Financial Development, Consumption Smoothing, and the Reduced Volatility of Real Growth

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Preliminary version (please do not quote)

Abstract

Over the past quarter century, much of the developed and developing world has experienced a great moderation. Economic growth is more stable than it once was. Using data from 13 OECD countries, we analyze the role of financial development in relieving household liquidity constraints, thereby allowing for smoother consumption and less volatile growth. In the paper we begin by documenting the combined reduction in the volatility of both consumption and real growth, together with the increases in credit extended to the private sector.

Our data allows us estimate the time variation in the proportion of the population that is limited to consuming, at most, its current income (thus violating the life-cycle permanent income hypothesis). We then proceed to make the case for a causal relationship: Financial development increases access to credit markets, enabling households to smooth their consumption, which reduces the volatility of consumption and real growth.

JEL classification: E21, E32, E44

Keywords: Growth volatility, liquidity constraints, private credit

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1 Introduction

Over the past quarter century, much of the developed and developing world has experienced a great moderation. Economics growth is more stable than it once was. Figures 1a and 1b portray the overall reduction in the volatility of real GDP growth and real consumption growth, respectively, for a sample of 13 OECD countries. Using subperiods of 16 quarters for the time span between 1971:I and 2006:II, the data displays two main observations: First, the standard deviation of both output and consumption growth has steadily decreased (except for the 1979:I-1982:IV period, during the second major oil crisis). Also, all 13 countries converge to low volatility levels (ranging from 0.5% to 1.5%) in the 2003:I-2006:II period.

Figure 1a: Real GDP growth volatility: 1971:I-2006:II

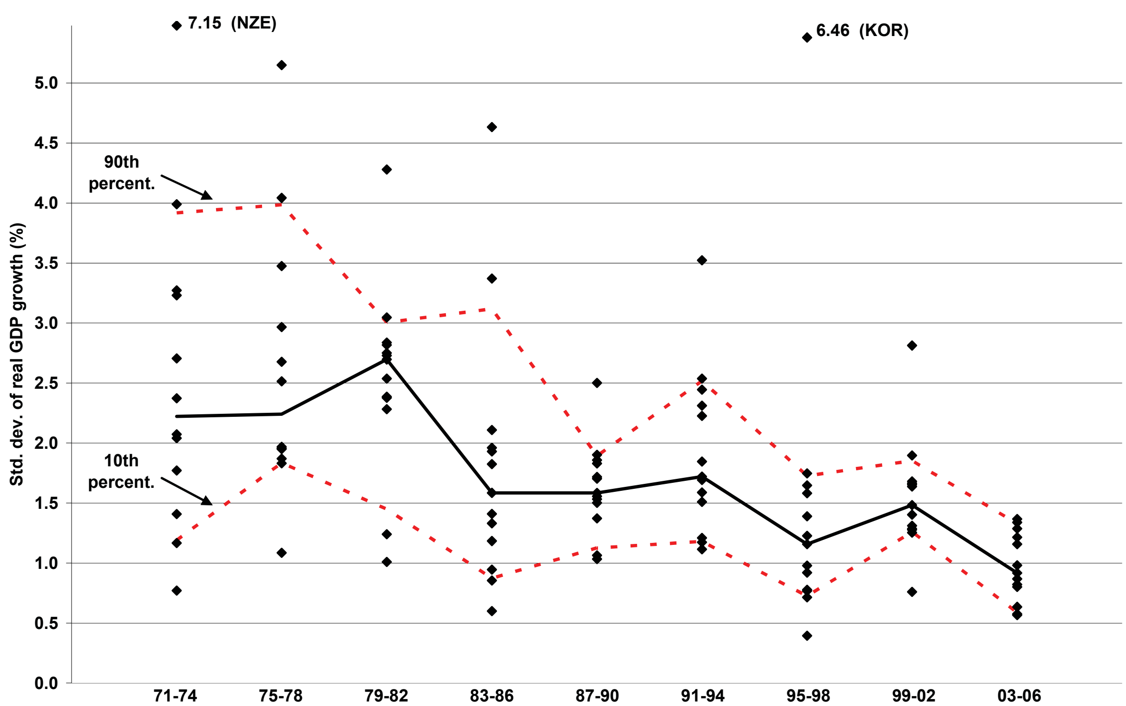
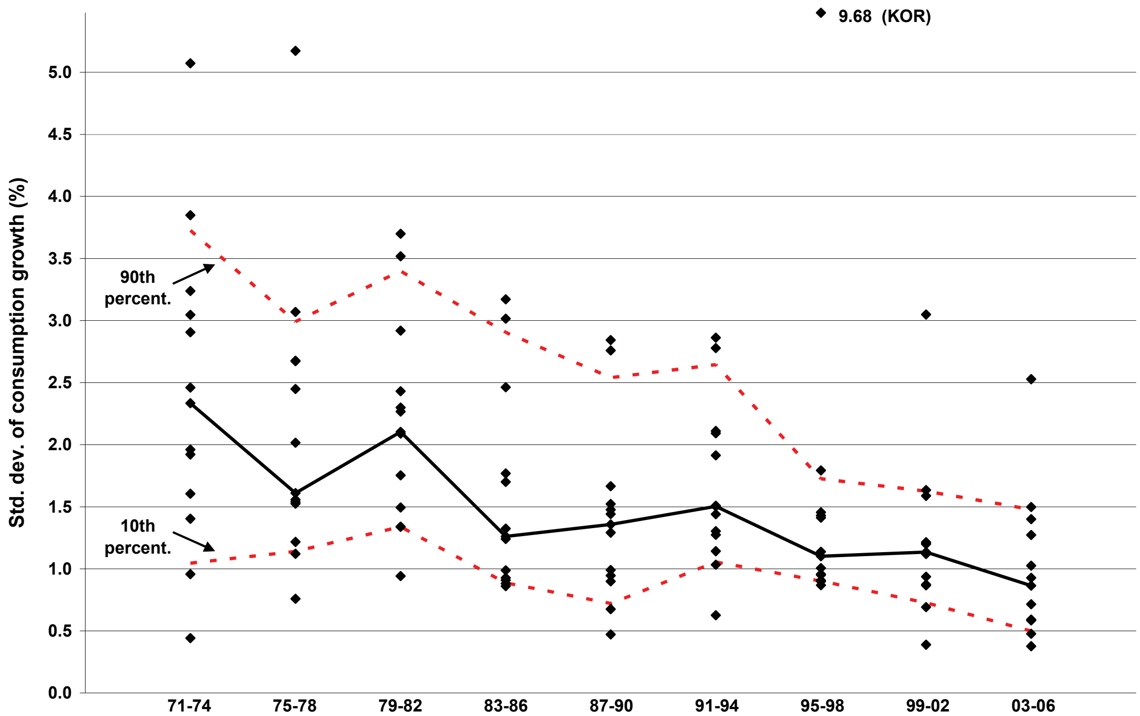


Figure 1b: Consumption growth volatility: 1971:I-2006:II



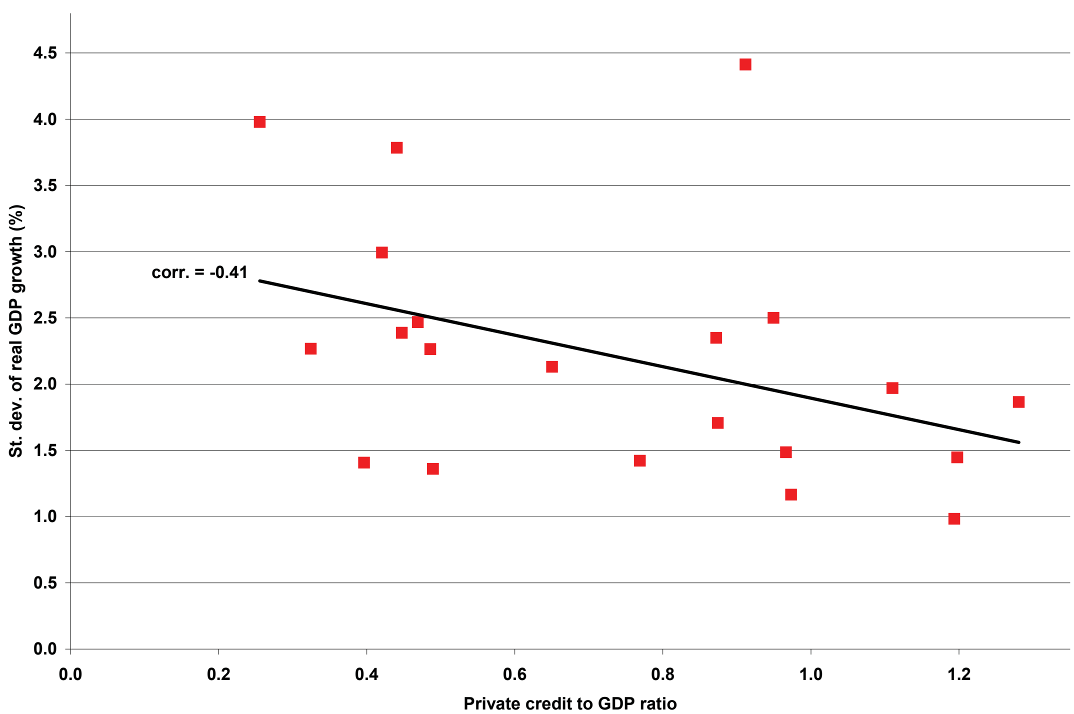
Recent work has looked at several potential explanations for this documented reduction in fluctuations. The list includes improved inventory management policies (McConnell and Pérez Quirós, 2000; Kahn, McConnell and Pérez Quirós, 2002; and McConnell and Kahn, 2005); more efficient monetary policy (Clarida, Galí and Gertler, 2000; Cecchetti, Flores-Lagunes and Krause, 2006); increased commercial openness (Barrell and Gottschalk, 2004); reduced impact of supply and demand shocks (Ahmed, Levin and Wilson, 2002; and Stock and Watson, 2002); and greater financial development, either in the form of financial innovation and improvements in risk sharing (Dynan, Elmendorf and Sichel, 2006), or through deeper financial markets (Denizer, Iyigun and Owen, 2002; and Beakaert, Harvey and Lundblad, 2006).[1]

[1] One exception is the work by Beck, Lundberg and Majnoni (2001), who find no robust link between financial development and reduced macroeconomic fluctuations when analyzing a sample of 63 countries.

In this paper we analyze in more detail the role that greater financial development has played in the changes in volatility of GDP. The preferred measure of financial development in the recent literature (Levine, Loayza, and Beck, 2000; and Levine, 2005) has been *private credit*, which includes the value of all credit that financial intermediaries issue to the private sector as a share of GDP. Aghion, Angeletos, Banerjee and Manova (2005) find that the effect of private credit in reducing growth volatility is significant and robust to various controls.

Employing the breaks in GDP growth volatility documented in Cecchetti, Flores-Lagunes and Krause (CFK, 2005), we compute the standard deviation of real growth and the average ratio of private credit to GDP for the subperiods before and after (and between the breaks, when multiple) for the 13 countries in our sample and plot them in Figure 2.[2] The correlation between these two variables is -0.41, very similar to the correlation of -0.48 estimated by Aghion, Angeletos, Banerjee and Manova (2005).

Figure 2: Real GDP growth volatility and private credit to GDP ratio



[2]CFK (2005) estimates the presence of (potentially) multiple structural breaks in GDP growth volatility for a cross section of countries, allowing for breaks in the mean and persistence of the series.

While the evidence suggests a stabilizing role of private credit on output fluctuations, it is less clear whether deeper financial markets have a direct impact on the volatility of GDP growth. Alternatively, financial development could either assist in stabilization by either providing a broader scope of action for monetary policy (Cecchetti and Krause, 2001; and Krause and Rioja, 2006), or allowing for smoother consumption by relieving household liquidity constraints. In this paper we concentrate on examining this latter channel. We directly estimate time variation in the proportion of the population that is liquidity constrained, that is, those who are limited to consuming, at most, its current income (thus violating the life-cycle permanent income hypothesis). We then proceed to make the case for a causal relationship: Increased access to credit enables households to smooth their consumption, which in turn reduces the volatility of consumption and real growth.

The remainder of the paper is organized as follows. Section 2 explains the method we employ to estimate the fraction of credit-constrained agents in the economy (λ). In Section 3 we analyze how λ is associated with the depth of financial intermediation markets; while Section 4 looks at how much changes in credit market access have contributed to smoother consumption. Section 5 concludes.

2 Estimating the Fraction of Credit-Constrained Agents

The usual approach in examining how financial development is related to output volatility looks at the following direct connection:

$$\uparrow pc \Longrightarrow \downarrow Stdev(g_y) \tag{1}$$

where pc represents credit extended to the private sector (as a fraction of GDP), while g_y is the real growth rate of output. Depending on the model specification, pc is typically instrumented by other variables (Aghion, Angeletos, Banerjee and Manova, 2005; and Beakaert, Harvey and Lundblad, 2006) to overcome potential endogeneity problems.

Our contention is that private credit allows for smoother consumption by relieving household liquidity constraints; in turn, smaller fluctuations in real consumption expenditures reduce the volatility of output growth. Formally:

$$\uparrow pc \Longrightarrow \downarrow \lambda \Longrightarrow \downarrow Stdev(g_c) \Longrightarrow \downarrow Stdev(g_y) \tag{2}$$

where λ represents the (income weighted) fraction of consumers who are credit constrained, and g_c is the real growth rate of consumption.

In order to study whether the transmission mechanism described in (2) is warranted, we need to generate estimates of λ. We follow the analysis performed by Campbell and Mankiw (1989, 1990) and Zeldes (1989) and assume the presence of two types of households. *Type 1* agents have limited access to credit markets and can only consume up to their present income, while *Type 2* agents can smooth out their consumption, consistent with the permanent income hypothesis. The fraction of total income that is received by Type 1 consumers is defined by λ.

For our analysis, we estimate the Campbell and Mankiw (CM, 1989) regression that allows for a general model, in which Type 1 agents are bound by the Keynesian consumption function, and Type 2 agents make their decisions based on intertemporal optimization. For each country this implies estimating the following equation:

$$\Delta c_t = \mu + \lambda \Delta y_t + \phi r_t + e_t, \tag{3}$$

where Δc is the four-quarter change in (log) consumption (i.e., the real growth rate of consumption, g_c); Δy represents the four-quarter change in (log) output (i.e., the real growth rate of output, g_y); and r is the ex-post real interest rate. A *priori*, we should expect λ to be bounded between 0 and 1, with a value closer to unity being indicative of a higher proportion of credit-constrained agents. The real interest rate directly affects consumption through its role of intertemporal substitution for agents with access to credit markets; and the sign of ϕ will depend on whether the substitution effect offsets the income effect.

Given that the disturbance term in (3) is likely to be correlated with Δy_t, valid instruments are necessary to consistently estimate λ. To that purpose, we instrument Δy_t by three of its own lags (excluding the immediately prior one), Δy_{t-2}, Δy_{t-3}, Δy_{t-4} as in CM (1989). In addition, we also introduce a new instrument to this literature in the index of domestic democratic capital (DDC), developed in Persson and Tabellini (PT, 2006).[3] Implicit in this approach is the assumption that the above set of variables impacts Δy directly; but the effect on Δc is only through the contemporaneous value of Δy.

In obtaining the estimates of λ presented in Table 1, we tried as much as possible to employ a general specification for all countries in our sample. However, statistical specification tests performed for each country (also shown in the table) indicate, not surprisingly, that some countries require slightly different instrument specifications.[4] We present the particular choice of instruments for each specific country in the first column of Table 1. In the second column we report the λ's resulting from estimating (3) for each of the OECD countries in our sample; instrumenting Δy using the set of instruments noted in column 1. The sample period used is 1973:I to 2004:IV.

[3]The DDC variable was constructed to correspond to the definition by PT (2006), assuming a depreciation rate of democratic capital of 1% per year. PT (2006) developed this index using data from the PolityIV Data Set (Marshall and Jaggers, 2002; updated in 2004) and relate it to economic growth. Their findings suggest that both current and lagged DDC are significant and robust explanatory variables for real GDP growth using a panel of 149 countries. This evidence gives support to employing DDC as an instrument for Δy to circumvent any potential endogeneity problems. In our estimates below, we employ the one-year lagged value of DDC as an instrument for Δy.

[4]In particular, we employ two commonly used specification tests in the context of instrumental variables estimation. The first reported test in Table 1, the "F-test of excluded instruments" gauges the relevance of the instruments. It is a significance F-test on the set of identifying instruments from the first-stage regression. Rejection of this test indicates that the set of identifying instruments are not "weak" in the sense of showing enough relationship with the variable they instrument for. The second test employed, the "Hansen Over-identifying Restrictions (OIR) Test", is a general misspecification test. A common interpretation of the rejection of this test in the present context is that the instruments are not "valid" in the sense that they are correlated with the error term of the second-stage equation. Therefore, failing to reject this test indicates that the instrument set used is "valid".

Table 1: Income-weighted fraction of credit constrained agents (λ)

Country	Instruments	Fraction of credit constrained agents, λ	F-test of excluded instruments	Hansen OIR test
Australia	Δy_{t-2}, Δy_{t-3}, DDC_{t-4}	0.714 *(0.00)*	9.61 *(0.00)*	3.93 *(0.14)*
Canada	Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, DDC_{t-4}	0.651 *(0.00)*	37.74 *(0.00)*	4.88 *(0.18)*
Denmark	Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, DDC_{t-4}	0.749 *(0.00)*	10.44 *(0.00)*	5.43 *(0.14)*
France	Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, DDC_{t-4}	0.594 *(0.00)*	45.05 *(0.00)*	2.59 *(0.46)*
Germany	Δy_{t-2}, Δy_{t-3}, DDC_{t-4}	0.816 *(0.00)*	23.05 *(0.00)*	5.26 *(0.07)*
Japan	Δy_{t-2}, Δy_{t-3}, Δy_{t-4}	0.771 *(0.00)*	62.10 *(0.00)*	5.25 *(0.07)*
Korea Rep.	Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, DDC_{t-4}	1.141 *(0.00)*	23.46 *(0.00)*	1.05 *(0.79)*
Netherlands	Δy_{t-2}, Δy_{t-3}, DDC_{t-4}	1.124 *(0.00)*	15.45 *(0.00)*	2.53 *(0.28)*
New Zealand	Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, DDC_{t-4}	0.920 *(0.00)*	5.56 *(0.00)*	6.81 *(0.08)*
Norway	Δy_{t-2}, Δy_{t-3}, DDC_{t-4}	0.747 *(0.01)*	7.53 *(0.00)*	3.84 *(0.15)*
Switzerland	Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, DDC_{t-4}	0.413 *(0.00)*	52.33 *(0.00)*	3.79 *(0.28)*
United Kingdom	Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, DDC_{t-4}	0.937 *(0.00)*	23.25 *(0.00)*	7.34 *(0.06)*
United States	Δy_{t-2}, Δy_{t-3}, Δy_{t-4}, DDC_{t-4}	0.590 *(0.01)*	35.93 *(0.00)*	2.46 *(0.48)*

Source: Own computations using data from OECD Main Economic Indicators, 2006 (10); OECD Economic Outlook 76; and International Financial Statistics. P-values computed from robust standard errors are in parentheses.

We observe in Table 1 that for the median country in our sample, 3 out of every 4 (income-weighted) agents has limited access to credit markets. Only for Korea and the Netherlands, do we obtain an estimate for λ above 1; for the remaining 11 countries, λ ranges between 0.41 (Switzerland) and 0.94 (the UK). Finally, in the last two columns we provide some statistical evidence as to the validity and relevance of our instruments: For all 13 countries the F-test of excluded-instruments is rejected at the 1% level indicating that the identifying instruments are relevant; while we cannot reject either at the 1% or 5% level the hypothesis that the instruments fail the over-identification restriction, suggesting the validity of the instrument set used to identify each country's parameters.

The estimated λ's in Table 1 are interesting in their own right, but we are more interested in studying how (and if) λ has changed over time, while also looking at the particular connection between financial development and the proportion of the population that has access to credit, which we pursue in the next section. We also explore the relative importance of changes in the fraction of credit-constrained agents on the documented decrease in consumption and output volatility, the topic of Section 4.

3 Financial Development and Access to Credit

To analyze how the depth of financial intermediation markets is associated with the fraction of credit constrained agents in the economy (λ) we estimate the evolution of the latter over time. To accomplish this task, we first perform 32-quarter rolling regressions of equation (3) to obtain time-varying estimates $\widehat{\lambda}_t$ for each individual country in our sample, using the same set of identifying instruments presented in Table 1. Unit root tests applied to both $\widehat{\lambda}$ and private credit to GDP ratio (pc) for each country strongly suggest the presence of a unit root in both variables.[5] Therefore, we proceed to estimate the following equation in differences:

$$\Delta\widehat{\lambda}_t = \alpha + \beta\Delta pc_t + \delta r_t + \varepsilon_t \ , \tag{4}$$

where $\Delta\widehat{\lambda}$ is the four-quarter difference of $\widehat{\lambda}_t$; Δpc is the four quarter difference in the private credit to GDP ratio;[6] and all other variables are defined as above. This equation is estimated

[5] Results for these tests are available upon request.

[6] We apply 4-quarter differences, instead of the standard one-lag differencing, to be consistent with equation (3). In addition, this allows us to circumvent any potential seasonality present in the data; particularly in the case of the private credit to GDP ratio.

for each country using ordinary least squares. We argue that the direction of the causality is that deeper financial markets (larger pc_t) reduce credit restrictions, λ (i.e., $\beta < 0$). We control for the relative price of credit: our prior is that a higher real interest rate should, overall, limit consumers' access to credit, thereby increasing λ ($\delta > 0$).

Table 2: Regression results: $\Delta\lambda = \alpha + \beta\Delta pc + \delta r + \varepsilon$

Country	Private credit to GDP (β)	Real interest rate (δ)	F-test (goodness of fit)	# of observations
Australia	-9.994 *(0.00)*	4.950 *(0.24)*	6.37 *(0.00)*	105
Canada	0.768 *(0.04)*	-0.256 *(0.56)*	3.11 *(0.05)*	89
Denmark	-2.409 *(0.05)*	-1.126 *(0.47)*	3.10 *(0.05)*	71
France	-0.689 *(0.07)*	-1.738 *(0.04)*	5.43 *(0.01)*	69
Germany	-2.276 *(0.05)*	-5.056 *(0.07)*	5.42 *(0.01)*	105
Japan	-0.970 *(0.10)*	1.750 *(0.22)*	2.12 *(0.16)*	105
Korea Rep.	-0.893 *(0.01)*	-1.538 *(0.12)*	4.74 *(0.01)*	81
Netherlands	2.106 *(0.15)*	11.569 *(0.03)*	14.39 *(0.00)*	56
New Zealand	3.269 *(0.02)*	6.083 *(0.02)*	8.29 *(0.00)*	49
Norway	-0.519 *(0.74)*	2.059 *(0.50)*	0.25 *(0.78)*	101
Switzerland	2.599 *(0.00)*	-2.383 *(0.05)*	5.00 *(0.01)*	69
United Kingdom	-2.340 *(0.05)*	1.031 *(0.85)*	2.51 *(0.09)*	59
United States	-1.304 *(0.42)*	-0.611 *(0.34)*	0.73 *(0.49)*	105
Panel	**-0.964** ***(0.05***	**0.787** ***(0.27)***	**2.24** ***(0.10)***	**1064**

Source: Own computations using data from OECD Main Economic Indicators, 2006 (10); OECD Economic Outlook 76; and International Financial Statistics. P-values computed from robust standard errors are in parentheses.

The results are reported in Table 2. For each country, the number of available time periods to estimate (4) depends on data availability on each of the variables used. As a result, different sample sizes are available for each country, indicated in the last column of the table. We find that the estimated β's are negative (as predicted) and statistically significant at the 10% level for 7 countries: Australia, Denmark, France, Germany, Japan, Korea, and the UK. Only for Canada, New Zealand and Switzerland do we find positive and statistically significant coefficients for β, while the remaining 3 countries have statistically insignificant coefficients. As for our control variable, the real interest rate does not appear to be positively related to λ for most countries.

Clearly, other elements such as particular financial institutions, arrangements between banks and costumers, demographics, income distribution, and many more could determine the population's access to credit market (or changes thereof), and these factors are likely to be specific to each individual country. Therefore, it is useful to control for country specific effects and test our hypotheses. The last row of Table 2 reports the results for an unbalanced panel. Controlling for fixed effects, we find evidence that the common β is negative and largely significant; while the real interest rate has a positive, albeit insignificant, impact on λ. It is also worthwhile to note that the effects are quite substantial: a 1% increase in the ratio of private credit to GDP reduces the fraction of credit constrained agents by nearly 1%.

Clearly, these results only provide suggestive evidence as to the first part of (2) - deeper financial markets are associated with increased access to credit markets. In the next section we analyze how important the role of lower credit constraints is in reducing fluctuations in consumption.

4 Explaining the Volatility of Consumption Growth

Our previous results enable us to now test our main hypothesis: Can the fraction of credit constrained agents in the economy (λ) explain the volatility of consumption and output growth? If so, How much has consumer access to credit markets (via a greater extent of financial depth) contributed to reducing fluctuations in output via smoother consumption?

We accomplish this task through a number exercises that rely on different assumptions that are needed to interpret the results. Importantly, in each of these alternative exercises we arrive to the same conclusion that supports our hypothesis.

The first exercise we undertake consists in using the subperiods provided by the GDP volatility breaks from CFK (2005). We compute the volatility of consumption and estimate λ in (3) in each of those subperiods for each of the countries in our sample that present such breaks. We then relate the changes in consumption volatility with the changes in λ for each country. We report these results in Table 3a.

Out of the 13 countries, 7 experienced at least one significant break in output volatility between 1973 and 2004. Considering first the countries with only one break, Australia, Denmark, Germany and the US experienced both a reduction in consumption volatility, measured by the standard deviation of consumption growth, $Stdev(g_c)$, and a decrease in the fraction of credit constrained agents, λ, consistent with (2). Meanwhile, Canada experienced a lower $Stdev(g_c)$, with no perceptible change in λ. As for the countries with multiple breaks, the estimates for the Netherlands suggest, after the first break, a fall in $Stdev(g_c)$, accompanied by an increase in λ; while after the second break, both $Stdev(g_c)$ and λ went up, consistent with (2). Finally, the UK experienced a steady decrease in the volatility of consumption over the three subperiods, with the estimate for λ first rising and then falling.

Thus, only the results for the periods separated by the first break, in the case of both the Netherlands and the UK, contradict our hypothesis of a positive relationship between the fraction of credit constrained agents and consumption volatility. Plotting the data from Table 3a in Figure 3, we find supporting evidence of a significantly positive correlation ($\rho = 0.52$)

between $Stdev(g_c)$ and λ.

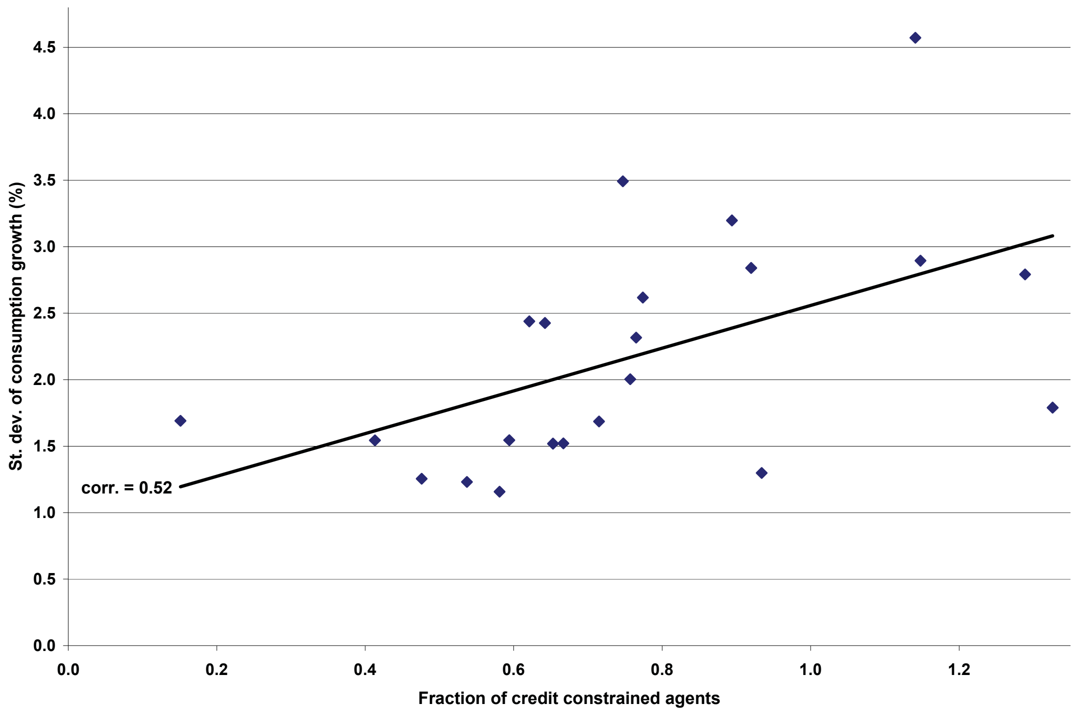
Table 3a: λ and Consumption growth volatility (breaks)

Country	Period	Qtr-beg.	Qtr.-end	Consumption Volatility (σ_C, in %)	Fraction of credit constrained agents, λ [st. dev in brackets]
Australia	*1st*	1971Q1	1984Q3	1.686	0.715 [0.348]
	2nd	1984Q4	2004Q4	1.565	0.667 [0.138]
Canada	*1st*	1975Q1	1987Q2	2.439	0.621 [0.104]
	2nd	1987Q3	2004Q4	1.553	0.653 [0.079]
Denmark	*1st*	1972Q1	1994Q3	2.894	1.148 [0.202]
	2nd	1994Q4	2004Q4	1.514	0.151 [0.242]
France	*entire*	1971Q1	2004Q4	1.576	0.594 [0.098]
Germany	*1st*	1971Q1	1993Q3	2.004	0.757 [0.141]
	2nd	1993Q4	2004Q4	1.292	0.476 [0.256]
Japan	*entire*	1971Q1	2004Q4	0.353	0.765 [0.071]
Korea Rep. *	*1st*	1977Q1	2004Q4	5.037	1.141 [0.115]
Netherlands	*1st*	1971Q1	1983Q4	2.618	0.774 [0.125]
	2nd	1984Q1	1994Q3	1.298	0.934 [0.224]
	3rd	1994Q4	2004Q4	1.717	1.326 [0.182]
New Zealand *	*entire*	1985Q1	2004Q4	3.536	0.920 [0.140]
Norway	*entire*	1972Q1	2004Q4	2.892	0.747 [0.245]
Switzerland	*entire*	1972Q1	2004Q4	1.570	0.413 [0.048]
United Kingdom	*1st*	1972Q1	1981Q2	3.198	0.894 [0.090]
	2nd	1981Q3	1991Q4	2.792	1.289 [0.095]
	3rd	1992Q1	2004Q4	1.191	0.537 [0.292]
United States	*1st*	1971Q1	1984Q2	2.426	0.642 [0.083]
	2nd	1984Q3	2004Q4	1.196	0.581 [0.101]

Source: Own computations using data from OECD Main Economic Indicators, 2006 (10); OECD Economic Outlook 76; and International Financial Statistics. GDP growth volatility break dates are from CFK (2005).

* For Korea Republic and New Zealand, even though CFK (2005) do identify breaks in the volatility of GDP growth (1980Q3 for Korea; and 1975Q3 and 1987Q3 for New Zealand), data on the money market interest rate only goes back to 1977:I and 1985:I, respectively. Therefore, we are unable to estimate λ prior to 1980 for Korea, and 1985 in the case of New Zealand.

Figure 3: Consumption volatility and fraction of credit-constrained agents



We now turn our attention to the 6 countries for which either no volatility break was identified (France, Japan, Norway and Switzerland), or data limitations do not allow us to use the break periods to define subperiods (Korea and New Zealand). For each of these countries, we simply split the sample into two equal parts, and estimate λ for each of the subperiods.

The estimates in Table 3b strengthen our previous findings: In all 5 countries that experienced a reduction in consumption volatility, the fraction of credit-constrained agents also fell; meanwhile, the estimate of λ rose in the case of Korea, which experienced a sharp increase in $Stdev(g_c)$. Including these observations also increases the correlation between $Stdev(g_c)$ and λ from 0.52 to 0.63.

Table 3b: λ and Consumption growth volatility (equal sample split)

Country	Period	Qtr-beg.	Qtr.-end	Consumption Volatility (σ_C, in %)	Fraction of credit constrained agents, λ [st. dev in brackets]
France	*1st*	1971Q1	1987Q4	1.756	0.653 [0.127]
	2nd	1988Q1	2004Q4	1.255	0.568 [0.082]
Japan	*1st*	1971Q1	1987Q4	2.437	0.924 [0.096]
	2nd	1988Q1	2004Q4	1.799	0.277 [0.092]
Korea	*1st*	1977Q1	1990Q4	2.653	0.515 [0.174]
	2nd	1991Q1	2004Q4	6.420	1.395 [0.084]
New Zealand	*1st*	1985Q1	1994Q4	2.917	0.696 [0.230]
	2nd	1995Q1	2004Q4	1.653	0.089 [0.414]
Norway	*1st*	1972Q1	1988Q2	3.709	1.474 [0.651]
	2nd	1988Q3	2004Q4	1.799	0.354 [0.159]
Switzerland	*1st*	1972Q1	1988Q2	1.817	0.434 [0.066]
	2nd	1988Q3	2004Q4	1.041	0.349 [0.091]

Source: Own computations using data from OECD Main Economic Indicators, 2006 (10); OECD Economic Outlook 76; and International Financial Statistics.

As a quick verification exercise, we note that the estimate of λ we obtain for the US for the subperiod 1971:Q1-1984:Q2 (0.642) is very similar to the value for the fraction of households with $\leq$ 2 months' average income, reported by Zeldes (1989) ($\lambda_z = 0.67$), and the estimates of CM (1989) when instrumenting income with lagged changes in the 3-month T-bill rate ($\lambda^1_{CM} = 0.698$; $\lambda^2_{CM} = 0.657$; depending on the model specification).

The evidence presented in Tables 3a and 3b is only suggestive, given that only two (at most three) subperiods are used to relate λ with the volatility of consumption. Therefore, as a second exercise to determine how much of the generalized decrease in consumption volatility can be explained by the agents' ability to smooth their expenditures, we turn to a variance decomposition approach. To accomplish this task, consider the fitted version of equation (3):

$$g_c \equiv \Delta c_t = \mu + \lambda \widehat{\Delta y}_t + \phi r_t + \epsilon_t, \tag{5}$$

where $\widehat{\Delta y}$ is the instrumented value for Δy in the first stage regression. Taking variances on both sides of (5)

$$\sigma_c^2 = \lambda^2 \sigma_{\widehat{y}}^2 + \phi^2 \sigma_r^2 + 2\phi\lambda\sigma_{\widehat{y},r} + \sigma_\epsilon^2, \tag{6}$$

where $\sigma_c^2 \equiv Var_t(g_c)$; $\sigma_{\widehat{y}}^2 \equiv Var_t(\widehat{\Delta y})$; $\sigma_r^2 \equiv Var_t(r)$; $\sigma_{\widehat{y},r} \equiv Cov_t(\widehat{\Delta y}, r)$; and $\sigma_\epsilon^2 \equiv Var_t(\epsilon)$. The above equation is well defined provided that $|\lambda| < 1$.

We now turn to the data to establish the relative importance of the changes in λ towards smoother consumption. We perform a comparative statics exercise, which is explained in the following steps:

- First, we employ as the original values for λ the estimates of period 1 from Tables 3a and 3b for all countries; except the Netherlands and the UK (the two countries with multiple breaks), for which we use the estimates of period 2 as our baseline; and define them as λ_1. Accordingly, the estimates of λ for period 2 (period 3, in the case of Netherlands and the UK) are defined as λ_2. Since (6) is only well defined whenever $|\lambda| < 1$, for the countries and subperiods where $\widehat{\lambda} > 1$ we impose a value of $\lambda = 0.9999$.

- Second, we estimate ϕ, $\sigma_{\widehat{y}}^2$, $\sigma_{\widehat{y},r}$, and σ_r^2 for the entire period for each of the 13 countries; and fix them for the entire analysis.

- Next, we calibrate σ_ϵ^2, so that it fits with period 1 (period 2, for the Netherlands and the UK) estimates of σ_c^2; the λ_1's; $\sigma_{\widehat{y}}^2$, $\sigma_{\widehat{y},r}$, and σ_r^2. We then maintain σ_ϵ^2 fixed for the duration of the analysis.

- Finally, we determine the change in σ_c that is due solely to a change in the fraction of credit-constrained agents from λ_1 to λ_2; and compare it to the *actual* change in σ_c between the two periods in question. This enables us to compute the relative contribution of the change in λ to the overall change in σ_c.

Table 4: Relative contribution of changes in λ to changes in σ_C

Country	Change in λ *	Actual change in σ_C	Change in σ_C due to change in λ	Relative contribution of change in λ
Australia	-0.048	0.121	0.087	71.6%
Canada	+0.033	0.630	-0.043	-6.8%
Denmark	-0.849	1.236	0.854	69.1%
France	-0.085	0.501	0.068	13.7%
Germany	-0.281	0.711	0.347	48.8%
Japan	-0.646	0.638	1.236	193.7%
Korea Rep.	+0.485	-3.767	-1.636	43.4%
Netherlands	+0.066	-0.419	-0.097	23.1%
New Zealand	-0.646	1.910	0.642	33.6%
Norway	-0.607	1.264	0.687	54.3%
Switzerland	-0.085	0.776	0.106	13.6%
United Kingdom	-0.463	1.600	0.292	18.3%
United States	-0.060	1.230	0.077	6.3%

Source: Own computations using data from OECD Main Economic Indicators, 2006 (10); OECD Economic Outlook 76; and International Financial Statistics.

* For estimates of **λ** > 1, we impose a value of **λ** = 0.9999.

The results of this exercise are reported in the last two columns of Table 4. For 10 out of the 13 countries, a smaller λ has contributed to reduced fluctuations in consumption; the measured relative contribution ranges from 6% (US); a median of 34% (New Zealand), all the way to over 100% (Japan). For the two countries which experienced higher volatility in the latter subperiod (Korea and the Netherlands), the contribution of a higher λ was 43%

and 23%, respectively. Only in the case of Canada, the slight increase in λ would predict more volatile consumption, instead of the observed reduction in σ_c.

Overall, these findings are suggestive that better access to credit, through deeper financial markets, has effectively reduced the volatility of consumption, which in turn - given the importance of consumer expenditures in GDP - has contributed towards more stable real growth.

5 Conclusions

Financial development is regarded as one of the factors contributing towards a reduction in the volatility of output growth. In this paper, we empirically examine one potential transmission mechanism in which deeper financial markets can result in the moderation of business cycles; namely, through consumption smoothing. Looking at evidence from 13 OECD countries, we find evidence suggesting that an increase in private lending alleviates credit restrictions on consumers. In turn, this better access to credit markets has allowed consumers to smooth their expenditures, resulting in more stable consumption growth.

6 Data Appendix

We collected data for 13 OECD countries; most of them between 1970:I-2006:II. A description of the variables used is presented in Appendix Table 1. As for the sources, data on *GDP*, *consumption*, *CPI*, and *Private Credit / GDP ratio* were obtained OECD Main Economic Indicators, 2006 (10) and the OECD Economic Outlook 76. Information on *money market interest rates*, comes from the International Financial Statistics 2006, Issue 10. Finally, *Domestic Democratic Capital* was constructed based on the index developed by Persson and Tabellini (2006), who employ data from PolityIV (Reference: Monty G. Marshall and Keith Jaggers. 2002. Polity IV Dataset. [Computer file; version p4v2002] College

Park, MD: Center for International Development and Conflict Management, University of Maryland, 2004 update)

Table A.1: Data Description

Country	Real GDP	Real Cons.	Real Int. Rate	Private Credit	Dem. Capital
Australia	1970Q1	1970Q1	1970Q1	1976Q3	1970
	2006Q2	2006Q2	2006Q2	2006Q2	2003
Canada	1970Q1	1970Q1	1975Q1	1970Q1	1970
	2006Q2	2006Q2	2006Q2	2006Q2	2003
Denmark	1978Q1	1970Q1	1972Q1	1978Q1	1970
	2006Q2	2006Q2	2006Q2	2006Q2	2003
France	1970Q1	1970Q1	1970Q1	1987Q4	1970
	2006Q2	2006Q2	2006Q2	2006Q2	2003
Germany	1970Q1	1970Q1	1970Q1	1970Q1	1970
	2006Q2	2006Q2	2006Q2	2006Q2	2003
Japan	1970Q1	1970Q1	1970Q1	1970Q1	1970
	2006Q2	2006Q2	2006Q2	2006Q2	2003
Korea Rep.	1970Q1	1970Q1	1977Q1	1970Q1	1970
	2006Q2	2006Q2	2006Q1	2006Q2	2003
Netherlands	1970Q1	1970Q1	1970Q1	1991Q1	1970
	2006Q2	2006Q2	2006Q2	2006Q1	2003
New Zealand	1970Q1	1970Q1	1985Q1	1984Q2	1970
	2006Q1	2006Q1	2006Q2	2006Q1	2003
Norway	1970Q1	1970Q1	1972Q1	1970Q1	1970
	2006Q2	2006Q2	2006Q2	2006Q2	2003
Switzerland	1970Q1	1970Q1	1972Q1	1987Q4	1970
	2006Q1	2006Q1	2006Q2	2006Q1	2003
United Kingdom	1970Q1	1970Q1	1972Q1	1990Q2	1970
	2006Q2	2006Q2	2006Q2	2006Q2	2003
United States	1970Q1	1970Q1	1970Q1	1970Q1	1970
	2006Q2	2006Q2	2006Q2	2006Q2	2003

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