34	0.22	0.46	3.32
Δ log Y_{it} (Disp. inc.)	1.92	2.91	2.57	1.35	0.32	2.21
Canada						
Δ log C_{it} (Nondur.)	2.24	1.96	2.80	1.81	0.27	1.52
Δ log Y_{it} (Disp. inc.)	2.33	3.78	2.94	1.69	0.43	2.79
OECD						
Δ log C_{it} (Total)	2.50	2.50	4.23	0.97	0.73	1.32
Δ log Y_{it} (GDP)	2.64	2.67	4.41	1.10	0.74	1.60
EU-15						
Δ log C_{it} (Total)	2.62	2.67	3.53	1.55	0.53	1.46
Δ log Y_{it} (GDP)	2.77	2.69	3.71	2.04	0.50	1.73

Notes:

All variables in percentages.

$\bar{x}_{..}$ is the panel mean. $\overline{\text{sd}(x_{it})}$ is the average regional standard deviation of x; $\bar{x}_{i.}$ is region i's specific mean. $\text{sd}(\bar{x}_{i.})$ is the standard deviation of the region specific means, while $\text{sd}(\bar{x}_{.t})$ is the standard deviation of the time specific averages.

Explicitly: $\overline{\text{sd}(x_{it})} = \frac{1}{N}\Sigma_i\sqrt{\frac{\Sigma_t(x_{it}-\bar{x}_{i.})^2}{T-1}}$, $\text{sd}(\bar{x}_{i.}) = \sqrt{\frac{\Sigma_i(\bar{x}_{i.}-\bar{x}_{..})^2}{N-1}}$, and $\text{sd}(\bar{x}_{.t}) = \sqrt{\frac{\Sigma_t(\bar{x}_{.t}-\bar{x}_{..})^2}{T-1}}$.

series for each region. The values of the test statistics indicate that the null hypothesis of a unit root cannot be rejected in most regions. Thus, these results are in concordance with the widely held view that the consumption and income series display unit root or near unit-root behaviour. Consequently, we perform our empirical analysis using first differenced data.[6]

Table 3 presents the results from estimating models 1 to 4, using region-level data for each country. As shown in the first column, the estimates of the excess sensitivity coefficient β from model 1 are positive and statistically significant at the 5% level. This implies that lagged changes in income help forecast changes in current consumption, contrary to the prediction of the PIH. The magnitude of β varies across countries, ranging from 0.07 for the Japanese prefectures to 0.51 for the Spanish provinces. These estimates must, however, be compared with caution as the time period covered differs for each country. For example, the estimate of β appears to be lower for the Spanish regions than the Spanish provinces. This is probably caused by the fact that the sample period for Spanish provinces is 1967–97 vs. 1985–96 for the Spanish regions. Obviously, the integration of the Spanish economy in the world credit market was fostered after the country joined the European Union – then the European Community – in 1986. Aside from these differences, it is worth noting that Spain and Italy have higher estimates of β than Japan, the UK and West Germany. Moreover, the estimates tend to be higher when using GRP (vs. disposable income) and total consumption (vs. non-durable consumption) data.

The second column of Table 3 presents the results from model 2, using region-specific income and consumption observations in the regression. It is apparent that the estimates of β are much smaller than those in the first column. In fact, the estimates are small in magnitude and insignificantly different from zero in many countries. The third and fourth columns present the results based on alternative ways of controlling for aggregate effects. Similar to those in the second column, the estimates of β are often insignificantly different from zero. Hence, these findings suggest that regions within a country suffer less from closed-economy constraints than the country as a whole, as hypothesized.

We next perform the same analysis for a sample of 22 OECD countries, treating each country as a region and the OECD as an aggregate. Aggregate per capita OECD consumption (income) is calculated as the ratio of total consumption (income) to total population in the 22 countries. The results indicate that controlling for aggregate effects decreases but does not eliminate excess sensitivity, implying that closed-economy constraints are almost as important for each of the OECD countries as for the OECD as a whole. We also repeat the exercise for the group of EU-15 countries. In this case,

[6]Detailed tables reporting unit-root tests are available from the authors upon request.

TABLE 3

Sensitivity of regional–provincial level consumption to lagged income

Country	*Model 1* *(1)*	*Model 2* *(2)*	*Model 3* *(3)*	*Model 4* *(4)*
United Kingdom (11 regions):	0.13*	0.06	0.05	0.06
Total cons. and disp. income	(0.06)	(0.05)	(0.05)	(0.05)
United Kingdom (11 regions):	0.23*	0.01	0.01	0.01
Total cons. and GRP	(0.06)	(0.06)	(0.06)	(0.06)
West Germany (11 landers):	0.20*	−0.02	−0.00	−0.02
Total cons. and GRP	(0.05)	(0.05)	(0.04)	(0.04)
Italy (20 regions):	0.36*	0.05*	0.05*	0.05*
Total cons. and GRP	(0.05)	(0.02)	(0.02)	(0.02)
Italy (20 regions):	0.32*	0.06	0.08*	0.05
Nondur. cons. and GRP	(0.05)	(0.04)	(0.03)	(0.04)
Spain (18 regions):	0.24*	−0.07	−0.06	−0.07†
Total cons. and disp. income	(0.06)	(0.04)	(0.04)	(0.04)
Spain (18 regions):	0.42*	0.06	0.07†	0.06
Total cons. and GRP	(0.05)	(0.05)	(0.04)	(0.04)
Spain (52 provinces):	0.37*	0.07†	0.06	0.07†
Total cons. and disp. income	(0.03)	(0.04)	(0.04)	(0.04)
Spain (52 provinces):	0.51*	0.01	0.02	0.01
Total cons. and GRP	(0.03)	(0.04)	(0.05)	(0.04)
Japan (47 prefectures):	0.07*	0.02	0.01	0.02
Total cons. and GRP	(0.02)	(0.02)	(0.02)	(0.02)
United States (50 states)	0.31*	0.15*	0.14*	0.15*
Total retail sales and disp. income	(0.04)	(0.05)	(0.05)	(0.05)
Canada (10 provinces)	0.21*	0.08*	0.07*	0.08*
Total cons. and disp. income	(0.03)	(0.03)	(0.02)	(0.03)
Canada (10 provinces)	0.17*	0.02	0.01	0.01
Nondur. cons. and disp. income	(0.03)	(0.02)	(0.02)	(0.03)
OECD	0.26*	0.20*	0.18*	0.18*
Total cons. and GDP	(0.03)	(0.03)	(0.03)	(0.03)
EU-15	0.27*	0.19*	0.17*	0.17*
Total cons. and GDP	(0.04)	(0.05)	(0.05)	(0.05)

Notes:

Standard errors in parentheses. *Significant at the 5% level. †Significant at the 10% level.

C_{it} and Y_{it} are consumption and GRP/income in region i in period t. C_t and Y_t are national consumption and GDP/income respectively.

Model 1: $\Delta \log C_{it} = \alpha_i + \beta\Delta \log Y_{i,t-1} + \varepsilon_{it}$.

Model 2: $\Delta(\log C_{it} - \log C_t) = \alpha_i + \beta\Delta(\log Y_{i,t-1} - \log Y_{t-1}) + \varepsilon_{it}$.

Model 3: $\Delta \log C_{it} = \alpha_i + \upsilon_t + \beta\Delta \log Y_{i,t-1} + \varepsilon_{it}$.

Model 4: $\Delta \log C_{it} = \alpha_i + \beta\Delta(\log Y_{i,t-1} - \log Y_{t-1}) + \gamma\Delta \log C_t + \varepsilon_{it}$.

The table presents estimates for β, the excess sensitivity parameter. All models estimated using feasible GLS assuming disturbances are heteroscedastic across regions. Estimation allows for regional fixed effects, α_i. In model 3, υ_t denotes time fixed effects.

controlling for aggregate effects leads to a further reduction in the estimate of β, although only marginally.[7]

In summary, we find a substantial reduction in the excess sensitivity coefficient for the individual countries once fluctuations in aggregate consumption and income are accounted for in the regression. This result suggests that regions within a country are relatively open in the sense that they can more easily borrow and import goods among themselves in response to region-specific income shocks. As such, the excess sensitivity observed when using aggregate national-level data may be explained by closed-economy constraints or the lack of integration of a country as a whole in the world credit/goods markets rather than lack of integration of regions within a country.

V. The importance of closed-economy constraints

The foregoing discussion suggests that closed-economy constraints may be responsible for the excess sensitivity phenomenon at the country level. In this section, we investigate whether variables that proxy for closed-economy constraints have important effects on the magnitude of the excess sensitivity coefficient, β.

We start with a measure relating to the commodities market. Figure 1 shows the cross-section relation between the estimates of β from model 1 – the model which does not control for aggregate effects – and a standard measure of openness: trade (exports plus imports) as a percentage of GDP.[8] In particular, we calculate the average trade to GDP ratio of each country for the time period considered in the estimation of the excess sensitivity coefficient. Note that as different periods are used in the individual country regressions, different periods are used as well to compute this trade average (see Table 1 for details). Likewise, as we find that the estimates of β are generally higher when using GRP than when using disposable income, we separate those countries with data on regional disposable income from those with GRP.

Figure 1 is instructive. Excess sensitivity is systematically larger the higher the measure of openness, with Japan being a clear outlier. The case of Spain is particularly revealing. Spain(1) covers regional data from 1985 to 1996 while Spain(2) consists of provincial data from 1967 to 1995. Average openness is higher for Spain(1) and the excess sensitivity is lower. For Japan, the low excess sensitivity coefficient can be explained by the fact that wealth accumulation and saving rates among Japanese households are high

[7]To further examine the robustness of the results to minor changes in specification, we also estimate these four models using the variables in levels instead of logs. The results are qualitatively similar and not reported here. Tables available from the authors upon request.

[8]Data from the World Development Indicators published by the World Bank.

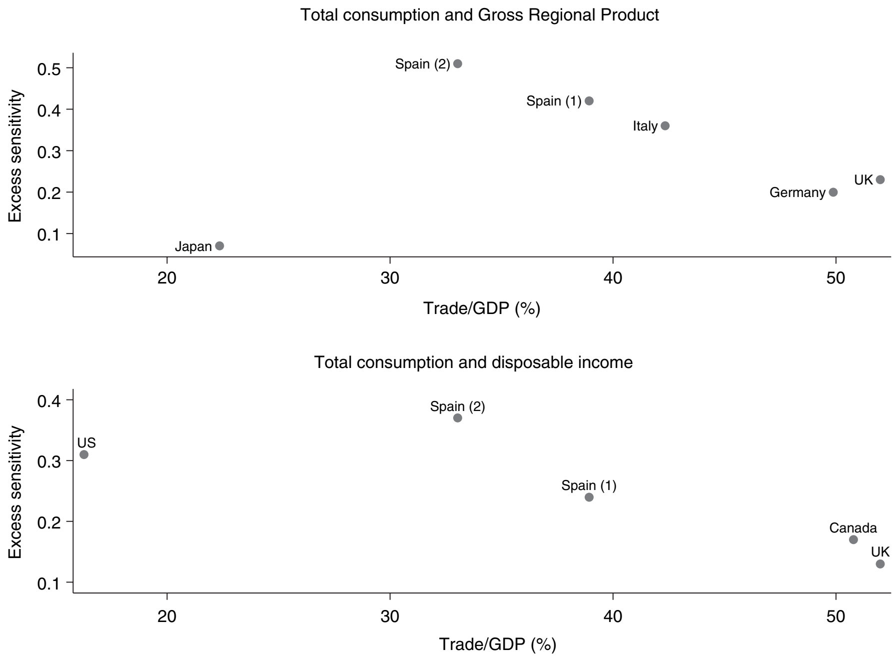

Figure 1. Excess sensitivity and trade

relative to other industrial countries.[9] Consequently, Japanese households have more means for consumption smoothing. To formally test the significance of the relationship between the variables, we calculate Pearson correlation coefficients. These are −0.98 (P-value = 0.002, excluding Japan) and −0.96 (P-value = 0.04, excluding the US) for Figure 1, top and bottom respectively, indicating an inverse relationship between openness and excess sensitivity.[10]

An alternative approach of examining the effects of openness on excess sensitivity, which takes advantage of the time-series aspect of our regional data, is to model β as a function of trade. Specifically, suppose that

$$\beta = \beta_0 + \beta_1 \text{trade}_{it}. \tag{5}$$

Substituting equation (5) into model 1 yields:

$$\Delta \log C_{it} = \alpha_i + (\beta_0 + \beta_1 \text{trade}_{it}) \Delta \log Y_{i,t-1} + \varepsilon_{it}, \tag{6}$$

trade_{it} denotes exports plus imports as a percentage of GDP, which is

[9] Hayashi (1997) shows that Japan's saving rate remains higher than the US saving rate, even after adjusting for differences in accounting methodologies. Moreover, he observes that wealth accumulation by Japanese households starts very early and continues on until very late in life.

[10] The numbers for the US are not entirely comparable as consumption is proxied by retail sales. Including the US, the correlation in the bottom panel of Figure 1 is −0.8 (P-value = 0.1).

normalized to have zero mean in order to make the estimated coefficient in the model more easily interpretable. β_0 represents the excess sensitivity coefficient when trade is equal to its average over the sample period, while β_1 represents the change in the coefficient due to trade. To the extent that frictions in the goods market are an important closed-economy constraint, we expect that more open economies should exhibit lower excess sensitivity, i.e. β_1 should be negative. The columns labelled '*Trade*' in Table 4 summarize the estimation results of equation (6).[11]

As expected, β_1 is significantly negative for all individual countries as well as for the OECD and the EU-15 groups, providing evidence that increases in trade lower the excess sensitivity in consumption.[12] Moreover, the size of β_1 varies by country (ranging from −0.011 for Canada to −0.066 for the UK), and is far from negligible in most cases. In the UK, for example, a 1% increase in trade over the sample average would lower the excess sensitivity coefficient by 0.066. However, this estimate should be interpreted with caution as we are not explicitly controlling for other factors that may reduce excess sensitivity such as a decrease in macroeconomic volatility.

Next, we proceed to examine whether closed-economy constraints relating to international credit markets are important. We follow previous studies in assuming that financial integration may increase financial efficiency, which should stimulate the demand for funds and increase the size of the domestic financial market.[13] In this regard, we measure the size of the domestic financial market, as is standard in the literature, by the value of the private credit provided by deposit money banks and other financial institutions as a percentage of GDP (credit).[14]

Figure 2 plots the estimates of β from model 1 against the credit variable. The pattern is not as clear as with the trade measure; Pearson correlations are −0.67 (P-value = 0.15) and 0.45 (P-value = 0.45) for Figure 2, top and bottom respectively. We now turn to the individual country regressions, estimating equation (6) with credit_{it} in place of trade_{it}:

$$\Delta \log C_{it} = \alpha_i + (\beta_0 + \beta_2 \text{credit}_{it}) \Delta \log Y_{i,t-1} + \varepsilon_{it}. \quad (7)$$

The columns labelled '*Credit*' in Table 4 report the regression results of equation (7). β_2 is significantly negative in most countries, implying that a

[11]Note that trade_{it} does not vary across regions in the individual country regressions but it does in the regressions with OECD and EU-15 data.

[12]We have tried including the level of the openness in the regression but, in doing so, the size of the coefficient for the interaction term decreases significantly and in many cases becomes statistically insignificant. These results are not surprising, however, because openness appears to be heavily trended in a way that the dependent variable ΔC_{it} is not. Therefore, we estimated equation (6) (excluding the level of openness) to avoid possible spurious regression results.

[13]See for example Levine and Zervos (1998) and Guiso *et al.* (2004).

[14]The 'credit' variable is obtained from the World Bank's Financial Structure Database.

TABLE 4

Excess sensitivity, trade and credit

	Trade (1)		Credit (2)		Both (3)		
	β_0	β_1 (×100)	β_0	β_2 (×100)	β_0	β_1 (×100)	β_2 (×100)
United Kingdom	0.16*	−2.37†	0.15*	0.17	0.17*	−2.05	0.10
Total cons. and disp. inc.	(0.06)	(1.44)	(0.06)	(0.15)	(0.06)	(1.53)	(0.16)
United Kingdom	0.29*	−6.64*	0.23*	0.13	0.29*	−7.56*	−0.25
Total cons. and GRP	(0.06)	(1.41)	(0.06)	(0.17)	(0.06)	(1.53)	(0.18)
West Germany	0.18*	−2.23*	0.18*	−1.05*	0.18*	−1.82*	−0.34
Total cons. and GRP	(0.05)	(0.67)	(0.05)	(0.38)	(0.05)	(0.91)	(0.52)
Italy	0.33*	−3.41*	0.28*	−11.51*	0.26*	−2.61*	−10.60*
Total cons. and GRP	(0.05)	(0.99)	(0.05)	(1.93)	(0.05)	(0.97)	(1.93)
Italy	0.28*	−4.18*	0.25*	−9.19*	0.23*	−3.61*	−8.00*
Nondur. cons. and GRP	(0.05)	(0.94)	(0.05)	(1.99)	(0.05)	(0.95)	(1.96)
Spain (regions)	0.23*	−0.64	0.32*	−5.09*	0.31*	−2.10	−5.23*
Total cons. and disp. inc.	(0.07)	(2.57)	(0.06)	(1.13)	(0.06)	(2.44)	(1.14)
Spain (regions)	0.38*	−2.75*	0.42*	−4.58*	0.37*	−2.76*	−4.64*
Total cons. and GRP	(0.06)	(1.29)	(0.05)	(0.88)	(0.05)	(1.19)	(0.87)
Spain (provinces)	0.32*	−3.29*	0.47*	−4.10*	0.43*	−1.36*	−3.38*
Total cons. and disp. inc.	(0.03)	(0.54)	(0.03)	(0.50)	(0.04)	(0.63)	(0.59)
Spain (provinces)	0.49*	−2.27*	0.59*	−3.60*	0.57*	−0.84†	−3.20*
Total cons. and GRP	(0.03)	(0.46)	(0.03)	(0.46)	(0.03)	(0.50)	(0.52)
Japan	0.04	−3.01*	0.06*	0.13*	0.03	−5.30*	−0.42*
Total cons. and GRP	(0.02)	(0.37)	(0.02)	(0.05)	(0.02)	(0.52)	(0.07)
United States	0.29*	−4.11*	0.31*	−0.12	0.38*	−9.39*	1.82*
Total cons. and disp. inc.	(0.04)	(0.88)	(0.04)	(0.24)	(0.04)	(1.36)	(0.36)
Canada	0.18*	−1.44*	0.16*	−0.77*	0.30*	3.57*	−2.44*
Total cons. and disp. inc.	(0.04)	(0.57)	(0.03)	(0.18)	(0.04)	(0.94)	(0.38)
Canada	0.17*	−1.07*	0.16*	−0.36*	0.24*	1.84*	−1.40*
Nondur. cons. and disp. inc.	(0.03)	(0.45)	(0.03)	(0.14)	(0.04)	(0.77)	(0.31)
OECD	0.25*	−1.01*	0.23*	−0.35*	0.24*	−0.80*	−0.26*
Total cons. and GDP	(0.03)	(0.21)	(0.03)	(0.08)	(0.03)	(0.22)	(0.08)
EU-15	0.25*	−0.95*	0.27*	−0.41*	0.26*	−0.78*	−0.24†
Total cons. and GDP	(0.04)	(0.22)	(0.04)	(0.13)	(0.04)	(0.23)	(0.14)

Notes:

Standard errors in parentheses. *Significant at the 5% level. †Significant at the 10% level.

Regression (1): $\Delta \log C_{it} = \alpha_i + (\beta_0 + \beta_1 \text{trade}_{it})\Delta \log Y_{i,t-1} + \varepsilon_{it}$.

Regression (2): $\Delta \log C_{it} = \alpha_i + (\beta_0 + \beta_2 \text{credit}_{it})\Delta \log Y_{i,t-1} + \varepsilon_{it}$.

Regression (3): $\Delta \log C_{it} = \alpha_i + (\beta_0 + \beta_1 \text{trade}_{it} + \beta_2 \text{credit}_{it})\Delta \log Y_{i,t-1} + \varepsilon_{it}$.

C_{it} and Y_{it} are consumption and GDP/income in region i in period t. C_t and Y_t are national consumption and GDP/income respectively. 'trade$_{it}$' is defined as exports plus imports as a percentage of GDP. 'credit$_{it}$' is the value of private credit by deposit money banks and other financial institutions as a percentage of GDP. The trade and credit variables are normalized to have zero mean for easier interpretation of the coefficients.

All models estimated using feasible GLS assuming disturbances are heteroscedastic across regions. Estimation allows for regional fixed effects, α_i.

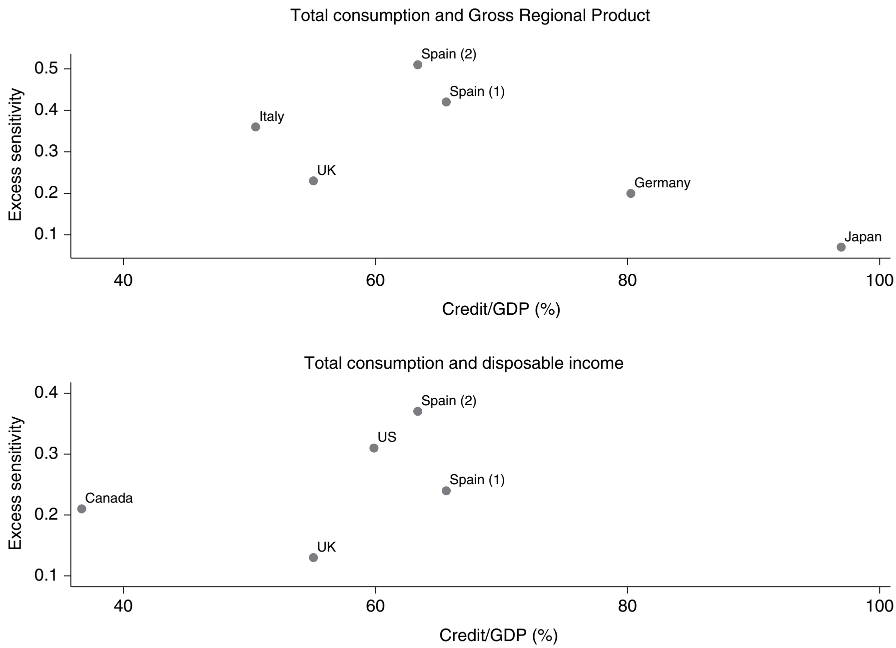

Figure 2. Excess sensitivity and credit

relaxation of international credit market imperfections would decrease excess sensitivity coefficients.[15] For the OECD and EU-15 as a whole, credit_{it} has the predicted negative effect on the excess sensitivity coefficient as well.

For completeness, we estimate the following regression which includes both the trade and credit variables in the same regression:

$$\Delta \log C_{it} = \alpha_i + (\beta_0 + \beta_1 \text{trade}_{it} + \beta_2 \text{credit}_{it}) \Delta \log Y_{i,t-1} + \varepsilon_{it}. \quad (8)$$

Results are reported in Table 4 under the columns labelled '*Both*'. It is evident that β_1 and β_2 remain negative and significant in many countries, suggesting that the two economic mechanisms may have an independent role in explaining excess sensitivity.[16]

One may argue that our proxy for financial integration may reflect mainly credit constraints (as opposed to closed-economy constraints), which are

[15] These results are in accord with those reported by Japelli and Pagano (1989) and Campbell and Mankiw (1991) among others, who found that the proportion of liquidity constrained consumers is inversely related to measures of credit availability.

[16] The two mechanisms could be complementary as well. Imports may adjust slowly because international credit markets are imperfect, which may explain why for Canada the sign of β_1 changes. For this country, the correlation between trade_{it} and credit_{it} is the highest, 0.8.

known to influence consumption behaviour and lead to rejection of the PIH. We perform another regression by taking advantage of our European sample of countries. In 1992, the European Union lifted capital controls and started the process of full financial integration. We create a dummy variable for financial integration, fi_t, which takes on the value of 0 before 1992 and 1 otherwise, and use it instead of our previous credit variable. Results are reported in Table 5. β_2 is negative and significant for all countries in the

TABLE 5

Excess sensitivity, financial integration and trade

	Financial integration (1)		*Financial integration and trade (2)*		
	β_0	β_2	β_0	β_2	β_1 (×100)
United Kingdom	0.13*	0.10	0.16*	0.08	−2.35
Total cons. and disp. inc.	(0.06)	(0.19)	(0.06)	(0.19)	(1.45)
United Kingdom	0.23*	−0.03	0.29*	−0.11	−6.69*
Total cons. and GRP	(0.06)	(0.21)	(0.06)	(0.20)	(1.41)
West Germany	0.27*	−0.35*	0.25*	−0.37*	−2.36*
Total cons. and GRP	(0.05)	(0.10)	(0.05)	(0.10)	(0.65)
Italy	0.48*	−0.59*	0.46*	−0.52*	−1.34
Total cons. and GRP	(0.05)	(0.10)	(0.05)	(0.11)	(1.05)
Italy	0.42*	−0.53*	0.38*	−0.41*	−2.57*
Nondur. cons. and GRP	(0.05)	(0.10)	(0.05)	(0.11)	(1.04)
Spain (regions)	0.31*	−0.33*	0.34*	−0.40*	3.36
Total cons. and disp. inc.	(0.06)	(0.11)	(0.07)	(0.12)	(2.72)
Spain (regions)	0.47*	−0.45*	0.51*	−0.56*	2.01
Total cons. and GRP	(0.05)	(0.09)	(0.06)	(0.12)	(1.60)
Spain (provinces)	0.47*	−0.86*	0.47*	−0.87*	0.06
Total cons. and disp. inc.	(0.03)	(0.07)	(0.03)	(0.08)	(0.60)
Spain (provinces)	0.59*	−1.14*	0.59*	−1.12*	−0.28
Total cons. and GRP	(0.03)	(0.09)	(0.03)	(0.09)	(0.45)
OECD	0.27*	−0.18*	0.26*	−0.07	−0.94*
Total cons. and GDP	(0.03)	(0.08)	(0.03)	(0.08)	(0.22)
EU-15	0.30*	−0.31*	0.27*	−0.19†	−0.79*
Total cons. and GDP	(0.04)	(0.10)	(0.04)	(0.10)	(0.23)

Notes:

Standard errors in parentheses. *Significant at the 5% level. †Significant at the 10% level.

Regression (1): $\Delta \log C_{it} = \alpha_i + (\beta_0 + \beta_2 fi_t)\Delta \log Y_{i,t-1} + \varepsilon_{it}$.

Regression (2): $\Delta \log C_{it} = \alpha_i + (\beta_0 + \beta_2 fi_t + \beta_1 \text{trade}_{it})\Delta \log Y_{i,t-1} + \varepsilon_{it}$.

C_{it} and Y_{it} are consumption and GDP/income in region i in period t. C_t and Y_t are national consumption and GDP/income respectively. 'trade$_{it}$' is defined as exports plus imports as a percentage of GDP. 'fi_t' is a dummy variable that takes the value 1 after 1992 when the EU lifted capital controls and started the process of full financial integration. The trade variable is normalized to have zero mean for easier interpretation of the coefficient. All models estimated using feasible GLS assuming disturbances are heteroscedastic across regions. Estimation allows for regional fixed effects, α_i.

TABLE 6

Excess sensitivity over time

	ζ_0	ζ_1 (×100)
United Kingdom	−0.02	1.86*
Total cons. and disp. inc.	(0.09)	(0.75)
United Kingdom	0.01	2.15*
Total cons. and GRP	(0.10)	(0.79)
West Germany	0.33*	−1.09*
Total cons. and GRP	(0.08)	(0.51)
Italy	0.58*	−3.07*
Total cons. and GRP	(0.09)	(1.14)
Italy	0.50*	−2.44*
Nondur. cons. and GRP	(0.09)	(1.12)
Spain (regions)	0.72*	−11.29*
Total cons. and disp. inc.	(0.10)	(2.12)
Spain (regions)	0.72*	−9.03*
Total cons. and GRP	(0.07)	(1.43)
Spain (provinces)	0.59*	−4.47*
Total cons. and disp. inc.	(0.04)	(0.60)
Spain (provinces)	0.70*	−3.41*
Total cons. and GRP	(0.04)	(0.53)
Japan	0.03	0.57*
Total cons. and GRP	(0.03)	(0.29)
United States	0.40*	−0.70
Total cons. and disp. inc.	(0.07)	(0.42)
Canada	0.33*	−1.03*
Total cons. and disp. inc.	(0.05)	(0.36)
Canada	0.22*	−0.39
Nondur. cons. and disp. inc.	(0.04)	(0.28)
OECD	0.41*	−1.11*
Total cons. and GDP	(0.04)	(0.21)
EU-15	0.44*	−1.15*
Total cons. and GDP	(0.05)	(0.27)

Notes:
Standard errors in parentheses. *Significant at the 5% level.
Regression: $\Delta \log C_{it} = \alpha_i + (\zeta_0 + \zeta_1 t)\Delta \log Y_{i,t-1} + \varepsilon_{it}$.
All models estimated using feasible GLS assuming disturbances are heteroscedastic across regions. Estimation allows for regional fixed effects, α_i.

European sample except the UK. For the OECD and the EU-15, the coefficient is negative and significant as well. When adding the trade interaction term to the regression, the effect of trade remains negative and significant for the individual European countries and the EU-15 group but it becomes insignificant for the OECD.

Our previous results suggest that as countries become more open and integrated in the world credit market, we should expect excess sensitivity to decrease. In order to check if this pattern is observed in our sample of countries, we run our final regression:

$$\Delta \log C_{it} = \alpha_i + (\beta_0 + \zeta t)\Delta \log Y_{i,t-1} + \varepsilon_{it}, \tag{9}$$

where ζ captures changes in the excess sensitivity coefficient over time. Table 6 summarizes the results.

For all countries except the UK and Japan, ζ is estimated to be negative and is significant in most cases. The results for the UK and Japan probably reflect the fact that the history of financial deregulation/liberalization in these countries is too complex to be modelled by a simple linear trend function. Hall (1991) documents that although the process of financial deregulation in the UK started in the early 1970s, the various policies implemented at the early stage – at times conflicting – seemed not to have reached their intended objectives. In the case of Japan, the financial system was slowly deregulated in the 1980s but the comprehensive financial system reform came in 1996, under the so-called Japanese Financial Big Bang (see Craig, 1998). Furthermore, for these two countries, trade does not seem to have increased significantly during our estimation period.[17]

In summary, estimates of the excess sensitivity coefficient are inversely related to standard measures of openness and financial market integration, and for most countries excess sensitivity has decreased over time. These facts may be explained by a relaxation of closed-economy constraints which may arise from either an increase in trade, a better integration of financial markets, or both.

VI. Concluding remarks

In this paper, we examine the relationship between changes in consumption and lagged changes in income using regional data from Italy, Japan, Spain, the UK, West Germany, Canada and the US. Hall's (1978) version of the PIH predicts that consumption should follow a martingale process, i.e. changes in consumption should be independent of lagged changes in income.

Our empirical findings reveal that region-level consumption exhibits excess sensitivity to lagged income. However, we also find that region-specific consumption exhibits substantially less sensitivity to lagged region-specific

[17]Fitting a linear trend to the trade measure for the different countries delivered the following coefficients: UK, 0.01; Germany, 0.48*; Italy, −0.25; Spain (regions), 0.73*; Spain (provinces), 0.64*; Japan −0.55*; Canada, 0.82*; and the US, 0.42*. An asterisk (*) denotes significant at the 5% level. In Italy, the series had a clear U-shape.

income. Thus, once aggregate income and consumption fluctuations are controlled for, the deviation from PIH consumption behaviour in macroeconomic data becomes smaller. We also document that estimated excess sensitivity coefficients are inversely related to standard measures of openness and credit market integration and that for most countries, excess sensitivity has decreased over time. Taken together, these findings are consistent with those reported in Ostergaard *et al.* (2002) for US states and Canadian provinces, and provide empirical support for the hypothesis that closed-economy constraints may be partly responsible for the excess sensitivity phenomenon.

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