On the Role of Domestic and International Financial Cyclical Factors in Driving Economic Growth

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Abstract

We investigate the effects of (domestic and international) financial cyclical factors on the US business cycle over the period 1890-2013 using an augmented stochastic version of the neoclassical growth model. In our setting, financial factors enter as determinants of the total factor productivity cyclical pattern. By means of static and dynamic estimations we find that (*i*) the inclusion of financial cyclical factors improves the model's performance; (*ii*) the sensitivity of economic growth to financial factors is time-varying; (*iii*) domestic financial factors have a key role in explaining short-run output fluctuations only in the first half of the 20th century; (*iv*) the US business cycle is found to be mainly driven by global financial factors (i.e., financial integration) over the last three decades.

JEL CODES: O40, E32, C32

Key Words: Stochastic Growth Model, Cyclical Fluctuations, Financial Factors, State-Space Model

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"*One of the most important problems in the field of finance, if not the single most important one,...is the effect that financial structure and development have on economic growth.*"

Goldsmith (1969)

1 Introduction

The relationship between finance and growth is still at the center of the academic and policy debate. Since the seminal work of Gurley and Shaw (1955),[1] several empirical studies have supported the hypothesis that financial development promotes economic growth (see, among others, King and Levine, 1993a; Rousseau and Wachtel, 1998; Beck et al., 2000; Amano, 2005). Most recent studies have also found that there are strong linkages between financial factors and macroeconomic aggregates (Goodhart and Hofmann, 2008; Tsouma, 2009; Haavio, 2012).

The existing empirical works on the relationship between finance and growth have been dominated by cross-country analyses until recently due to the lack of very long term time series data. In this respect, cross-sectional analyses tend to deliver results that rely on a relatively short period. Moreover, cross-country regressions do not account for time-variation in the finance-growth relationship and exhibit additional non-negligible drawbacks (Demetriades and Hussein, 1996; Arestis and Demetriades, 1997).[2] Differently, the number of works investigating the relationship between financial development and economic growth in a country-by-country setting by means of time-series regressions is rather scarce, in particular when it comes to examine the determinant of growth in the pre-war period (i.e., before the '50s). However, the existing time series-based works exhibit a couple of weaknesses. First, they rely exclusively on empirical approaches and do not have any theoretical support. Actually in the theory of growth, financial development (or, more in general, financial variables) may affect economic growth improving the productivity of capital. To this end,

[1]The authors argue that the development of financial institutions boosts households' savings and firms' investments. This, of course, stimulates the real economic activity.

[2]The most important limitation is that using cross-country regressions the investigator is able to estimate only the so-called "average effect" described in Arestis and Demetriades (1997). See also Demetriades and Arestis (1996) on this point.

it is important also to understand whether these effects are only transitory ("level effect") or permanent ("growth effect"). Second, the international dimension of financial development (i.e., global financial integration) has not yet been considered in the literature. Said differently, existing studies have employed mainly domestic financial factors.

In this paper we focus on the effects of financial development and domestic and international financial factors on the US economy. More importantly, we examine whether these effects are time-varying. To do so, we employ an empirical framework based on a neoclassical growth model and more than 100 years of data. This analysis is motivated by the following facts: (*i*) the more recent empirical literature emphasizes that the shape of the effect of financial variables on economic growth is not constant over time (see Rousseau and Wachtel, 2011; Beck et al., 2014; Bezemer et al., 2016; Amano, 2005, 2013); (*ii*) country-by-country time-series analyses show that the relationship between finance and growth differs across countries (see Shan et al., 2001; Amano, 2005, 2013; Peia and Roszbach, 2015);[3] (*iii*) empirical facts suggest that (*a*) the proposed empirical framework is suitable for the country and the period under investigation and (*b*) financial variables tend to heavily influence economic growth.

We differ from existing studies on the finance-growth nexus in several respects. First, motivated by the empirical evidence on the US economy, we let both domestic and international financial factors be potential drivers of economic growth. In particular, as domestic financial factors we use: (*i*) the domestic short-term interest rate and (*ii*) a financial development proxy (defined as stock market transaction plus money supply normalized by total population). The world interest rate is instead employed to capture the international dimension of financial development (i.e., financial integration). This choice is motivated by the increasing trend in the degree of financial integration process observed in the last decades among developed economies (see Figure B.1, Appendix B). We therefore assume that countries today may be more sensitive to international financial shocks than to local financial

[3]This result is consistent, for example, with Luintel and Khan (2004) who show lack of correspondence between panel and country-specific estimates; hence, the generalizations based on panel results may offer incorrect inferences for several countries of the panel. Likewise, Cooray et al. (2013) find that even in presence of countries of the same region, the growth determinants differ from one country to another. The single-country approach in growth regression is also consistent with Pack (1994) and Commission on Growth (2010) suggestions.

ones. Second, we develop a novel theoretical framework that reflects key empirical facts and has the nice property to collapse into a state-space representation. Specifically, in the spirit of Lee et al. (1997), we develop a stochastic version of the exogenous growth model where financial factors – expressed in deviations from their long-run trend – are drivers of the TFP cyclical pattern.[4] Third, we estimate our empirical model in a pure time-varying context. This allows us to investigate whether the structure of the finance-growth nexus has changed over time. In this respect, we also estimate several different specifications of the model. Fourth, a dynamic model selection procedure (using Akaike weights) is executed to investigate which model (and then financial variable) is "statistically better" in explaining output growth.

Our US-based analysis produces several novel empirical insights. First, we provide empirical support for the existence of an exogenous constant growth rate of per capita GDP. Via the Bai-Perron algorithm, a test on the mean-shift of US GDP per capita growth reveals zero breaks in the time series.[5] This implies that the so-called level effect operates for the US. Second, the cyclical pattern of financial variables (i.e., financial market development index, domestic interest rate and world interest rate) plays an important role in explaining growth dynamics over the years. Third, we find that the role played by these financial cyclical components varies over time. More importantly, our empirical findings suggest that international financial factors (i.e., financial integration) contributed most to explaining output growth in the US over the last 30 years. We argue that the observed heterogeneity and time-varying components can lead to uncertainty around the model that has to be selected to capture long-term growth dynamics.

Despite its complexity, we believe that our framework represents a useful tool to get a better understanding of the role of financial cyclical factors in the growth rate of the economy. Most broadly, our results suggest that policymakers should not exclude from their analysis the cyclical phase of exogenous financial factors once a new policy measure is

[4]Modeling the finance-growth transmission channel allows for a very flexible setup where one can exploit the role of additional financial aggregates. Theoretical endogenous models that capture different channels to which financial factors influence the real GDP, have been developed, among others, by Bencivenga and Smith (1991), King and Levine (1993b), and Blackburn and Hung (1998).

[5]This result is also in line with more recent empirical analyses (see, for instance, Papell and Prodan, 2014; Sobreira et al., 2014).

introduced. This because medium-run effects are state-dependent. Finally, we stress that our augmented stochastic exogenous growth framework gives rise to a methodology that can be easily employed – given the availability of a large number of long-term macroeconomic series – to study historical growth dynamics in many other countries.[6]

The rest of paper is organized as follows. Section 2 reviews the empirical (and theoretical) literature on the finance-growth nexus. Section 3 shows some empirical regularities for the US over the period 1870-2013. The augmented-stochastic neoclassical growth model and the related empirical strategy are presented in sections 4 and 5, respectively. Section 6 shows and discusses the empirical results for different model specifications. Section 7 concludes.

2 Literature review

Our paper fits into a long-standing and still growing literature that examines the implications of the financial system dynamics for economic growth. In this section, we review empirical and theoretical studies investigating the relationship between financial development and growth. Existing empirical works on the finance-growth nexus are classified into two different groups: (*i*) "cross-country" analyses (Section 2.1) and (*ii*) "country-by-country" time-series analyses (Section 2.2). The first group comprises all papers using cross-country growth regression, whereas the second one considers only papers studying the finance-growth relation from a single-country perspective in a time-series framework. Our paper is also more distantly related to the theoretical literature on business and financial/credit cycles. We (briefly) review this literature in Section 2.3.

2.1 Finance and growth: cross-country analyses

The cross-country growth literature investigating the effect of finance on growth is rather vast. A non exhaustive list of works and their related main results are summarized in Table 1.[7]

[6] See, among others, the Piketty and Zucman (2014)' database available at http://http://gabriel-zucman.eu/capitalisback/.

[7] Given the massive number of works produced in this literature, we have decide to focus exclusively on papers published in the last 30 years and studying the effect of finance on output mainly within high-income

One of the earliest paper is Kormendi and Meguire (1985) who relate the growth rate of 47 countries (for the period 1950-1977) to a bunch of macroeconomic and financial aggregates: initial per capita income, the population growth rate, the output volatility, money growth volatility, and money supply growth, the growth of government spending and of export, and inflation. The authors show that a rise in the money supply growth volatility reduces output growth. Most importantly, this "seems to be the largest single contributing factor in explaining growth." In a cross-country analysis setting and by relying on a group of 32 developing countries for the period 1961-1988, Fischer (1993) shows that growth is negatively related to the black market exchange premium (a measure of exchange market distortion).

A first attempt to develop an empirical framework aimed at investigating the relationship between finance and growth is due to King and Levine (1993a). In this seminal paper, the authors use four different financial development proxies.[8] By means of cross-country regressions King and Levine (1993a) find that the coefficient of all four indicators of financial development is positive and highly statistically significant, suggesting a positive relationship between financial development and growth. Levine and Zervos (1998) study the empirical relationship between various measures of stock market development, banking development, and growth. The authors find that stock market liquidity (i.e., stock turnover and value traded) and market capitalization influence positively economic growth. Based on the empirical evidence, Levine and Zervos (1998) argue that "financial factors are an integral part of the growth process." Rajan and Zingales (1998) support the empirical evidence of a robust, positive, causal effect of finance on industry growth.[9] In particular, Rajan and Zingales (1998) find that the growth rate of industries with a greater reliance on external financing is higher in countries with more public disclosures. Beck et al. (2000) use liquid liabilities

countries. We have therefore excluded from the review important papers such as Christopoulos and Tsionas (2004) who focus exclusively on developing countries.

[8]The first is liquid liabilities of financial institutions as share of GDP (i.e., a measure of the size of financial intermediaries). The second is the ratio of bank credit to the sum of bank and central bank credit (i.e., a measure of degree to which banks versus the central bank allocate credit). The third is the ratio of private credit to total credit, and the fourth is private credit as a share of GDP (i.e., measures of the extent to which banking sector funds to the private sector).

[9]The primary variable of interest in Rajan and Zingales (1998) is represented by the interaction between external dependence of industry i and financial development. The interaction variable tests how the sectors (who needs of external finance) grow given the level of financial development. If the sign is positive, as expected, it means that the sectors that are more dependent on external finance grow thanks to the help of financial development.

(divided by GDP), private credit (divided by GDP), and the ratio of bank credit to the sum of bank and central bank credit as proxies of financial development indicators. The authors find a positive and statistically significant relation between the proposed financial development measures and real GDP per capita.

Table 1: The Finance-Growth Nexus: Cross-country Analyses.

Authors	Fin. Var.	Dep. Var.	Panel	Estimation Technique	Results
Kormendi and Meguire (1985)	ΔM, σ_M.	ΔY	47 countries, 1950-1977.	Cross-section regression	$\sigma_M \overset{(-)}{\rightarrow} \Delta Y$
Fischer (1993)	BMP	ΔY	32 developing countries, 1961-1988	Cross-section and panel regression	$BMP \overset{(-)}{\rightarrow} \Delta Y$
King and Levine (1993a)	LLY, $BANK$, $PRIV$, $PRIVY$	ΔY	77 countries, 1960-1989	Cross-section and pooled cross-country regression	$LLY \overset{(+)}{\rightarrow} \Delta Y$, $BANK \overset{(+)}{\rightarrow} \Delta Y$, $PRIV \overset{(+)}{\rightarrow} \Delta Y$, $PRIVY \overset{(+)}{\rightarrow} \Delta Y$
Levine and Zervos (1998)	CAP, $TURN$, VTR, $WINT$, VOL, $PRIVL$	Δy	47 countries, 1976-1993	Cross-section regression	$PRIV \overset{(+)}{\rightarrow} \Delta y$, $TURN \overset{(+)}{\rightarrow} \Delta y$, $VTR \overset{(+)}{\rightarrow} \Delta y$, $CAP \overset{(+)}{\rightarrow} \Delta y$
Rajan and Zingales (1998)	$CCAPY$, DIN, $LEND$	ΔY_i	41 countries, 1980-1990	Cross-section regression	$CCAPY \overset{(+)}{\rightarrow} \Delta Y_i$, $DIN \overset{(+)}{\rightarrow} \Delta Y_i$, $LEND \overset{(+)}{\rightarrow} \Delta Y_i$
Beck et al. (2000)	$PRIV$, LLY, $BANK$	Δy	63 (cross-section) and 77 (panel regression) countries, 1960-1995	Cross-section regression and panel regression	$PRIV \overset{(+)}{\rightarrow} \Delta y$, $LLY \overset{(+)}{\rightarrow} \Delta y$, $BANK \overset{(+)}{\rightarrow} \Delta y$
Rousseau and Wachtel (2001)	$M3Y$, $M3LY$, $CRED$	ΔY	84 countries, 1960-1995	Panel regression	Only in periods of medium or low inflation: $M3Y \overset{(+)}{\rightarrow} \Delta Y$, $M3LY \overset{(+)}{\rightarrow} \Delta Y$, $CRED \overset{(+)}{\rightarrow} \Delta Y$
Shen and Lee (2006)	$PRIV$, LLY, SPR, $VTRY$, $TURNC$, $CAPY$	Δy	48 countries, 1976-2001	Panel regression	$VTRY \overset{(+)}{\rightarrow} \Delta y$, $CAPY \overset{(+)}{\rightarrow} \Delta y$, $TURNC \overset{(+)}{\rightarrow} \Delta y$
Apergis et al. (2007)	$M3Y$, $CRED$	y	65 countries, 1975-2000	Panel cointegration	$M3Y \overset{(+)}{\leftrightarrow} \Delta y$, $CRED \overset{(+)}{\leftrightarrow} \Delta y$
Rousseau and Wachtel (2011)	$M3Y$, $M3LY$, $CRED$	Δy	84 countries, 1960-2004	Cross-section regression and panel regression	Only for 1960-1989: $M3Y \overset{(+)}{\rightarrow} \Delta y$, $M3LY \overset{(+)}{\rightarrow} \Delta y$, $CRED \overset{(+)}{\rightarrow} \Delta y$
Beck et al. (2014)	$VADD$, $CRED$	Δy	77 countries, 1980-2007	Cross-section regression and panel regression	Mainly for 1995-2007 and for high-income countries: $VADD \overset{(+)}{\rightarrow} \Delta y$
Chortareas et al. (2015)	$PRIV$, FOP	y	37 countries (20 advanced), 1970-2007	Panel cointegration	Exclusively for high-income countries: $PRIV \overset{(+)}{\rightarrow} y$, $FOP \overset{(+)}{\rightarrow} y$
Bezemer et al. (2016)	$CRED$, $\Delta CRED$	Δy	46 countries, 1990-2011	Panel regression	$CRED \overset{(-)}{\rightarrow} \Delta y$ (especially for real estate credit), $\Delta CRED \overset{(+)}{\rightarrow} \Delta y$ (especially for nonfinancial credit)

Notes: ΔM = Money supply growth; σ_M = standard deviation of money supply shocks; BMP = Black market exchange rate premium; LLY = Ratio of liquidity liabilities to GDP; $BANK$ = Ratio of deposit money bank domestic assets to deposit money bank domestic assets plus central bank domestic assets; $PRIV$ = Ratio of claims on the non-financial private sector to total credit; $PRIVY$ = Ratio of claims on the non-financial private sector to GDP; CAP = Market capitalization; $TURN$ = Stock turnover; VTR = value traded; $WINT$ = International capital market integration; VOL = Stock return volatility; $PRIVL$ = Ratio of loans to the private sector to GDP; $CCAPY$ = Ratio of domestic credit plus stock market capitalization to GDP; DIN = Disclosure index; $LEND$ = Bank lending to the private sector; LLY = Ratio of liquid liabilities to GDP; $M3Y$ = Ratio of M3 to GDP; $M3LY$ = Ratio of M3 less M1 to GDP; $CRED$ = Ratio of total credit to GDP; SPR = Spread of borrowing and lending interest rates; $VTRY$ = Ratio of stock traded to GDP; $TURNC$ = Ratio of stock value traded to market capitalization; $CAPY$ = Ratio of market capitalization to GDP; $VADD$ = Ratio of gross value added of financial sector to GDP; FOP = Ratio of stock of flows of foreign assets and liabilities to GDP; $\Delta CRED$ = Annual change of credit relative to lagged GDP; ΔY = Growth of real GDP; Δy = Real per capita GDP growth; ΔY_i = Real growth rate of value added at industry level; y = Real per capita GDP. Arrows represent causal directions. "↔" denotes bidirectional causality. "+/-" indicates the sign of causality. Column "**Results**": statistically significant estimates only.

Rousseau and Wachtel (2001) studies the trilateral relationships between finance-inflation-growth for a group of 84 countries over the post-1960 period. The results indicate that the effect of financial development on growth is positive and statistically significant only in periods where inflation is not very high. Shen and Lee (2006) shows that only stock market development variables (captured by stock traded and market capitalization) influence positively the output growth. Instead, variables measuring banking development show negative or zero effect on economic growth.[10]

Thanks to the availability of useful nineteenth century time series data for a large number of countries,[11] there has been an increasing interest in examining the finance-growth nexus by means of panel estimations. Examples in this direction are Apergis et al. (2007) and Chortareas et al. (2015). Apergis et al. (2007), via a panel cointegrating analysis and taking heterogeneity into account, show that a bi-directional causality between financial variables and growth exists. Chortareas et al. (2015), using panel cointegration methods allowing for cross-sectional dependence, find that the direction of causality runs from financial development to output for the advanced economies. In addition, they find that financial openness is the most important factor, for mature economies, in driving output in the long-run.

Differently from previous works, Rousseau and Wachtel (2011), Beck et al. (2014), Bezemer et al. (2016) show that the finance-growth relationship may change over time and among country groups. Rousseau and Wachtel (2011) find that the finance-growth relationship is robust in the period 1960-1989 and disappears over the subsequent 15 years. The authors conclude that the underlying relationship is unstable and may reappear in the next future. Beck et al. (2014) show that the strength of the finance-growth relationship exhibits important variations over time and among country groups. The cross-section estimation for various group of countries show that the size of financial sector (measured by the ratio of value added of financial sector to GDP) is strongly positive for high income countries, especially after the second half of 1990s. Bezemer et al. (2016) emphasize the fact that different financial measures may have a different effect on GDP growth. In particular, the authors find that

[10] The authors advance that the reason could be the presence of non-linear effects of bank development on growth.

[11] See, for instance, the rich database released by the World Bank (http://data.worldbank.org/country).

credit stocks supporting asset markets have a negative influence on growth, whereas credit flows (i.e., change of credit stocks relative to GDP) not supporting mortgage have a positive impact on growth.

2.2 Finance and growth: country-by-country analyses

If compared to cross-sectional or panel analyses, the number of country-specific studies focusing on the finance-growth relationship is rather limited. A list of main country-by-country empirical works is reported in Table 2. One of the first study investigating the financial development-growth relationship is Jung (1986). He selects 56 countries (where 19 are developed) and conducts (for each country) a Granger causality test between financial development measures and real per capita GDP growth for the postwar period. In particular, for the US, the results show a bi-directional causality between monetization variable (M2/GDP) and output growth. Arestis and Demetriades (1997), instead, study the causality in a VECM framework for Germany and US. The authors find evidence of real GDP contributing to banking and stock market development, but not the opposite.

Rousseau and Wachtel (1998) analyze the links between various indicators measuring the intensity of financial intermediation and economic output for the US and other four developed countries over the period 1870-1929. The authors perform a Toda and Phillips (1994) test on the VAR model and show that all financial intensity measures Granger cause output, while output does not Granger cause any of the employed financial indicators. Neusser and Kugler (1998) examine the finance-growth relationship by using financial sector GDP and manufacturing sector GDP as proxies for financial development and economic growth, respectively. In addition, the authors investigate the finance-productivity relationship. Specifically, Neusser and Kugler (1998) study such relationship for a group of 13 countries over the post-war era. For the US, they observed that finance causes (in Granger sense) manufacturing GDP and productivity, and not viceversa.

Arestis et al. (2001), using quarterly data for five developed economies spanning the period 1972-1998, find that financial development does not cause real GDP in the US. Instead, there seems to be a weak causality running from growth to financial development. These results for the US are in stark contrast to those observed for the other countries (i.e., France, Germany, and Japan). Shan et al. (2001), using quarterly data and splitting the sample in two sub-periods (1974-1998 and 1986-1998), find for the US a clear evidence of a bi-directional causality.

Table 2: THE FINANCE-GROWTH NEXUS: MAIN SINGLE-COUNTRY (TIME-SERIES) WORKS.

Authors	**Fin. Var.**	**Dep. Var.**	**Panel**	**Estimation technique**	**Results**
Jung (1986)	$MMBY$, $M2Y$	Δy	56 countries, 1948-1981 (US sample)	Granger causality test	For the US: $M2Y \overset{(?)}{\leftrightarrow} \Delta y$
Arestis and Demetriades (1997)	$CAPY$, $CRED$, σ_{sp}	y	2 countries, 1979q1-1991q4 (US sample)	Weak exogeneity test in VECM	For the US: $Y \overset{(+)}{\rightarrow} CAPY$; $Y \overset{(+)}{\rightarrow} CRED$
Rousseau and Wachtel (1998)	CBA, $CBSA$, FIA, MMB	Y	5 countries, 1870-1929 (US sample)	Toda-Phillips test	For the US: $CBA \overset{(+)}{\rightarrow} Y$; $CBSA \overset{(+)}{\rightarrow} Y$; $FIA \overset{(+)}{\rightarrow} Y$; $MMB \overset{(+)}{\rightarrow} Y$
Neusser and Kugler (1998)	Y_F	Y_M, TFP_M	13 countries, 1960-1993 (US sample)	Toda-Phillips test	For the US: $Y_M \overset{(+)}{\rightarrow} Y_F$; $Y_F \overset{(+)}{\rightarrow} TFP_M$
Arestis et al. (2001)	$CAPY$, $CRED$, σ_{sp}	Y	5 countries, 1972q2-1998q1 (US sample)	Weak exogeneity test in VECM.	For US: $Y \overset{(+)}{\rightarrow} CAPY$
Shan et al. (2001)	$PRIV$, STK	y	10 countries, 1974q1-1998q1 (US sample).	Toda-Yamamoto test	For the US: 1974-1998: $PRIV \overset{(?)}{\leftrightarrow} y$; 1986-1998: $PRIV \overset{(?)}{\leftrightarrow} y$, $STK \overset{(?)}{\leftrightarrow} y$
Amano (2005)	CBY, $PRIV$, $MULT$, $M2Y$	y	3 countries, 1874-1999 (US sample)	Toda-Phillips test	For the US: 1874-1920: $CBY \overset{(?)}{\rightarrow} y$, $PRIV \overset{(?)}{\leftrightarrow} y$; 1953-1999: $CBY \overset{(?)}{\rightarrow} y$, $y \overset{(?)}{\rightarrow} SPRIV$, $MULT \overset{(?)}{\leftrightarrow} y$, $M2Y \overset{(?)}{\rightarrow} y$
Amano (2013)	SMP, SCP	y	3 countries, 1888-1999 (US sample)	Weak exogeneity test in VECM	For the US: Prewar period: $SMP \overset{(+)}{\rightarrow} y$, $SCP \overset{(+)}{\rightarrow} y$; Postwar period: $SMP \overset{(+)}{\leftrightarrow} y$
Peia and Roszbach (2015)	$CAPY$, $PRIV$	y	22 countries, 1973q1-2011q1 (US sample)	Toda-Phillips, Weak exogeneity, Toda-Yamamoto and Granger causality test	For the US: $\Delta y \overset{(+)}{\rightarrow} PRIV$, $CAP \overset{(+)}{\rightarrow} \Delta y$

Notes: $MMBY$ = Ratio of currency to M1; $M2Y$ = Ratio of M2 to GDP; $CAPY$ = Ratio of stock market capitalization to GDP; $CRED$ = Ratio of total credit to GDP; σ_{sp} = Standard deviation of stock market prices; CBA = Assets of commercial banks; $CBSA$ = CBA plus saving institutions; FIA = $CBSA$ plus assets of insurance companies and credit cooperatives; MMB = Monetary base; Y_F = Real GDP of financial sector; $PRIV$ = Ratio of loans to the private sector to GDP; STK = Stock market index (by banks) to GDP; CBY = Ratio of deposits of commercial banks to GDP; $MULT$ = Ratio of M2 to monetary base; $M2Y$ = Ratio of M2 to GDP; SMP = Ratio of stock transactions plus M2 to population; SCP = Ratio of stock transactions plus commercial bank claims to population; Δy = Real per capita GDP growth; y = Real per capita GDP; Y = Real GDP; Y_M = Real GDP of manufacturing sector; TFP_M = Real TFP of manufacturing sector. Arrows represent causal directions. "$\leftrightarrow$" denotes bidirectional causality. "+/-" indicates the sign of causality, whereas "?" indicates that the sign is not reported. Column "**Results**": statistically significant estimates only.

Amano (2005) examines the causal relationships between financial development and per capita GDP growth for the US, UK, and Japan in two sub-periods: 1874-1920 and 1953-1999. For the US, the author shows that financial development tends to lead and causes output in both sub-samples. In an updated version, Amano (2013) examines the causal relationships for the same group of countries but splitting the full sample in pre-war and post-war years. Via weak exogeneity test in a VECM, the author shows that for the US both of the financial development measures considered (stock transaction plus M2 and stock transaction plus banks' claims on private sector, both normalized by population) tend to lead and cause output in the pre-war era, whereas a bi-directional causality among "stock transaction plus M2" measure and GDP appears in the post-war period.

Peia and Roszbach (2015) investigate the finance-growth nexus differentiating between stock market and banking sector development. The authors, using different statistical tests (i.e., Toda and Phillips (1994), Toda and Yamamoto (1995), and simply Granger causality) for each of the 22 countries considered, show that results change across countries. In particular, focusing on the US, their empirical evidence suggest that stock market development causes output growth, while the causality between banking sector development and growth goes in opposite direction.

2.3 On the theory of financial cycles[12]

The role played by the financial sector in driving business cycle fluctuations has been also investigated from a theoretical point of view. Seminal contributions on the impact on financial dynamics on the business cycle are Minsky (1974) Minsky (1978) and Kindleberger (1978). Needless to say, the 2008-2009 Great Recession has then contributed most to the increasing number of theoretical works aimed at embedding financial/credit cycles into RBC frameworks. This with the ultimate goal of examining the impact of financial and credit shocks on the real economic activity. Examples are Christiano et al. (2010); Kiyotaki and Moore (2012); Jermann and Quadrini (2012); Ajello (2016); Azariadis et al. (2016); Del Negro et al. (2017); Rouillard (2018).[13]

Christiano et al. (2010) introduce financial frictions and banking sector into a standard DSGE model with nominal rigidities. The authors find that a financial shock plays an important role in generating economic fluctuations in the Euro Area and US. More importantly, they find that over a the business cycle frequency, this shock explains more than 60 percent of volatility of investment in the US. Kiyotaki and Moore (2012) builds a DSGE model in

[12] We thank an anonymous referee for valuable comments that improved the contents of this section.

[13] In this respect, there are also several empirical studies showing that financial variables have a key role in driving macroeconomic fluctuations (see Al-Zoubi, 2008, 2017; Al-Zoubi et al., 2018).

which differences in the liquidity of distinct assets create a link between asset prices and macroeconomic aggregates. The authors find that liquidity shocks have a large negative impact on investment and output. Jermann and Quadrini (2012) develop a business cycle model with debt and equity financing to investigate the macroeconomic effects of financial shocks. They find that financial shocks (i.e., tightening firms' borrowing constraints) contribute significantly to the observed dynamics of real and financial variables.

Ajello (2016) estimates a model with financial frictions a lá Kiyotaki and Moore (2012) in which competitive financial intermediaries transfer resources between entrepreneurs with heterogeneous skills. His estimated model generates a boom in both the economy and the stock market following a positive financial shock. Azariadis et al. (2016) observe that unsecured credit shows great volatility and often leads output. To explain this stylized fact, the authors develop a dynamic general equilibrium model in which the unsecured component of the loan depends on a firm's expectations of the future availability of unsecured credit. The model is able to generate self-fulfilling credit cycles and shocks to expected future unsecured credit conditions have persistent effects on credit, productivity and output. In the spirit of Ajello (2016), Del Negro et al. (2017) extend the model of Kiyotaki and Moore (2012) and investigate the impact of a large financial shock of the order of magnitude observed in the 2008 recession. Their simulations suggest that the negative effect of this shock on output may be mitigated using liquidity facilities. Rouillard (2018) presents a standard two-country RBC model with non-separable preferences between consumption and leisure and domestic financial frictions. The authors show that positive financial shocks create important fluctuations in the labor wedge, pushing firms to demand more labor.

Overall, there seems to be also a large and growing number of theoretical studies showing strong linkages between financial cycles and macroeconomic fluctuations. This has led to an increasing consensus among economists that financial dynamics represent an important source of business cycle fluctuations. However, the aforementioned theoretical works tend to assume unrelated financial and productivity shocks. Actually, in this work we assume productivity to be directly driven by financial cycle, consistent with the empirical evidence that will be provided in Section 3.

3 The facts

In this section, we report some empirical regularities related to the US economic growth and the finance-growth relationship for the period 1870-2013. The implications drawn from these empirical facts will be then used - as assumptions - to develop our augmented stochastic neoclassical growth model.[14]

The first stylized fact on output growth is key to identify whether an exogenous growth or an endogenous growth framework fits best the long-run US economic growth path. The neoclassical (exogenous) growth model postulates stable equilibrium with a long-run constant output growth rate (Solow, 1956). Following changes in variables affected by government policy, the growth rate of the economy increases temporarily and then turns back to its original value.[15] In endogenous growth models, such policy changes should lead to permanent change in growth rates ("growth effect"). To this end, according to the exogenous growth theory, income per capita has to be characterized by a "linear trend" pattern with no shifts.

The second stylized fact helps to capture those financial factors causing (in Granger sense) GDP per capita growth. Differently from existing empirical works, which have focused on the impact of financial development on economic growth, here we analyze impacts on growth of two alternative financial variables, namely (*i*) the short-term nominal interest rate (i.e., proxy of monetary policy rate) and (*ii*) the world interest rate. Both measures can influence growth through effects not captured by a financial development indicator. In general, the short-term interest rate influences output via its impact on the allocation of resources and on the cost of borrowing (i.e., term structure).[16] Obviously, both effects have an impact on capital accumulation and productivity growth. Intuitively, the world interest rate has implication for the domestic output especially in the presence of highly integrated capital markets. Under financial integration, domestic agents are more sensitive to changes in international conditions and do not react only to domestic financial shocks.

Figure B.1 in Appendix B shows that financial integration rose significantly starting from

[14] Details on data are provided in Appendix A.

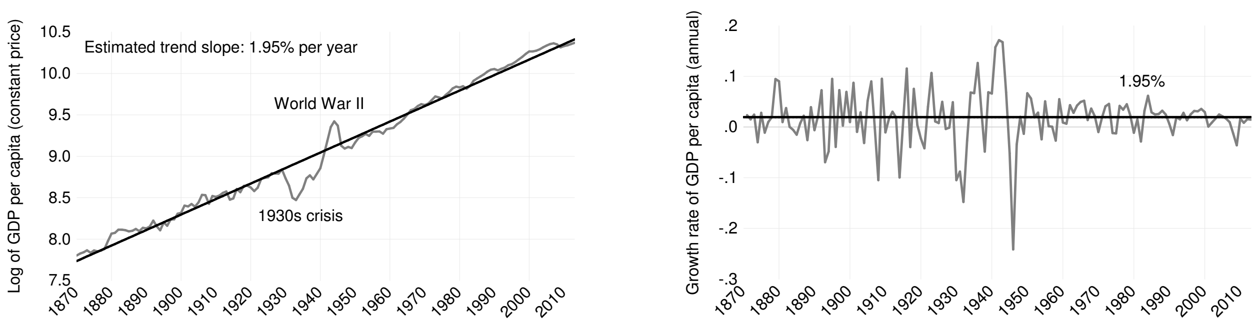
[15] The only change is the level of output which is permanently higher if the policy intervention produces a positive stimulus on the economy ("level effect").

[16] Note that the allocation of resources depends on the effect of changes in the short-term rate on inflation and relative prices. See, for example, (Sørensen and Whitta-Jacobsen, 2010, pages 574–584).

the '70s, reached its peak before the 2008-2009 subprime crisis, and (slowly) declined during the most recent EU sovereign debt crisis (Figure B.1, Panel B). The world interest rate – a weighted average of interest rates of major countries – tends to capture shocks occurring in international financial markets that might propagate in domestic – US in our case – economy.[17] Needless to mention, the dynamics depicted in Figure B.1 can have significant impacts on the US economy. More importantly, these impacts can be of different sign (or different magnitude) in different points in time.

Figure 1: Long-run stylized facts on the US economy.

Panel A　　　　　　　　Panel B



Notes: PANEL A: Trend of the log of the US real GDP per capita. PANEL B: US annual growth rate of the real GDP per capita. Horizontal black line in PANEL B indicates mean regimes. Regimes are identified via the Bai-Perron technique.

Stylized fact I. *The neoclassical (exogenous) growth hypothesis fits the* US *economic growth path.*

Figure 1 (Panel A) shows that the log of the US GDP per capita behaves as a trend-stationary variable. Except for the Great depression and World War II periods, the series fluctuates around a deterministic trend and exhibits a mean value of 1.95%. In the spirit of Russell (2011), Gallegati and Ramsey (2013), Clementi et al. (2015), we detect the presence of shifts by applying the Bai and Perron (BP) algorithm (see Bai and Perron, 1998, 2003) to the growth of real GDP per person (see Appendix C for details on the BP procedure). The BP algorithm selects zero breaks in the per capita real GDP growth series (see Figure 1 Panel B) with an estimated constant growth value of 1.95%. These two empirical regularities are

[17] By focusing on the UK, Mumtaz and Surico (2009) show a fall in the world (short-term) interest rate generates a one-step-ahead rise in domestic consumption and output.

in line with the recent findings of Papell and Prodan (2014) and Sobreira et al. (2014). A constant GDP per capita growth rate support the idea of employing and an exogenous growth framework to examine the effect of financial variables on long-term economic growth.

Stylized fact II. *Domestic and international financial factors influence the US growth.*

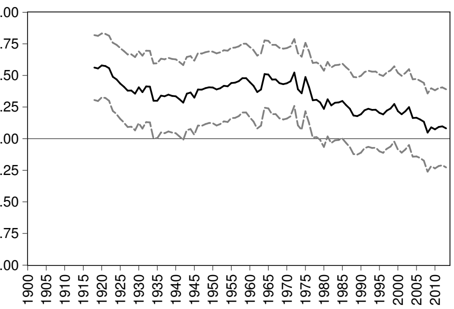
Domestic and international financial factors tend to co-move with output growth. Figure 2 (Panel A) reports the dynamic correlation between financial market development, fmd_t, (i.e., the sum of stock market transactions, smt_t, and money supply, $m2_t$, divided by population, L_t) and output growth. As expected, the correlation is positive and relatively high. More importantly, it is statistically significant until the end of the '70s. In Panel B we report the correlation between the lagged short-term interest rate and output growth. The correlation is negative and statistically significant during the period 1920-1940 and again after the mid-70s. Only starting from the mid-00s we observe a rapid increase in the correlation, which becomes positive (but statistically insignificant). Finally, Panel Panel C depicts the evolution of the correlation between the lagged world interest rate and output per capita growth. As expected, the correlation is negative ranging from -0.2 to -0.6. More importantly, it is highly statistically significant starting from the mid-70s.

Overall, lagged financial variables seem to lead the business cycle. We corroborate this finding by performing a Granger causality test between financial variables and output growth. Results are reported in Table 3. To capture potential changes in the causality, we run the test over two different periods: pre-war and post-war. Estimates show evidence of an unidirectional causality from i^{US} and i^{W} to output growth. Financial development is found to Granger causes output growth only in the pre-war era, whereas a bi-directional causality is instead observed in the post-war period.

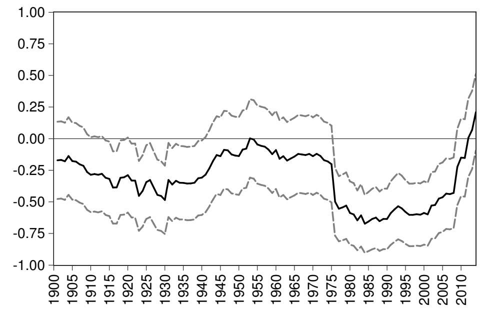
As a final robustness check, we perform an impulse response functions (IRFs) analysis to examine the response of output growth to shocks in financial variables in the post-war era. Both Cholesky-orthogonalized IRFs and Jordà (2005, 2009) local projections indicate that shocks to international financial factors (i.e., a world interest rate shock) propagate strongly to output growth and not viceversa (Figure F.1, Appendix F). More importantly, a world interest rate shock is found to explain a non-negligible fraction of output fluctuations. For

Figure 2: Dynamic Correlation: Output Growth vs. Financial Factors.

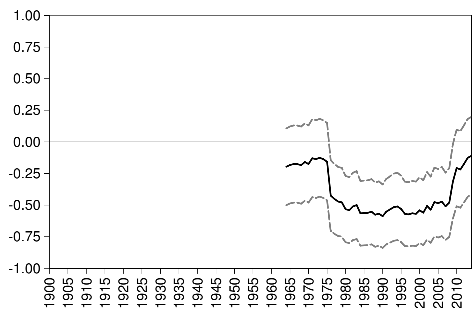
Panel A: $\rho\left(\Delta\log(y_t), \Delta\log(fmd_t)\right)$



Panel B: $\rho\left(\Delta\log(y_t), \Delta(i^{US}_{t-1})\right)$



Panel C: $\rho\left(\Delta\log(y_t), \Delta(i^{W}_{t-2})\right)$



Notes: PANEL A: Dynamic correlation between the growth of the US real GDP per capita ($\Delta\log(y_t)$) and the growth of the financial development indicator ($\Delta\log(fmd_t)$). PANEL B: Dynamic correlation between the growth of the US real GDP per capita ($\Delta\log(y_t)$) and the first difference of one-lagged US nominal short term interest rate ($\Delta(i^{US}_{t-1})$). PANEL C: Dynamic correlation between the growth of the US real GDP per capita ($\Delta\log(y_t)$) and the first difference of the two-lagged world short term interest rate ($\Delta(i^{W}_{t-2})$). The financial variable enters in correlation as simultaneous (lagged) if the maximum cross-correlation value corresponds to a contemporaneous (lead) of the variables to contemporaneous output growth. Correlations are computed using a rolling-window of 30 years. Dashed gray lines indicates 90% confidence intervals. Details on data sources and construction are reported in Appendix A.

Table 3: Granger causality: Output Growth vs. Financial Factors

	Δy vs Δfmd	Δy vs Δi^{US}	Δy vs Δi^{W}
Lag 1 (pre-war)	$\Delta y \xleftarrow[(0.04)^{**}]{+} \Delta fmd$	$\Delta y \xleftarrow[(0.06)^{**}]{-} \Delta i^{US}$	NA
Lag 1 (post-war)	–	$\Delta y \xleftarrow[(0.00)^{***}]{-} \Delta i^{US}$	$\Delta y \xleftarrow[(0.05)^{**}]{-} \Delta i^{W}$
Lag 2 (pre-war)	$\Delta y \xleftarrow[(0.08)^{*}]{+} \Delta fmd$	$\Delta y \xleftarrow[(0.09)^{*}]{-} \Delta i^{US}$	NA
Lag 2 (post-war)	$\Delta y \xleftarrow[(0.08)^{*}]{+} \Delta fmd$, $\Delta y \xrightarrow[(0.07)^{*}]{+} \Delta fmd$	$\Delta y \xleftarrow[(0.01)^{**}]{-} \Delta i^{US}$	$\Delta y \xleftarrow[(0.09)^{*}]{-} \Delta i^{W}$

Notes: Δy indicates the log difference of per capita real GDP; Δfmd denotes the log difference of financial market development index; Δi^{US} is the first difference of US short-term interest rate; Δi^{W} represents the first difference of world interest rate. Details on data sources and construction are reported in Appendix A. Arrows represent causal directions. Prewar corresponds to the period 1890-1949, whereas Postwar is the period 1950-2013. Data for i^{W} starts in 1933 and, for this reason, the correlation in Prewar sample cannot be calculated for lack of observations. "NA" indicates Not Available. P-values are reported below the arrows. ***, **, * denote significance at 1%, 5%, and 10% levels, respectively.

instance, at 5 year horizon, the contribution of a i^w shock to the volatility of output is as high as 32% (Table F.1).

Based on the empirical evidence produced in this section, one can draw the following implications:

In an exogenous growth framework (**fact I**): (i) *the per capita output growth, in the long-run, increases at the same rate of the (exogenous) technology; (ii) the technology follows a trend-stationary pattern. Since in an exogenous growth framework business fluctuations are generated by technological (temporary) changes,* **fact II** *suggests that an important source of these changes are due to both domestic and international financial factors.*

4 Theoretical framework

We assume, as in the standard Solow growth model, output Y_t to be produced using a two-factors Cobb-Douglas production function

$$Y_t = K_t^{\alpha} \left(A_t L_t\right)^{1-\alpha} \quad \text{with } 0 < \alpha < 1, \tag{1}$$

where K_t and L_t denote physical capital and labor, respectively, and A_t is the usual disembodied technology. The parameter α is the share of capital in a competitive Cobb-Douglas economy. Physical capital evolves as follows

$$K_t = I_{t-1} + (1-\delta) K_{t-1}, \tag{2}$$

where δ is the usual depreciation rate ($0 < \delta < 1$). In Eq. (2), investment, $I_t = sY_t$, and the savings rate, s, is constant. The evolution of capital per effective labor unit, $k_t = K_t/(A_t L_t)$, is then given by

$$\Delta \log\left(k_t\right) = -\Delta \log\left(A_t L_t\right) + \log\left(s k_{t-1}^{-(1-\alpha)} + (1-\delta)\right). \tag{3}$$

In line with the empirical facts reported in Section 3, the stochastic process determining technology is given by

$$\begin{aligned} &\log(A_t) = a_0 + gt + \sum_{k=1}^{N} \beta_k \log(\widetilde{x}_{k,t}) + u_{a,t} \\ &u_{a,t} = \phi_a u_{a,t-1} + \epsilon_{a,t} \quad |\phi_a| < 1,^{18} \end{aligned} \tag{4}$$

where $\log(\widetilde{x}_{k,t}) \doteq \log(x_{k,t}/\bar{x}_{k,t})$, $k = 1, \cdots, N$, denote N exogenous financial variables expressed as a log-difference from their long-term trends[19] and the numerical constants β_k, $k = 1, \cdots, N$, indicate the magnitude of the influence of the exogenous variables on technology. Notice that our setting allows for the inclusion of exogenous financial variables with a different lag structure, e.g. $\log(\widetilde{x}_{k,\,t-1})$, accounting for the fact that some variables may not produce immediate effects on technology. As in Lee et al. (1997), the technology shock, $u_{a,t}$, captures all those factors that might produce TFP changes. We stress that our technology specification differs from the one proposed by Lee et al. (1997) and Binder and Pesaran (1999). In their setting growth can be affected by either the technology growth rate, g, or the exogenous shock, $u_{a,t}$ and not by any other exogenous variable. Differently, we aim to identify additional variables that could have strong influences on the dynamics of the technology innovation (i.e., growth). By relying on this technological progress, in Appendix D we show that the logarithm of the real GDP per capita takes the following form:

$$\begin{cases} \log(y_t) = \mu + \lambda \log(y_{t-1}) + g(1-\lambda)t + (1-\alpha)\sum_{k=1}^{N} \beta_k \log(\tilde{x}_{k,t}) + (\alpha - \lambda)\sum_{k=1}^{N} \beta_k \log(\tilde{x}_{k,t-1}) + e_t \\ u_{a,t} = \phi_a u_{a,t-1} + \epsilon_{a,t}, \end{cases} \tag{5}$$

where μ and e_t are defined as follows:

$$\mu = -\alpha h + \lambda g + (1-\lambda)\left[a_0 - \frac{\alpha}{1-\alpha}\left[\log(n+g+\delta-h) - \log(s)\right]\right] \tag{6}$$

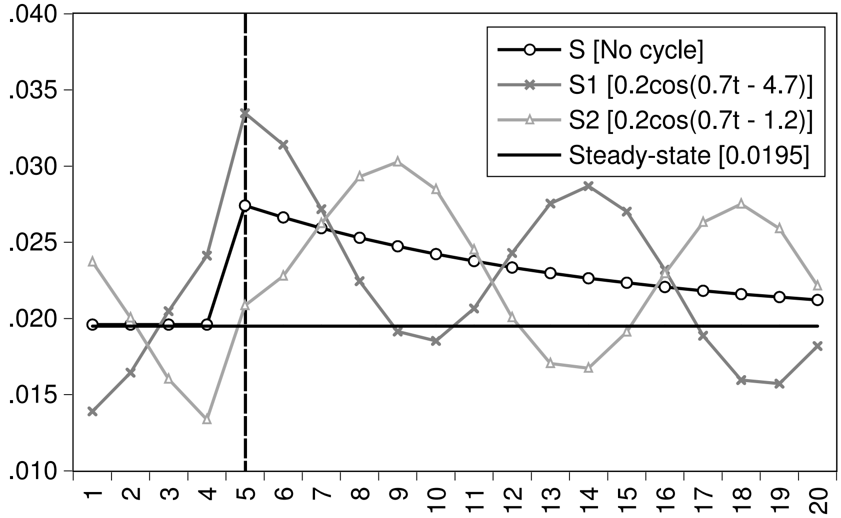
[18] This value is consistent with empirical evidence and, of course, ensures the existence of a steady-state.
[19] In terms of our notation, for any given variable x, $\bar{x}$ denotes the variable's trend or long-run value.

$$e_t = [(1-\lambda) - (1-\alpha)(1-\phi_a)]\, u_{a,t-1} + (1-\alpha)\epsilon_{a,t}. \tag{7}$$

In particular, notice that the term e_t depends on $u_{a,t}$ and the technology shock is modeled as in Eq. (4). Notice that the model is derived by assuming constant population and saving rate. But what if s or n shifts once in the steady state? Let us refer to Appendix E, where the law of motion around the steady state of our stochastic economy framework is derived. Eq. (E.7) shows that when there is a shift in the saving rate and/or population growth rate, the per capita output growth exhibits a transitory growth different from that one of equilibrium. Anyway, this difference disappears over the long-run. So, the dynamics described by Eq. (5) holds at some points in time in the future.

Eq. (E.7) is interesting for another reason. It suggests that the cyclical patterns should not be excluded from our analysis. As an example, Figure 3 shows that cyclical factors may exacerbate the effects of a permanent change in s_t – in a positive or negative sense depending on the phase of the cycle – on economic growth. In other words, the presence of the cyclical term implies that the effects of policy interventions are state-dependent. Indeed, starting at time $t = 0$ from a steady-state growth rate equal to 0.0195, we suppose that in period $t = 5$ the s_t permanently increases from 0.2 to 0.25. In the long-run we have that $\mathbb{E}(\Delta \log(y_\infty))$ $= 0.0195$, but in the medium run the stimulus effect on output strongly depends on the underlying cyclical phase of known factors affecting the growth rate. In good times (Figure 3, S1), the medium run effects on growth are very strong. The opposite is true in bad times (Figure 3, S2). It is thus relevant for a policy-maker while implementing a new policy to be properly informed about the phase of the economic cycle. The dynamics described in Figure 3 is also of general interest. The policy-maker may stimulate the growth rate in the medium run also influencing the cyclical phase of some factors entering in the growth dynamics. For example by favoring the development of financial markets or by deviating the monetary policy rate from its normal trend. As we are going to investigate empirically in the next two sections, this last aspect is very complex because the cyclical factors influencing the growth rate vary over time creating model's instability.

Figure 3: Effect of a change in the saving rate in the per capita output growth.



Notes: The reported simulation exercise is based on Eq. (E.7) and assumes that no exogenous shocks hit the economy (i.e., $\Delta u_{a,t} = 0$). Results are reported for two cycles exhibiting the same amplitude but different phases.

5 Empirical strategy

Motivated by stylized fact II, we select three exogenous financial factors ($\log(\tilde{x}_t)$): (i) the US nominal short-term interest rate $\left(i_t^{\mathrm{US}}\right)$[20]; (ii) the financial market development index (fmd_t); (iii) (a proxy) for the nominal short-term world interest rate $\left(i_t^{\mathrm{W}}\right)$. We thus study the effects of monetary policies and developments in financial markets on the cyclical pattern of technology and economic growth[21]. To do so, we estimate different versions of the system defined in Eq. (5). These model versions are summarized in Table 4.

The system represented in Eq. (5) can be expressed in a state-space form, where the first equation represents the signal equation, and the second one is the state equation. The state space model is estimated by maximum likelihood via the Kalman recursion. For identification

[20]Within the Solow's framework, the steady state of the real interest rate r is a function of the depreciation rate of capital, productivity and population growth, and the saving rate. From Fisher relation, we know that $\mathrm{r}_t = \mathrm{i}_t - \pi_t^e$, with π_t^e the expected inflation rate. In steady state, we expect that a relation like this holds: $\mathrm{r}_{ss} = \mathrm{i}_t^* - (\pi_t^e)^*$, where i_t^* and $(\pi_t^e)^*$ represent the long-run trend of the nominal interest rate and of the expected inflation rate, respectively. If this is true, we can then express the log-deviations from equilibrium as $\log(\mathrm{r}_t) - \log(\mathrm{r}_{ss}) = (\log(\mathrm{i}_t) - \log(\mathrm{i}_t^*)) - (\log(\pi_t^e) - \log(\pi_t^e)^*))$. A short-term interest rate above its trend value implies a disincentive to invest in new machinery and equipment with a negative impact on technology improvements.

[21]Data sources and details on the computation of the cyclical components are reported in Appendix A.1 and A.2, respectively.

Table 4: Model Specification

Model Specification	**Exogenous variables**		
	i^{US}	fmd	i^{W}
Model I	YES, 1-2		
Model II		YES, 0-1	
Model III			YES, 1-2
Model IV	YES, 1-2	YES, 0-1	
Model V		YES, 0-1	YES, 1-2
Model VI			

Notes: Numbers on the side of "YES" identify models' lags.

purposes, some restrictions are needed. We therefore fix some parameters to values reported in the existing literature. Precisely, the annual depreciation rate δ is set equal to 12.5% as suggested by Epstein and Denny (1980), Kollintzas and Choi (1985), and Bischoff and Kokkelenberg (1987). As pointed out by Lee et al. (1997) and Binder and Pesaran (1999), the parameter h is supposed to be positive but small, in order to defined output as in Eq. (D.14). Therefore, we impose $h = 0.03$. Based on the analysis conducted in Appendix C, we assume the population growth n and the saving rate s to be stationary processes shifting around their long-run average. Finally, to set a proper value for α, (i.e., the capital share in a competitive Cobb-Douglas economy), we follow existing studies. In particular, Gollin (2002) indicates a value of $\alpha = 0.33$ for most developed countries, whereas Piketty and Zucman (2014) suggest a value of $\alpha = 0.26$ for the US over the long-run.[22]

Using this information, two strategies are implemented in order to identify the core parameters λ, g, α, and β_i. The first one consists in fixing α equal to the value suggested by the literature and subsequently estimating λ. In the second strategy we instead let α to be estimated first.

The λ is then implicitly retrieved according to Eq. (D.9). Thus models in Table 4 are estimated using the two aforementioned strategies and different combinations of values for α, s, and n: (i) $\alpha = 0.26$ with constant n and s (version I); (ii) $\alpha = 0.33$ with constant n and s (version II); $\alpha = 0.26$ with mean-shift pattern of n and s (version III); $\alpha = 0.33$ with mean-shift pattern of n and s (version IV); α free to be estimated with mean-shift pattern of n and s (version V). Given that the period investigated in this work is rather

[22] See also the TableUS11 in the spreadsheet of US data available at http://gabriel-zucman.eu/capitalisback/.

long, a rolling estimation is then executed to examine whether the model is robust over time (i.e., model uncertainty). We decide to consider a window of 30 observations so as to obtain reliable estimates and give novel insights on the evolution of the parameters. In addition, a dynamic version of AIC model selection (in the finite sample correction) using Akaike weights is implemented in order to select (in each window) the "best" model.

6 Empirical results

Tables 5 and 6 report the estimation results of the state-space models defined in Table (4). Overall, estimates are in line with the stylized facts presented in Section 3. In particular, we find: $\beta_{fmd} > 0$, $\beta_{i^{US}}$ and $\beta_{i^{W}} < 0$.[23] Let us stress that our evidence are robust to (i) different values of α and (ii) different specifications of s and n (i.e., using s and n as constant or as mean-shift parameters yield similar results). For most of the models diagnostic tests are satisfactory. There is no trace of autocorrelation in the residuals, except for version V of MODEL VI. Normality test of residuals is rejected – due to excess of kurtosis – for MODEL I (version IV), MODEL III (all versions), MODEL IV (versions II, IV, and V), MODEL V (all versions), and MODEL VI (version IV). The skewness test of normality is instead accepted at 10% confidence level (results on this test are available upon request). Since excess kurtosis is not a big concern compared to skewness in likelihood-based analysis (see Juselius (2006), page 26), we prefer to be parsimonious in using dummies for correcting large outliers.[24]

The AIC criteria is then used to compare the various non-nested models. Standard empirical practices indicate to accept a model on the basis of the "raw" AIC statistics only, where AIC is defined as AIC=2K - 2log($\mathcal{L}(\bar{\theta}|y)$). The best model is then represented by the specification exhibiting the lower AIC score. Since the term $-2\log(\mathcal{L}(\bar{\theta}|y))$ is affected by large T, we divide for T the AIC obtaining a statistic (the one reported in Table 5 and 6, denoted AIC$_c$) adjusted for sample size allowing us to fairly compare a large variety of models. The AIC$_c$ related to all the specification of MODEL V is lower that the one of all

[23]It is important to note here that the estimation results are immune to the endogenous bias. This is due to the fact that i^{US} and i^{W} enter as lagged variables and, besides, in the signal equation fmd Granger causes the output growth (see Section 3).

[24]It is also important to note that problem of non-normality (caused by large outliers) in US output time series over long-horizons is recognized by many authors as discussed in Franke (2014).

the other models. Loosely speaking, we can assume MODEL V to be the best model.

As previously mentioned, since we use more than 100 years of data, a dynamic analysis is worth to be carried on. To this end, we estimate our models using a rolling window of 30 years. To examine model's performance the Akaike weights methodology is employed.[25] Figure 4 depicts the $\log_{10}(\mathrm{ER}_i)$ of models in Table 4 calculated in each window.[26] Dynamic estimates suggest that model's performances change over time. Specifically, we observe that starting from the mid-50s the different specifications of our novel framework performs better than the classical version without cyclical components (i.e., MODEL VI). This suggests that financial cyclical components play an important role in explaining growth dynamics. Notice that the performance of our augmented model varies across time and specifications generating instability in the TFP dynamics. Put it differently, the model's performance depends on the choice of the factor influencing the TFP (and, consequently, growth). Moreover, it is related to the cyclical phase of the economy. Loosely speaking, there cannot be a unique static and true model shaping TFP process. These results are clear from Figure 5, which depicts the dynamics of cyclical component parameters of MODELS I-IV. The results are in line with the dynamic correlation analysis conducted in Section 3 and depicted in Figure 2. We observe that the domestic interest rates (i^{US}) has played a significant role during the '20s and '30s and after the '70s. This is true for MODEL I and MODEL IV. The financial development index (fmd) is positive and statistically significant until the last 30-35 years (see PANEL B, D, and F). The world interest rate i^W has played a more important role than other factors in driving the short-run output growth over the last 30-35 years, as indicated in PANELS C and G. Domestic factors have played an important role especially until the second half of the '70s. From the early '80s, international factors – proxied by the dynamics of world interest rate – have played a dominant role.

Let us remark that estimates in Figure 5 reflect the dynamic correlation analysis conducted in Section 3 and, more importantly, seem to be in line with existing empirical findings. Concerning the domestic interest interest rate effects, the influence of short-term interest rate fluctuations on output dynamics in the pre-war period fits with the "monetarist view" ar-

[25]For further details on this model selection approach see the Appendix G.

[26]For brevity's sake, we plot only version I. For all the other specifications similar conclusions are drawn. Results are available upon request from the authors.

Table 5: Summary output of the estimated parameters: Model I, Model II and Model III

	Model I (1890-2013)					**Model II (1934-2013)**					**Model III (1890-2013)**				
Parameters	I-I	I-II	I-III	I-IV	I-V	II-I	II-II	II-III	II-IV	II-V	III-I	III-II	III-III	III-IV	III-V
α	0.260	0.330	0.260	0.330	0.213	0.260	0.330	0.260	0.330	0.250	0.260	0.330	0.260	0.330	0.375
					(0.055)					(0.051)					(0.067)
λ	0.853***	0.852***	0.848***	0.852***	0.901	0.834***	0.834***	0.830***	0.837***	0.906	0.868***	0.867***	0.853***	0.855***	0.920
	(0.045)	(0.044)	(0.049)	(0.048)		(0.041)	(0.040)	(0.046)	(0.046)		(0.069)	(0.068)	(0.080)	(0.082)	
g	0.019***	0.019***	0.019***	0.018***	0.019***	0.019***	0.019***	0.019***	0.019***	0.019***	0.020***	0.020***	0.020***	0.020***	0.020***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
$\beta_{i^{\mathrm{US}}}$	-0.039	-0.044	-0.037	-0.041	-0.038										
	(0.029)	(0.032)	(0.027)	(0.029)	(0.025)										
β_{fmd}						0.192***	0.210***	0.181***	0.193***	0.181***					
						(0.040)	(0.043)	(0.039)	(0.042)	(0.038)					
$\beta_{i^{\mathrm{W}}}$											-0.081	-0.093	-0.082	-0.093	-0.098
											(0.056)	(0.062)	(0.055)	(0.059)	(0.060)
ϕ_a	0.730***	0.706***	0.792***	0.800***	0.799***	0.643***	0.613***	0.741***	0.756***	0.743***	0.750***	0.731***	0.733***	0.741***	0.786***
	(0.108)	(0.110)	(0.080)	(0.073)	(0.102)	(0.113)	(0.111)	(0.084)	(0.076)	(0.111)	(0.124)	(0.131)	(0.182)	(0.185)	(0.174)
Model selection															
AIC$_c$ (normalized)	-3.598	-3.598	-3.579	-3.566	-3.585	-3.763	-3.764	-3.729	-3.706	-3.714	-4.013	-4.014	-4.006	-3.986	-3.966
Goodness of fit															
$\bar{\mathrm{R}}^2$	0.997	0.997	0.997	0.997	0.997	0.997	0.997	0.997	0.997	0.997	0.996	0.996	0.996	0.996	0.996
Diagnostic test (p-values)															
LB	0.098	0.098	0.065	0.049	0.050	0.109	0.091	0.104	0.076	0.090	0.000	0.000	0.000	0.000	0.000
Q(2)	0.057	0.057	0.100	0.120	0.035	0.104	0.103	0.163	0.195	0.067	0.815	0.814	0.852	0.914	0.879
Q(4)	0.145	0.146	0.224	0.258	0.098	0.188	0.188	0.260	0.299	0.123	0.320	0.314	0.337	0.388	0.472

Notes: Model I and Model II are estimated over the period 1934-2013 due to the constraint on the availability of the data. Model II is estimated over the period 1890-2013. In versions I and II of each model specification, the annual saving rate s and the annual growth rate of the population g are kept constant and equal to their average over the entire period. In versions III, IV, V, instead, they are set equal to the mean-shift values reported in Table C.1, Appendix C. Moreover, in version V the parameter α is estimated and the parameter λ calculated according to Formula (D.9). The parameter h is set to 0.03 in all models. Two intervention variables are inserted into the signal equation: The first one for the year of the recession in 1930s and the second one for the World War II. The Huber-White type standard errors are reported between brackets. The starting values for the parameters is set to: $\alpha = 0.33$, $\lambda = 0.7$, $g = 0.019$, $\beta^{\mathrm{US}} = -0.05$, $\beta^{\mathrm{W}} = -0.08$, $\beta^{\mathrm{FMD}} = 0.15$, $\phi_a = 0.8$. $\bar{R}^2$ indicates the adjusted R square. AIC$_c$ indicates the AIC with the finite sample correction. N indicates the Jarque-Bera Normality test. Q indicates the Box-Liung statistic of the autocorrelation of order 2 and 4 of the residuals. The Q statistic is corrected for taking into account the number of estimated hyper-parameters (i.e., the disturbance variance). ***, **, and * denote significance at 1%, 5% and 10% levels, respectively.

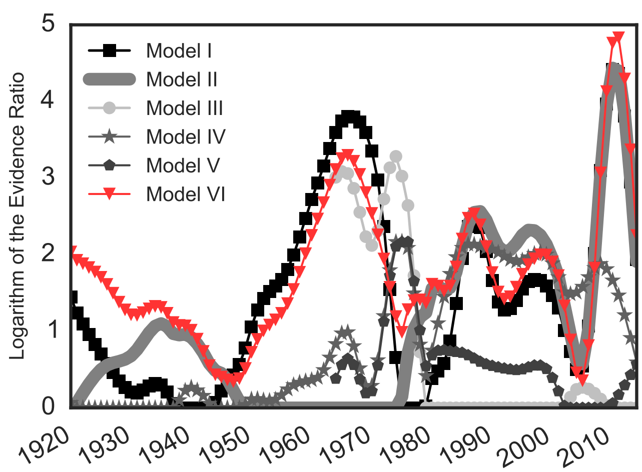
Table 6: Summary output of the estimated parameters: Model IV, Model V and Model VI

	Model IV (1890-2013)					Model V (1934-2013)					Model VI (1890-2013)				
Parameters	I-I	I-II	I-III	I-IV	I-V	II-I	II-II	II-III	II-IV	II-V	III-I	III-II	III-III	III-IV	III-V
α	0.260	0.330	0.260	0.330	0.257	0.260	0.330	0.260	0.330	0.380	0.260	0.330	0.260	0.330	0.207
					(0.052)					(0.061)					(0.055)
λ	0.836***	0.836***	0.832***	0.839***	0.906	0.852***	0.852***	0.829***	0.834***	0.921	0.851***	0.850***	0.846***	0.850***	0.900
	(0.042)	(0.040)	(0.047)	(0.046)		(0.065)	(0.065)	(0.076)	(0.082)		(0.045)	(0.044)	(0.048)	(0.047)	
g	0.019	0.019	0.019	0.019	0.019	0.020	0.020	0.021	0.021	0.021	0.019	0.019	0.019	0.018	0.019
	(0.001)	(0.001)	(0.001)	(0.001)		(0.001)	(0.001)	(0.001)	(0.001)		(0.001)	(0.001)	(0.001)	(0.001)	
$\beta_{i^{\mathrm{US}}}$	-0.044	-0.049	-0.043	-0.048	-0.047										
	(0.031)	(0.035)	(0.029)	(0.032)	(0.029)										
β_{fmd}	0.191***	0.210***	0.182***	0.195***	0.183***	0.142***	0.154***	0.151***	0.161***	0.173***					
	(0.038)	(0.042)	(0.037)	(0.041)	(0.037)	(0.044)	(0.048)	(0.043)	(0.047)	(0.049)					
$\beta_{i^{\mathrm{W}}}$						-0.124***	-0.140***	-0.130***	-0.145***	-0.151***					
						(0.052)	(0.058)	(0.053)	(0.057)	(0.058)					
ϕ_a	0.663***	0.633***	0.756***	0.770***	0.762***	0.708***	0.689***	0.651***	0.663***	0.717***	0.723***	0.698***	0.787***	0.795***	0.793***
	(0.115)	(0.112)	(0.084)	(0.074)	(0.108)	(0.097)	(0.103)	(0.148)	(0.161)	(0.157)	(0.110)	(0.112)	(0.081)	(0.074)	(0.105)
Model selection															
AIC_c (normalized)	-3.764	-3.764	-3.731	-3.709	-3.716	-4.130	-4.127	-4.137	-4.106	-4.061	-3.601	-3.602	-3.583	-3.570	-3.588
Goodness of fit															
$\bar{\mathrm{R}}^2$	0.997	0.997	0.997	0.997	0.997	0.996	0.996	0.996	0.996	0.996	0.996	0.996	0.996	0.996	0.996
Diagnostic test (p-values)															
LB	0.058	0.047	0.061	0.042	0.028	0.000	0.000	0.001	0.003	0.000	0.091	0.092	0.059	0.046	0.066
Q(2)	0.083	0.083	0.139	0.172	0.056	0.738	0.743	0.727	0.834	0.798	0.066	0.066	0.109	0.128	0.039
Q(4)	0.141	0.140	0.203	0.239	0.092	0.329	0.319	0.282	0.325	0.445	0.167	0.168	0.247	0.280	0.111

Notes: Model IV and Model VI are estimated over the period 1934-2013 due to the constraint on the availability of the data. Model V is estimated over the period 1890-2013. In versions I and II of each model specification, the annual saving rate s and the annual growth rate of the population g are kept constant and equal to their average over the entire period. In versions III, IV, V, instead, they are set equal to the mean-shift values reported in Table C, Appendix C.1. Moreover, in version V the parameter α is estimated and the parameter λ calculated according to Formula (D.9). The parameter h is set to 0.03 in all models. Two intervention variables are inserted into the signal equation: The first one for the year of the recession in 1930s and the second one for the World War II. The Huber-White type standard errors are reported between brackets. The starting values for the parameters is set to: $\alpha = 0.33$, $\lambda = 0.7$, $g = 0.019$, $\beta^{\mathrm{US}} = -0.05$, $\beta^{\mathrm{W}} = -0.08$, $\beta^{\mathrm{FMD}} = 0.15$, $\phi_a = 0.8$. $\bar{R}^2$ indicates the adjusted R square. AIC_c indicates the AIC with the finite sample correction. N indicates the Jarque-Bera Normality test. Q indicates the Box-Liung statistic of the autocorrelation of order 2 and 4 of the residuals. The Q statistic is corrected for taking into account the number of estimated hyper-parameters (i.e., the disturbance variance). ***, **, and * denote significance at 1%, 5% and 10% levels, respectively.

guing that changes in money supply lead to output variations (see, for example, Friedman and Schwartz, 1963; Christiano et al., 2003).[27] Differently, the effectiveness of interest rate fluctuations over the 1975-2005 period seems to be in line with Tatom (1984) (for post-oil shock era) and Stock and Watson (2002) (for the great moderation era). The effectiveness of world interest rate in influencing the business fluctuations is then consistent with Volosovych (2011, 2013) who shows that financial integration starts rising in the mid-'70s.[28] Finally, the declining relevance of the financial development index in capturing output growth corroborates the aforementioned discussion on the increasing role played by international factors. In other words, developed economies are today more exposed to international/global shocks and only marginally driven by domestic financial factors.

Figure 4: Dynamic Model Selection



Notes: Dynamic calculation of $\log_{10}(\mathrm{ER}_i)$ for MODELS I-V, specification I. Best performance := $\log_{10}(\mathrm{ER}) = 0$. Additional details are provided in Appendix G.

7 Concluding remarks

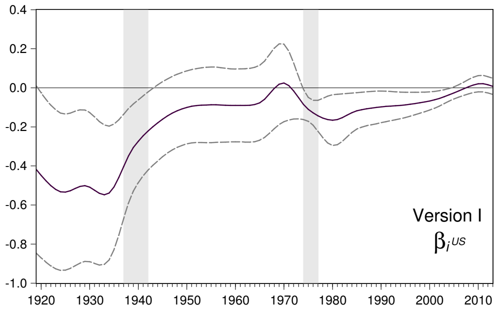
This paper investigates the role of domestic and international financial factors in driving US macroeconomic fluctuations over the period 1890-2013. Differently from existing studies

[27] At the time, the FED had money supply as target of monetary policy. Given the well-known relationships between money supply and interest rate (see Tatom, 1984), the reasoning is compatible with our empirical finding.

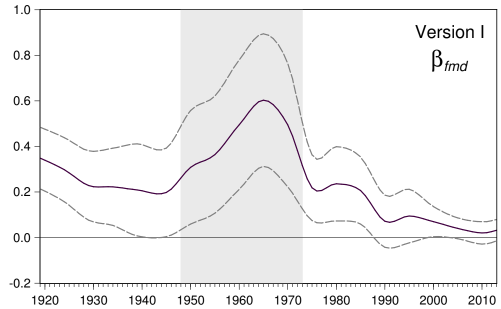
[28] The observed deterioration of this effect in the aftermath of Great Recession (Panel C and G of Figure 5) is in line with Billio et al. (2017)'s findings.

Figure 5: Dynamic Estimates of Cyclical Component Parameters of MODELS I-V

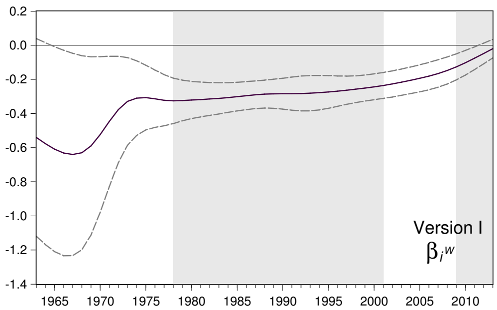
Panel A: MODEL I [β_{iUS}]



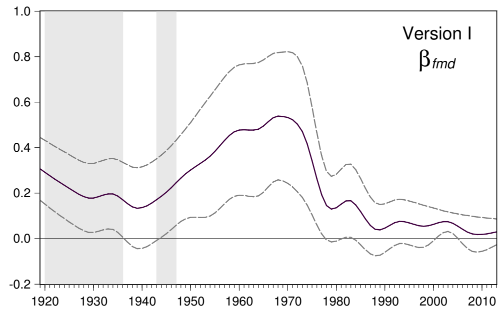
Panel B: MODEL II [β_{fmd}]



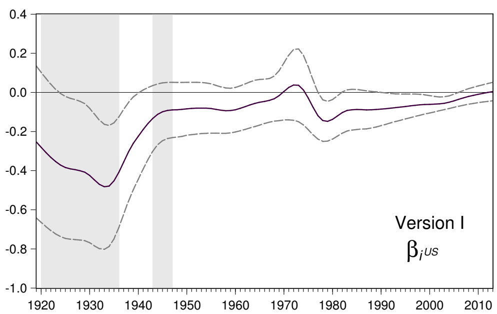
Panel C: MODEL III [β_{iW}]



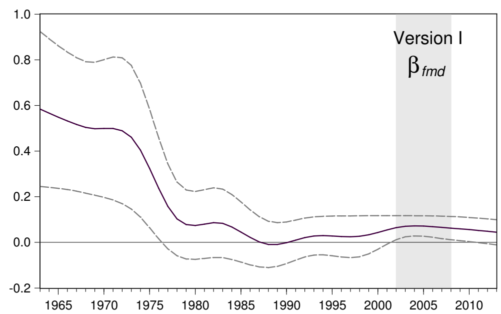
Panel D: MODEL IV [β_{fmd}]



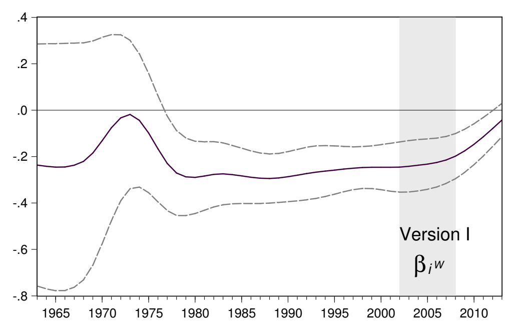
Panel E: Eq. MODEL IV [β_{iUS}]



Panel F: MODEL V [β_{fmd}]



Panel G: MODEL V [β_{iW}]



Notes: Dynamic estimates of cyclical components MODELS I-VI. The dotted lines indicate the confidence bands at 5% of significance level. The grey shaded areas indicate the period in which the relative equation is the best according to the Akaike weights criteria depicted in Figure 4.

on the finance-growth nexus, our empirical strategy is based on a newly developed exogenous growth model where financial factors enter as cyclical components in the TFP growth process. The resulting empirical model, which collapses to a standard state-space representation, is flexible enough to (dynamically) investigate the role played by a large variety of financial/banking/credit factors in shaping growth dynamics.

We first show key stylized facts on the US growth rate supporting our augmented stochastic neoclassical growth framework. By means of simple dynamics estimations, we then find that the role of each financial cyclical factor in explaining growth is not constant over time creating instability on the model's specification. On the one hand, domestic financial factors are found to significantly influence macroeconomic fluctuations only in the first half of the 20th century. On the other hand, as financial integration rises the contribution of global factors to explaining growth increases.

Taken together, the empirical results presented in this paper have implications for the effectiveness of policy interventions aimed at stimulating growth. In particular, our results suggest that policymakers should not exclude from their analysis the cyclical phase of exogenous financial factors once a new policy measure is introduced.

We conclude by arguing that, despite its analytically complexity, our novel framework allows for a user-friendly empirical estimation procedure. We thus believe that our setting can be used to investigate economic dynamics in other countries. In this respect, further research should aim at detecting the presence of cross-country common cyclical components.

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A Data appendix

A.1 Data sources

Table A.1 reports the source of data as well as notations used in our research.

Table A.1: Sources of the data

Real GDP per capita (y_t)
From 1870 to 2008: Bolt and Zanden (2014). After 2008: The World Bank website.
Population (L_t)
From 1870 to 2005: Bolt and Zanden (2014). After 2005: The World Bank website.
Saving rate (s_t)
From 1870 to 1988: Maddison (1992). After 1988: The World Bank website.
US nominal short-term interest rate (i_t^{US})
From 1870 to 2013: Measuring Worth Project website (http://www.measuringworth.com).
G7 nominal short-term world interest rate (i_t^{W})
From 1933 to 1970: Homer and Sylla (1996). After 1970: OECD website.
GDP deflator p_t
From 1888 to 1993: Amano (2013). After 1993: The World Bank website.
US amount of annual stock transactions in dollar (stm_t)
From 1988 to 1993: Amano (2013). After 1993: The World Bank website.
Money supply ($m2_t$)
From 1988 to 1993: Amano (2013). After 1993: The International Monetary Fund (IMF).

Notes: Saving rate is measured as the gross private saving divided by GDP. G7 stands for Canada (CAN), France (FR), Germany (DE), Italy (IT), Japan (JPN), United Kingdom (UK), United States (US).

A.2 Data construction

In this section we describe the procedure employed to construct the time series used in the empirical analysis. Following Gagnon and Unferth (1995) and Volosovych (2011), the world interest rate, i_t^{W}, is captured by the first Principal Component extracted from the G7 interest rates data matrix. The Financial Market Development (fmd_t) indicator is defined as

$$fmd_t \doteq \frac{stm_t + m2_t}{p_t \mathrm{L}_t}, \tag{A.1}$$

where stm_t is the amount of annual stock transactions, $m2_t$ is the money supply, p_t is the GDP deflator and L_t is the population.

In order to obtain the (log) differences between i_t^{US} and i_t^{W} with respect to their long-run trend, we use an univariate trend-cycle decomposition. Precisely, each of the observed time series is decomposed into trend (ϑ_t) and cycle (ψ_t) components. For instance, in the i_t^{US} case, we have

$$\begin{aligned} i_t^{\mathrm{US}} &= \vartheta_t + \psi_t + \epsilon_t, \\ \vartheta_{t+1} &= \vartheta_t + \xi_t, \end{aligned}$$

where ϵ_t and ξ_t are independent Gaussian shocks. The cycle component has a trigonometric form with frequency associated with the length of the (business) cycle.

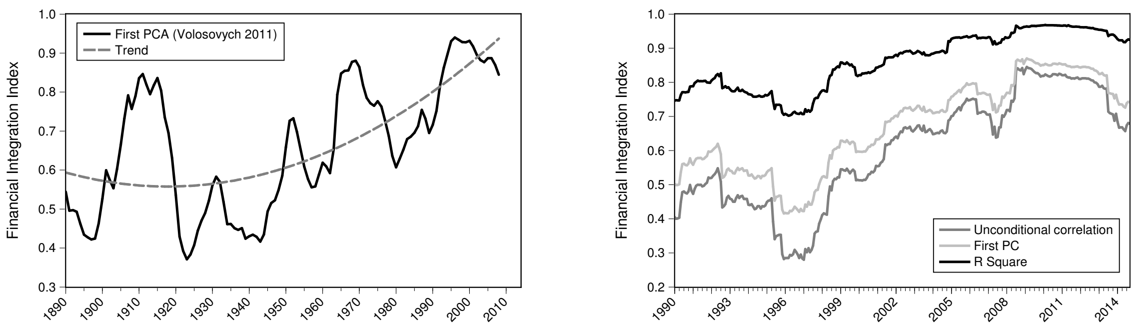
The (log) difference between fmd_t with respect to its long-run trend is identified as the residual of a regression between $\log(fmd_t)$ and a cubic trend.

Finally, the (log) difference between the real GDP per capita with respect to its long-run trend is identified as the residual of a regression between $\log(\mathrm{y}_t)$ and a linear trend (see Section 3). The population growth, i.e., $n_t = \Delta\log(L_t)$, and the saving rate s_t series are considered as mean-shift variables according to analysis conducted in Appendix C. The Bai and Perron (see Bai and Perron (1998, 2003)) algorithm identifies the dates of breaks in the population growth and in the saving rate series so as to minimize the sum of the squared residuals and thereby identify the number of regimes.

B Additional figures

Figure B.1: International financial integration dynamics

Panel A | Panel B



Notes: PANEL A reports the financial integration pattern over the long-run as computed by Volosovych (2011). Panel B depicts the evolution of the equity market integration process over the last 20 years. Integration dynamics in PANEL B are computed using three different measures: (i) unconditional correlation, (ii) First Principal Component (1st PC), and (iii) cross-country average adjusted R-square. The 1st PC is computed as in Volosovych (2011, 2013). The R-square is computed following Pukthuanthong and Roll (2009).

C Bai Perron test for shifting means

Let g_t be either the growth rate of a variable Z_t or the ratio between the generic variable Z_t and GDP (i.e., growth rate:= $g_t = \log(Z_t) - \log(Z_{t-1})$; ratio:= $g_t = \frac{Z_t}{GDP_t}$) To investigate shifts in the mean of a time series we can start from the following structure:

$$g_t = \gamma_t + \epsilon_t, \quad t = 1, ..., T \tag{C.1}$$

where γ_t is a time-varying intercept and $\epsilon_t \sim iid(0, \sigma^2)$. A standard approach for modeling γ_t – developed by Bai and Perron (1998, 2003) – is to assume that the series under investigation is stationary around a small set of discrete breaks in its mean.[29] Loosely speaking, Z_t behaves as a piecewise stationary process. According to Eq. C.1, γ_t can be rewritten as

$$\gamma_t = \gamma_0 + \sum_{j=0}^{m} \gamma_j I_{\tau_j} \tag{C.2}$$

where I_{τ_j} is Heaviside indicator function – indicating that $I_{\tau_j} = 1$ if $t > \tau_j$ and 0 otherwise –, and m denotes the number of discrete breaks occurring in the unconditional mean of the Z_t series. Bai and Perron (1998, 2003) method represents a generalization of Andrews and Ploberger (1994) methodology to allow for $m > 1$ breaks occurring at unknown dates. The Bai-Perron procedure consists in determining the optimal number and location of the structural break points τ_j $(j = 1, ..., m)$ by minimizing the within-regime sums of squares. The appropriate number of breaks corresponds to the one achieving the lowest Bayesian information criterion score.[30] The estimated regimes for population growth and saving rate series are reported in Table C.1.

Table C.1: Estimated regimes

	Dates of the regimes	**Estimate**	**Standard error**
Population growth			
Regime 1	1871 – 1891	0.022762***	0.000565
Regime 2	1892 – 1914	0.018771***	0.000540
Regime 3	1915 – 1944	0.011130***	0.000473
Regime 4	1945 – 1965	0.016168***	0.000565
Regime 5	1966 – 2014	0.010128***	0.000370
Saving rate			
Regime 1	1870 – 1919	0.186673***	0.005464
Regime 2	1920 – 1940	0.165513***	0.003140
Regime 3	1941 – 1985	0.220311***	0.006118
Regime 4	1986 – 2014	0.182621***	0.004071

Notes: Population growth is measured as the change in the natural logarithm of the population (i.e., $n_t = \Delta \log(L_t)$); the saving rate corresponds to gross private saving divided by GDP (i.e., $s_t = \frac{S_t}{GDP_t}$).

[29] Examples of application of this procedure to detect shifts in the mean of macroeconomic time series are Russell (2011), Russell and Chowdhury (2013), and Clementi et al. (2015).

[30] In testing for breaks, Bai and Perron (2003) suggest to use a trimming value of 0.15 and to set the maximum number of breaks $m = 5$.

D Derivation of the system of Eqs. (5)

To derive the output equation, we follow Lee et al. (1997), Binder and Pesaran (1999) and Kutan and Yigit (2007). First, we rewrite $\Delta\log(A_t)$ and $\Delta u_{a,t}$ as follows.

$$\begin{aligned}\Delta\log(\mathrm{A}_t) &= g + \sum_{k=1}^{N}\beta_k\Delta\log(\tilde{x}_{k,t}) + \Delta u_{a,t}\\ \Delta u_{a,t} &= -(1-\phi_a)u_{a,t-1} + \epsilon_{a,t}.\end{aligned} \tag{D.1}$$

Then, we impose $\Delta\log(\mathrm{L}_t) = n$. Using Eqs. (D.1) and $\Delta\log(\mathrm{L}_t) = n$ in Eq. (3) yields:

$$\Delta\log(\mathrm{k}_t) = -(n+g) - \sum_{k=1}^{N}\beta_k\Delta\log(\tilde{x}_{k,t}) - \Delta u_{a,t} + \log(s\mathrm{k}_{t-1}^{-(1-\alpha)} + (1-\delta)). \tag{D.2}$$

As far as the notion of steady state is concerned, we acknowledge that, in our framework, the steady-state of the economy is obtained by assuming that each stochastic process and all (exogenous) variables are equal to their long-average value (i.e., $u_{a,t} = \log(\tilde{x}_{k,t}) = 0$, $t \in \{1,\dots,T\}$). We then linearise Eq. (D.2) around $\mathbb{E}[\log(\mathrm{k}_\infty)]$, where k_∞ is the random variable that underlies the steady-state distribution of k_t. By taking expectation on both sides of Eq. (D.2) we obtain:

$$(n+g) = \mathbb{E}\left[\log(se^{-(1-\alpha)\log(\mathrm{k}_\infty)} + 1 - \delta)\right]. \tag{D.3}$$

The function $f(\log(\mathrm{k}_\infty)) \doteq \log(se^{-(1-\alpha)\log(\mathrm{k}_\infty)} + (1-\delta))$ is a convex function of $\log(k_\infty)$. Then, Jensen's inequality implies:

$$(n+g) = \log(se^{-(1-\alpha)\mathbb{E}[\log(\mathrm{k}_\infty)]} + 1 - \delta) + h. \tag{D.4}$$

The parameter h is a strictly positive number which depends on the degree of the curvature of the function f. From Eq. (D.4) we easily obtain an expression for $\mathbb{E}[\log(\mathrm{k}_\infty)]$:

$$\mathbb{E}[\log(\mathrm{k}_\infty)] = \frac{1}{1-\alpha}\left[\log(s) - \log(e^{n+g-h} - 1 + \delta)\right] \tag{D.5}$$

which can be used to linearise Eq. (D.2). Specifically, let ξ_t be the approximation error. Then, the expansion of the non-linear term in Eq. (D.2) around $\mathbb{E}[\log(\mathrm{k}_\infty)] \doteq \mathrm{k}_{ss}$ yields

$$\log(se^{-(1-\alpha)\log(\mathrm{k}_{t-1})} + 1 - \delta) = \gamma - (1-\lambda)\log(\mathrm{k}_{t-1}) + \xi_t, \tag{D.6}$$

where

$$(1-\lambda) = \frac{s(1-\alpha)e^{-(1-\alpha)\mathrm{k}_{ss}}}{se^{-(1-\alpha)\mathrm{k}_{ss}} + 1 - \delta} \tag{D.7}$$

and

$$\gamma = \log(se^{-(1-\alpha)\mathrm{k}_{ss}} + 1 - \delta) + (1-\lambda)\mathrm{k}_{ss}. \tag{D.8}$$

Using Eq. (D.5), $(1-\lambda)$ and γ simplify as follows:

$$(1-\lambda) = (1-\alpha)\left[1 - (1-\delta)\,e^{-(n+g-h)}\right] \tag{D.9}$$

and

$$\gamma = n + g - h - \left[1 - (1-\delta)e^{-(n+g-h)}\right]\left[\log(e^{n+g-h} - 1 + \delta) - \log(s)\right]. \tag{D.10}$$

Notice also that for small values of n, g, δ, and h Eqs. (D.9) and (D.10) take the following form:

$$(1-\lambda) \approx (1-\alpha)(n + g + \delta - h) \tag{D.11}$$

$$\gamma \approx n + g - h - \frac{(1-\lambda)}{(1-\alpha)}\left[\log(s) - \log(n + g + \delta - h)\right] \tag{D.12}$$

Now, we derive the univariate representation for the output per capita. The production function in Eq. (1) can be expressed in terms of the logarithm of output per capita, $\log(\mathrm{Y}_t/\mathrm{L}_t) \doteq \log(\mathrm{y}_t)$, as

$$\log(\mathrm{y}_t) = \alpha\log(\mathrm{k}_t) + \log(\mathrm{A}_t). \tag{D.13}$$

Using Eqs. (D.2) and (D.1) – and the related approximations in Eqs. (D.11) and (D.12) – jointly with Eq. (D.13), we obtain:

$$\begin{aligned}
\Delta\log(\mathrm{y}_t) &= \alpha\Delta\log(\mathrm{k}_t) + \Delta\log(\mathrm{A}_t)\\
&= \alpha\left(-(n+g) - \sum_{k=1}^{N}\beta_k\Delta\log(\tilde{x}_{k,t}) - \Delta u_{a,t} + \gamma - (1-\lambda)\log(\mathrm{k}_{t-1})\right)\\
&+ g + \sum_{k=1}^{N}\beta_k\Delta\log(\tilde{x}_{k,t}) + \Delta u_{a,t}\\
&= \alpha(\gamma - (n+g)) + (1-\alpha)\sum_{k=1}^{N}\beta_k\Delta\log(\tilde{x}_{k,t}) + (1-\lambda)a_0 + (1-\lambda)gt - (1-\lambda)g\\
&+ (1-\lambda)\sum_{k=1}^{N}\beta_k\log(\tilde{x}_{k,t-1}) + (1-\lambda)u_{a,t-1} - (1-\lambda)\log(y_{t-1}) + g + (1-\alpha)\Delta u_{a,t} + (1-\lambda)u_{a,t-1}\\
&\approx -\alpha h - (1-\lambda)\frac{\alpha}{1-\alpha}\left[\log(n+g+\delta-h) - \log(s)\right] + (1-\lambda)a_0 + \lambda g + (1-\lambda)gt - (1-\lambda)\log(y_{t-1})\\
&+ (1-\alpha)\sum_{k=1}^{N}\beta_k\Delta\log(\tilde{x}_{k,t}) + (1-\lambda)\sum_{k=1}^{N}\beta_k\log(\tilde{x}_{k,t-1}) + (1-\alpha)\Delta u_{a,t} + (1-\lambda)u_{a,t-1}.
\end{aligned} \tag{D.14}$$

Rearranging terms yields:

$$\log(\mathrm{y}_t) = \mu + \lambda \log(\mathrm{y}_{t-1}) + g(1-\lambda)t + (1-\alpha)\sum_{k=1}^{N}\beta_k \log(\tilde{x}_{k,t}) + (\alpha-\lambda)\sum_{k=1}^{N}\beta_k \log(\tilde{x}_{k,t-1}) + e_t, \tag{D.15}$$

where

$$\mu = -\alpha h + \lambda g + (1-\lambda)\left[a_0 - \frac{\alpha}{1-\alpha}\left[\log(n+g+\delta-h) - \log(s)\right]\right] \tag{D.16}$$

and

$$e_t = \left[(1-\lambda) - (1-\alpha)(1-\phi_a)\right] u_{a,t-1} + (1-\alpha)\epsilon_{a,t}. \tag{D.17}$$

Putting together the Eqs. (D.15), (D.16), and (D.17), it is immediate to obtain the system of equation (5).

E Law of motion around steady state

As pointed out at the end of Section 3, we present the theoretical framework assuming constant values for the population annual growth rate and for the annual saving rate. Actually, these rates are mean-shift processes. So, we can identify a number of windows over the entire secular period considered where these growth rates are constants. We indicate with t a year inside an arbitrary window and with $k_{ss|t} \doteq \mathbb{E}_t\left[\log\left(k_\infty\right)\right]$ where $\mathbb{E}_t\left[\cdot\right]$ indicates the expectation with respect the information available at time t. Using Eqs. (D.2) and Eq. (D.6) and assuming negligible the error ξ_t the law of motion for the logarithm of the capital per effective labor around the steady state $k_{ss|t}$ can be rewritten as:

$$\log(k_t) = k_{ss|t} + \lambda(\log(k_{t-1}) - k_{ss|t}) - \Delta u_{a,t} - \sum_{k=1}^{N}\beta_k \Delta \log(\tilde{x}_{k,t}) - h. \tag{E.1}$$

Hence, making usage of Eq. (D.13) we obtain:

$$\log(y_t) = \alpha\left(k_{ss|t} + \lambda(\log(k_{t-1}) - k_{ss|t}) - \Delta u_{a,t} - \sum_{k=1}^{N}\beta_k \Delta \log(\tilde{x}_{k,t}) - h\right) + \log(A_t). \tag{E.2}$$

In order to derive the law of motion for the logarithm of per capita output we make use of the Eq. (D.13) to obtain expressions for $\log(k_{t-1})$, $k_{ss|t}$, and hence for $\log(k_{t-1}) - k_{ss|t}$:

$$\log(k_{t-1}) = \frac{\log(y_{t-1}) - \log(A_{t-1})}{\alpha} \tag{E.3}$$

$$k_{ss|t} = \frac{\log(y_{ss|t}) - \log(A_{ss|t})}{\alpha} \tag{E.4}$$

$$\log(k_{t-1}) - k_{ss|t} = \frac{\log(y_{t-1} - \log(y_{ss|t}) + \log(A_{ss|t}) - \log(A_{t-1})}{\alpha} \tag{E.5}$$

So, Eq. E.2 reads as follow:

$$\begin{aligned}
&\log(y_t) = \log(y_{ss|t}) - \log(A_{ss|t}) + \lambda \left(\log(y_{t-1}) - \log(y_{ss|t}) + \log(A_{ss|t}) - \log(A_{t-1})\right) \\
&- \alpha\Delta u_{a,t} - \alpha \sum_{k=1}^{N} \beta_k \Delta \log(\tilde{x}_{k,t}) - \alpha h + \log(A_t) \\
&= \log(y_{ss|t}) + \lambda \left(\log(y_{t-1}) - \log(y_{ss|t})\right) + (1-\lambda) \left(\log(A_{t-1}) - \log(A_{ss|t})\right) - \alpha\Delta u_{a,t} \\
&- \alpha \sum_{k=1}^{N} \beta_k \Delta \log(\tilde{x}_{k,t}) - \alpha h + \Delta \log(A_t) \\
&= g + \log(y_{ss|t}) + \lambda \left(\log(y_{t-1}) - \log(y_{ss|t})\right) + (1-\lambda) \left(\log(A_{t-1}) - \log(A_{ss|t})\right) + (1-\alpha)\Delta u_{a,t} \\
&+ (1-\alpha) \sum_{k=1}^{N} \beta_k \Delta \log(\tilde{x}_{k,t}) - \alpha h
\end{aligned} \tag{E.6}$$

Subtracting $\log(y_{t-1})$ from both sides of Eq. E.6, we get:

$$\begin{aligned}
&\Delta \log(y_t) = g - (1-\lambda) \left[\left(\log(y_{t-1}) - \log(y_{ss|t})\right) - \left(\log(A_{t-1}) - \log(A_{ss|t})\right)\right] \\
&+ (1-\alpha)\Delta u_{a,t} + (1-\alpha) \sum_{k=1}^{N} \beta_k \Delta \log(\tilde{x}_{k,t}) - \alpha h
\end{aligned} \tag{E.7}$$

Eq. E.7 helps to capture dynamics in the case of a change in the steady-state conditions generated, for instance, by a change in s and/or n. The per capita output starts a temporary growth pattern different respect to the normal growth characterized by g and short-run oscillations caused by shocks and $\Delta \log(\tilde{x}_{k,t})$. This dynamics is only temporary, because the quantity multiplied for $(1-\lambda)$ can be interpreted as an error correction term bringing output per capita growth back to its normal pattern.

F IRFs analysis

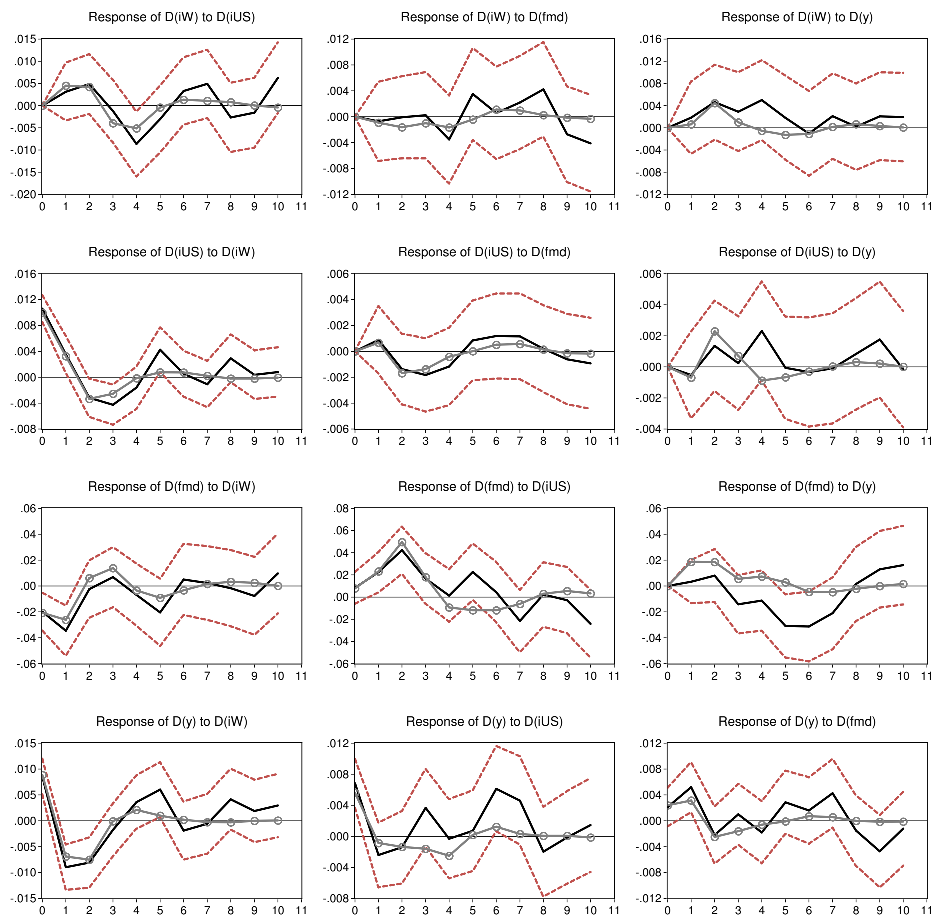
Figure F.1 depicts the impulse response function from a VAR(2) model of output growth on financial factors.[31] To avoid any potential misspecification of the data generating process IRFs are based on local projections as suggested by Jordà (2005, 2009). Output growth displays an immediate increase with a subsequent rapid fall from one year after the shock to i^W. The effect becomes statistically insignificant from three years after the shock. Differently,

[31] The lag order in VAR is selected according to AIC criterion. It is important to stress that the results presented in this section are robust to different variables ordering. The following order is considered: world interest rate, (domestic) short-term interest rate, financial development index, output.

shocks to i^{US} and fmd do not have significant effects on output growth. More importantly, there is no evidence of a significant impact of output growth shocks on financial factors.

Table F.1 documents the 5-year and 10-year ahead forecast error variance decomposition analysis. Based on the 5-year forecasting horizon, we see that 31.8% of the forecast error variance for the output growth is accounted for by innovations in the world interest rate. The other financial variables explain instead less than 12% of output variability. Overall, there is little evidence of output growth innovations contributing significantly financial factors' volatility.

Figure F.1: Response to Cholesky one S.D. Innovations (1950-2013)



Notes: $D(y)$ indicates the log of first difference of per capita GDP; $D(fmd)$ denotes the log difference of financial development index; $D(i^{US})$ is the first difference of the US short-term interest rate; $D(i^{W})$ represents the first difference of world interest rate. VAR-based IRFs are obtained by estimating a VAR(2) and identified with a standard Cholesky decomposition. The order for the Cholesky decomposition: $D(i^{W})$; $D(i^{US})$; $D(fmd)$; $D(y)$. Solid black lines: local projections IRFs Jordà (2005). Solid-dotted grey lines: VAR IRFs. Dashed-red lines: 90.0% Marginal confidence bands as described in Jordà (2009).

Table F.1: Forecast error variance decomposition.

Variance explained in	Horizon	By innovations in			
		ε_y	ε_{fmd}	$\varepsilon_{i^{US}}$	ε_{i^W}
$\Delta(y)$					
	5	56.7	4.2	7.3	31.8
	10	56.4	4.3	7.5	31.8
$\Delta(fmd)$					
	5	5.1	62.9	23.0	9.0
	10	5.2	61.4	24.1	9.3
$\Delta(i^{US})$					
	5	3.4	2.6	33.2	60.9
	10	3.6	2.8	33.4	60.2
$\Delta(i^{W})$					
	5	2.0	0.7	7.4	89.9
	10	2.3	0.9	7.6	89.2

Notes: Δy indicates the log difference of real GDP per capita; Δfmd denotes the log difference of financial market development index; Δi^{US} is the first difference of US short-term interest rate; Δi^{W} represents the first difference of world interest rate. Percentage of the forecast error variance explained by innovations in Δy, Δfmd, Δi^{US}, and Δi^{W}.

G Model selection using Akaike weights

In order to select the model that approximates the true process best, we use the AIC model selection using Akaike weights (see Anderson, 2007). We outline this criteria in the following.

The first step is to determine, for each model M_i, $i = 1, \cdots, K$, the AIC with the finite sample correction, defined as

$$\mathrm{AIC}_c^i = -2\log(\mathcal{L}_i) + 2V_i + \frac{2V_i(V_i+1)}{T - V_i - 1},$$

where $\mathcal{L}_i$ is the maximum likelihood for the candidate model i, V_i is the number of free parameters of the model, T is the sample size. Then, one computes the differences in AIC_c with respect to the AIC_c of the best candidate model, that is

$$\Delta_i(\mathrm{AIC}_c) = \mathrm{AIC}_c^i - \min_{i\in\{1,\cdots,K\}}(\mathrm{AIC}_c^i).$$

$\Delta_i(\mathrm{AIC})$ takes into account the relative performance of the models. From the differences in AIC_c the Akaike weights $w_i(\mathrm{AIC}_c)$

$$w_i(\mathrm{AIC}_c) = \frac{\exp(-\frac{1}{2}\Delta_i(\mathrm{AIC}_c))}{\sum_{k=1}^{K}\exp(-\frac{1}{2}\Delta_k(\mathrm{AIC}_c))}$$

are computed. Weight $w_i(\mathrm{AIC}_c)$ can be interpreted as the probability that M_i is the best model, given the data and the set of candidate models. Finally, the logarithm of the so-called

Evidence Ratio, $\mathrm{LER}_i = \log_{10}(\mathrm{ER}_i)$

$$\mathrm{ER}_i = \frac{w_{best}(\mathrm{AIC}_c)}{w_i(\mathrm{AIC}_c)}$$

is computed. The LER_i is a quantitative measure of the strength of the evidence of the best model vs. any other models (i.e., a relatively small value of LER_i suggests that model i represents the best model).