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Finance-augmented business cycles: A robustness check

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Abstract

Recent literature has highlighted the importance of considering the financial cycle for the estimation of business cycles. The applied estimation approaches, however, differ widely and cyclical estimates are therefore difficult to compare. In this paper, we assess the robustness of finance-augmented business cycle estimates to different trend specifications for Japan, the UK, and the US. In line with earlier studies, we confirm that the inclusion of financial variables strongly affects the estimates of the business cycle, resulting in larger amplitudes and more persistent dynamics than traditional cycle estimates. While the dynamics of the cyclical component does not depend much on the model used, its amplitude shows strong sensitivity to the underlying assumptions of the trend model.

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1. Introduction and previous literature

The global financial crisis and the following deep recession have highlighted the tight link between financial and economic developments. Prior to the crisis, financial developments played only a minor role in the information set of economic policy-makers. As a response to the global financial collapse, recent literature suggests that financial variables should explicitly be taken into account for the estimation of business cycles (see, for instance, Borio et al. 2013, and Bernhofer et al. 2014).

Theoretical studies showed that the effects of various types of shocks, either financial or real, can be magnified and transmitted through the financial accelerator and other related mechanisms such as credit cycles, leverage cycles and developments of housing markets (Bernanke et al. 1999, Kiyotaki and Moore 1997, Mendoza 2010, Iacoviello 2005, Iacoviello and Pavan 2013).[1] Nolan and Thoenissen (2009) used a calibrated model to show that financial shocks (i.e. shocks to the efficiency of the financial sector) in the United States (US) are tightly connected with the inception of recessions, even more so than total factor productivity (TFP) or monetary shocks. Financial shocks also account for a large part of the variance of Gross Domestic Product (GDP), and remain contractionary even after recessions have ended. A recent paper by Iacoviello (2015) also underlines the financial, rather than real character of business cycles in the context of a Dynamic Stochastic General Equilibrium (DSGE) model. Alongside its interlinkages with credit markets, the housing market itself also plays an essential role in the business cycle (see, e.g., Leung 2004, Iacoviello and Minetti 2008, Iacoviello 2010), not only because housing wealth is an important component of national wealth, but also because housing investment is a very volatile component of aggregate demand, and further, because housing markets show significant cyclical fluctuations and volatility (Bernanke and Gertler 1995).

The empirical evidence mostly confirms the linkages among credit, housing and business cycles suggested by theory, but the magnitude of the effects differs widely. Early studies on the impact of credit conditions on business cycles using times series methods—mostly vector autoregressions (VARs)—confirmed the important role of credit for the business cycle (Stock and Watson 1989, Friedman and Kuttner 1992, 1993, Kashyap et al. 1993).[2] Borio et al. (2001) showed extensive descriptive evidence on the various linkages between finance and real economic activity. They concluded that at least ex-post, a financial cycle is clearly apparent, as periods of strong economic growth tend to be associated with significant increases in the ratio of credit to GDP, and recessions with declines in this ratio. Similarly, strong credit growth tends to be linked to increases in equity and property prices. Balke (2000) examined the role of credit as a nonlinear propagator of shocks by using nonlinear impulse response functions from threshold VARs and found that shocks to credit have a larger effect on output in a "tight" credit regime as compared to

[1] Davis and Van Nieuwerburgh (2014) have recently reviewed the literature on the interconnection of macroeconomics, finance, and housing.

[2] Still, Ramey (1994) found rather mixed results.

"normal" times. Helbling et al. (2011) analyzed the importance of credit market shocks for the G-7 countries using a series of VAR models. Their variance decompositions showed that credit market shocks are as important as productivity shocks in driving business cycles. In a similar vein, several other studies used VAR models to analyze the impact of credit supply shocks on the business cycle (see, for instance, Gilchrist et al. 2009, Peersman 2012, Meeks 2012, Gambetti and Musso 2012, Hristov et al. 2012, Gilchrist and Zakrajsek 2012, Bijsterbosch and Falagiarda 2015), or Factor VAR models to explore the global impact and the synchronization of credit supply shocks (Eickmeier and Ng 2015, Miranda-Agrippino and Rey 2015). In general, these studies confirm a significant impact of credit shocks on the business cycle as well as their international transmission.

With respect to the impact of housing, a number of papers have documented a significant co-movement between housing markets and the macroeconomy (see Leung 2004 for a summary). While fluctuations in residential investment tend to lead real GDP in the US (Leamer 2007, Davis and Heathcote 2005), they coincide with the business cycle in most other industrialized countries (Kydland et al. 2012). Furthermore, housing market swings generally show considerable co-movement with consumption expenditures (Iacoviello and Neri 2010). Musso et al. (2011) analyzed the role of the housing market in the macroeconomy by using structural VARs. They found that adverse (mortgage) credit supply shocks in the US tend to decrease consumption and that positive non-monetary housing demand shocks tend to rise consumption. Against this backdrop, several VAR-based studies found a substantial effect of the housing market on business cycles (see, e.g., Bagliano and Morana 2012, Cesa-Bianchi 2013, Walentin 2014).

Finally, some empirical studies examined the linkages between credit and housing markets and their influence on business cycles. Stock and Watson (1999) found a high degree of co-movement in house prices, bank credit and real activity for the US when applying a band-pass filter to the respective times series and examining correlations at various leads and lags across those time series. More recently, Goodhart and Hofmann (2008) investigated the links between money, credit, house prices and economic activity in industrialized countries by applying a fixed-effects panel VAR model to data of 17 industrialized countries. They found a multidirectional link between credit aggregates, house prices and the macroeconomy. Moreover, this effect seems to become stronger towards the end of their sample, and shocks to money and credit are stronger in the case of booming house prices. Igan et al. (2011) examined the cyclical co-movement in advanced economies by using a dynamic generalized factor model. They concluded that house prices tend to lead credit and business cycles over the long term, while in the short to medium term the relationship varies across countries. The lead-lag relationships between house price, credit and real activity cycles vary considerably across countries, suggesting that the financial accelerator mechanisms also differ in their magnitude and dynamics.

The role of developments in credit, housing, and asset markets in the origin of the recent financial crisis has also been noted by several authors. Babecký et al. (2013) found that domestic housing prices, share prices, credit growth, and private credit constitute good early warning indicators of the costs of crisis. Jordá et al. (2014) concluded that mortgage boom-bust episodes are followed by considerably slower GDP growth, irrespective of whether a financial crisis occurred. Claessens et al. (2012) found strong interactions between business and financial cycles by using simple characteristics of cyclical phases such as duration, amplitude and slope. One of their main findings is that the duration and amplitude of recessions and recoveries are often shaped by linkages between business and financial cycles. Moreover, according to their results, financial disruptions tend to increase the length and depth of recessions, and recoveries following asset price busts tend to be weaker. Rapid growth in credit and house prices, on the contrary, is often associated with stronger recoveries. Schüler et al. (2015) used a multivariate spectral approach to characterize a synthetic financial cycle based on credit, residential property prices, equity prices, and bond yields and found that this indicator outperforms the credit-to-GDP gap in predicting financial crises on a horizon of up to three years. Finally, Iacoviello (2015) noted that financial shocks were at the core of the last recession in the US, accounting for two-thirds of the decline in private GDP during the 2007–2009 recession.[3]

Previous empirical research has also documented the existence of long financial cycles. Drehmann et al. (2012) characterized low frequency financial cycles, which are mainly driven by credit and property prices. Aikman et al. (2015) provided further evidence for low frequency credit cycles and formulated a theoretical framework in which small changes in fundamentals can yield large swings in credit. Stremmel (2015) developed and compared several synthetic medium-term financial cycle measures for 11 European countries, concluding that the best fitted financial cycle measure for these countries combines information on the credit-to-GDP ratio, credit growth and the house-prices-to-income ratio. The periodicity of housing cycles is also longer and amplitudes are larger than those of typical business cycles (see Leung 2004, and the articles reviewed there).

Despite of its high relevance, however, information on the financial cycle has typically not been incorporated in the design of macroeconomic policies. In particular, output gap estimates are normally based on the concept of inflation-neutral potential output (Borio et al. 2013). Incorporating financial information in the estimation of business cycles may therefore provide superior information for policymaking (Borio 2014). In this context, Borio et al. (2013) suggested extending the concept of inflation-neutral potential output by explicitly incorporating financial information in the estimation of business cycles. They proposed a finance-augmented dynamic (FAD) Hodrick-Prescott (1997, HP) filter where the financial cycle is proxied by the growth rates of house prices and credit to non-financial sectors. Bernhofer et al. (2014) estimated a structural unobserved components (SUC) model with a smooth trend component specification (Harvey 1989, Harvey and Jaeger 1993). The authors used the finance-augmented output gaps to estimate a Phillips curve (Borio et al. 2013) and an Okun's law (Bernhofer et al. 2014), thereby improving

[3] Other studies emphasizing the importance of credit developments and their role in crises include Reinhart and Rogoff (2011), Claessens et al. (2011), Gourinchas and Obstfeld (2012), Schularick and Taylor (2012), and Mendoza and Terrones (2012).

the fit of both macroeconomic relationships significantly as compared to standard approaches. However, some properties of the business cycle estimate such as its amplitude and duration may be sensitive to the filter used, particularly when additional explanatory variables are taken into account. In the context of finance-augmented output gaps, however, the sensitivity of the respective estimates to underlying trend specifications has not been tested so far. This note aims at filling this gap in the literature.

We estimate the cyclical components of real GDP in Japan, the United Kingdom (UK), and the US using a SUC model in the spirit of Harvey (1989) and Harvey and Jaeger (1993). Following Borio et al. (2013), we incorporate financial information captured by the growth rate of house prices and credit to non-financial sectors in the specification of the cyclical component. Our model's trend specification nests those of the FAD-HP filter by Borio et al. (2013) and the SUC model with smooth trend by Bernhofer et al. (2014). The proposed trend specification can accommodate sharp movements in GDP series associated with events such as deep recessions (Crespo-Cuaresma and Fernández-Amador 2013). We test the robustness of the estimates under these particular trend specifications and compare the results of those different specifications for the three countries mentioned above.

The three countries are selected as illustrative examples as they are crucial for the determination of global liquidity and worldwide macroeconomic conditions. They have experienced pronounced boom-bust cycles caused by tight credit conditions and the real estate sector that resulted in significant deleveraging, while both timing and policy responses have differed widely across the selected economies. We confirm previous research results that the inclusion of financial variables strongly affects the estimated dynamics of the business cycle. We also find that the dynamics do not depend much on the model or specification used in the estimation. Notwithstanding, the amplitude of cyclical components shows a strong sensitivity to the underlying assumptions of the trend model, which is highly important for the use of finance-augmented output gaps in economic policy.

2. The model

We extend the SUC model of Harvey (1989) and Harvey and Jaeger (1993) and decompose the logarithm of (quarterly) real GDP y_t as

$$y_t = \tau_t + \psi_t + u_t \ , \tag{1}$$

where τ_t is the trend component, ψ_t is the finance-augmented cyclical component, and u_t is the (white noise) irregular component. In its most general specification, the trend follows a random walk with a drift where the drift also follows a random walk. Thus,

$$\begin{aligned} \tau_t &= \tau_{t-1} + \beta_{t-1} + \varepsilon_t^\tau, \ \ \varepsilon_t^\tau \sim \mathbf{NID}(0, \sigma^2_{\varepsilon^\tau}), & (2) \\ \beta_t &= \beta_{t-1} + \varepsilon_t^\beta, \ \ \varepsilon_t^\beta \sim \mathbf{NID}(0, \sigma^2_{\varepsilon^\beta}) & (3) \end{aligned}$$

This specification of the trend nests various interesting cases. The variances $\sigma^2_{\varepsilon^\tau}$ and $\sigma^2_{\varepsilon^\beta}$ govern the stochastic variability in the shift and the slope of the trend, respectively. If $\sigma^2_{\varepsilon^\tau} > 0$ and $\sigma^2_{\varepsilon^\beta} > 0$, this component induces an $\mathsf{I}(2)$ trend on y_t. When $\sigma^2_{\varepsilon^\tau} > 0$ and $\sigma^2_{\varepsilon^\beta} = 0$, τ_t is a random walk with drift. The case $\sigma^2_{\varepsilon^\tau} = 0$ and $\sigma^2_{\varepsilon^\beta} > 0$ defines a smoothly changing trend. $\sigma^2_{\varepsilon^\tau} = 0$, $\sigma^2_{\varepsilon^\beta} = 0$, and the initial state of β different from zero ($\beta_0 \neq 0$) imply a deterministic linear trend. The model (1)-(3) nests the HP filter when $\sigma^2_{\varepsilon^\beta}/\sigma^2_{\varepsilon^y} = \lambda = 1600$ for quarterly data, and $\sigma^2_{\varepsilon^\tau} = \psi_t = 0$. Thus, the HP estimate of the cyclical component is the smoothed irregular component (Harvey 1989, Harvey and Jaeger 1993).[4] The finance-augmented cyclical component follows a damped stochastic sine-cosine wave specified as

$$\begin{aligned} \psi_t &= \gamma_1 House_t + \gamma_2 Credit_t + \rho \cos\omega\psi_{t-1} + \rho\sin\omega\psi^*_{t-1} + \kappa_t \qquad \kappa_t \sim N(0, \sigma^2_\kappa), \quad (4) \\ \psi^*_t &= -\rho\sin\omega\psi_{t-1} + \rho\cos\omega\psi^*_{t-1} + \kappa^*_t \qquad \kappa^*_t \sim N(0, \sigma^2_\kappa) \quad (5) \end{aligned}$$

$\rho \in [0, 1]$ is the damping factor, $\omega \in (0, \pi)$ is the cycle's frequency measured in radians, and $\mathbf{\Sigma}_\kappa = \text{diag}(\sigma^2_\kappa, \sigma^2_\kappa)$, i.e. the disturbances of the cyclical component are assumed independent and of equal variance. The cycle given by (4)-(5) follows a stationary ARMA(2,1) process and the constraints on the parameter space given above restrict the roots of the lag polynomial to lie in the region of the parameter space that leads to pseudo-cyclical behavior in ψ_t. We follow Borio et al. (2013) and include two proxies for the financial cycle in the specification of the cyclical component; $House_t$ refers to the growth rate of real house prices and $Credit_t$ is the growth rate of real credit.

The model can be written in state-space form and estimated by maximum likelihood via the Kalman (1961) filter and prediction error decomposition, constrained by the above-mentioned restrictions over the parameters space and well-behavior of the cyclical component. In particular, the estimation is carried out as follows. First, we draw initial states from uniform distributions and estimate the model using the Kalman filter. Likelihood is maximized across draws subject to the constraint of well-behaved parameters and states estimates. Besides the restrictions over the parameters space mentioned above, we discard estimations that result in cyclical components with long tails at the end of the sample period. Afterwards, the smoothed estimate of the cyclical component is retrieved.

The proposed model extends and nests the approaches of Borio et al. (2013) and Bernhofer et al. (2014). Borio et al. (2013) augment the HP filter by including both credit and house price growth rates in the cyclical component and by adding the lag of the cycle. They restrict the noise-to-signal ratio to that of the HP filter, i.e. the trend captures long-term swings of 8 years or more. However, the HP filter has been subject to criticism by Harvey and Jaeger (1993) and Cogley and Nason (1995), since it may produce spurious cycles and its effects depend on the properties of the observed series to be decomposed. We do not restrict the cyclical component to an upper bound of 8 years. Instead, our model implicitly specifies a band of frequencies which corresponds to the business cycle. Therefore, we do not restrict the frequency of the cycle and stay

[4] When $\lambda \to \infty$, the HP filter approaches to linear detrending.

closer to the original definition of Burns and Mitchell (1946) according to which a full cycle must last for at least six quarters and a maximum of 12 years. If the financial cycle is longer than the business cycle and if we allow for some effect of the former on the latter, the financial cycle may widen the duration of the business cycle to some extent. While Bernhofer et al. (2014) assume a smoothly-changing trend component, the general specification employed in the present paper can account for sharp movements in GDP series connected to crises in transition economies, shocks to trend growth in real variables in emerging markets, or deep recessions that are associated with more volatility in the trend component (see Crespo-Cuaresma and Fernández-Amador 2013). In the following section, we illustrate the differences across various approaches to highlight the impact of differing trend specifications.

Real GDP data are from the International Monetary Fund's IFS database. House prices are from the Bank of International Settlements (BIS) property price statistics. Credit is total credit to non-financial sectors (incl. cross-border credit) provided by the BIS (see Dembiermont et al. 2013 for details). House price and credit series are deflated using IMF consumer prices. All data are at a quarterly frequency. The sample covers the periods 1979q2–2012q4 (Japan), 1988q2–2012q4 (UK), and 1979q2–2012q4 (US).

3. Results

Table I displays the coefficients of the financial indicators estimated using the model specified by (1)-(5) (full-trend), the smooth trend model of Bernhofer et al. (2013) (BFAGS), and the FAD-HP filter. Both the BFAGS and the FAD-HP filters are special cases of the full-trend model. In particular, the BFAGS model incorporates the a priori restriction that $\sigma^2_{\varepsilon^\tau} = 0$ and $\sigma^2_{\varepsilon^\beta} > 0$ such that it defines a smoothly changing trend component. The FAD-HP filter contains the restrictions that correspond to the trend of the HP filter; that is, $\sigma^2_{\varepsilon^\beta}/\sigma^2_{\varepsilon^y} = 1600$ and $\sigma^2_{\varepsilon^\tau} = \psi_t = 0$ such that the cyclical component is the (smoothed) irregular component u_t from equation (1). The FAD-HP filter adds the term δu_{t-1} in equation (1) to account for autocorrelation in the cyclical component. Consistent with the estimates of Borio et al. (2013) and to avoid nonstationarity problems in the cycle, we restrict the parameter of the lagged irregular component δ to be equal or less than 0.85.[5]

The growth rates of both credit and house prices have contemporaneous effects that are mostly significant, particularly in the full-trend specification. In comparison, the coefficients in the FAD-HP and BFAGS model are not conclusive about the significance of house prices and credit. When the contemporaneous effects for the BFAGS model are significant, their magnitude is rather similar to the full-trend model, especially for the UK and the US. The total effects remain ambiguous, however, since they depend on the estimated autocorrelation of the sine-cosine wave model of the cyclical component.

[5] However, our FAD-HP filter differs from Borio et al. (2013), since we do not include lagged credit and house price growth rates and our estimation is not based on a Bayesian framework.

Thus, for illustrative purposes, Figure 1 presents the smoothed cyclical components from the three models and the cycle estimated with a standard HP filter.

Table I: Contemporaneous effect of the financial indicators

	Japan			**United Kingdom**			**United States**		
	FAD-HP	BFAGS	Full-trend	FAD-HP	BFAGS	Full-trend	FAD-HP	BFAGS	Full-trend
House	0.0623	0.2400*	0.6866***	0.1086***	0.0489	0.1007***	0.0780	0.1372***	0.1475***
	(0.1308)	(0.1352)	(0.1820)	(0.0191)	(0.0326)	(0.0308)	(0.0549)	(0.0343)	(0.0390)
Credit	0.1684*	0.0622	0.8588***	0.0460*	0.1127***	0.1461***	0.2478***	0.1997***	0.1840**
	(0.0979)	(0.1093)	(0.1524)	(0.0248)	(0.0275)	(0.0298)	(0.0827)	(0.0658)	(0.0733)

In all three economies, the estimates taking into account financial information present rather similar dynamics, but differ widely from a standard HP filter. In line with earlier studies, finance-augmented cyclical components show more persistent and accentuated deviations from the potential, reflecting the main financial developments in the economies considered. The credit crunch led to extreme drops in the output gap in the beginning of the 1990s in Japan and in 2007–08 in the UK and the US. Subsequently, the process of deleveraging has retained real GDP significantly below potential. Thus, the stagnation of these economies after their respective financial collapses (Ueda 2012) is clearly captured. In particular, the output gap in Japan has remained negative since the beginning of the 1990s until the end of the sample, reflecting the depression associated with the problems in the banking sector (Hoshi and Kashyap 2004, Ueda 2012), which is also consistent with the theory of zombie banks and firms (Caballero et al. 2008). Noteworthy, in contrast to the standard HP filter, the finance-augmented estimates show persistently negative output gaps at the end of the sample, in line with current official estimates.

While the dynamics of finance-augmented cyclical components are rather similar across models, they significantly differ in their amplitude and persistence. These differences can be crucial for the determination of appropriate monetary and fiscal policy reactions. In this context, it may be convenient to specify a sufficiently general model for the trend when estimating finance-augmented output gaps, since it avoids imposing a priori constraints on the estimation. Furthermore, the full-trend model is more general and "lets the data speak", i.e. it is subject to fewer limitations than the other two models; it still nests the other two specifications. The output gap estimated by the full-trend model accounts for a sharp decline in Japan at the beginning of the 1990s, coincident with the financial collapse. On the contrary, this adjustment is slower and smaller in magnitude for the FAD-HP and BFAGS estimates. The full-trend estimates are more similar to those of the BFAGS in the end of the sample in Japan. Regarding the UK, the FAD-HP estimate differs from the other two models only during 2002–2008 and in the end of the sample. Finally, all models including financial indicators show similar results for the US. Nevertheless, the estimates from the BFAGS model differ from those of the full-trend

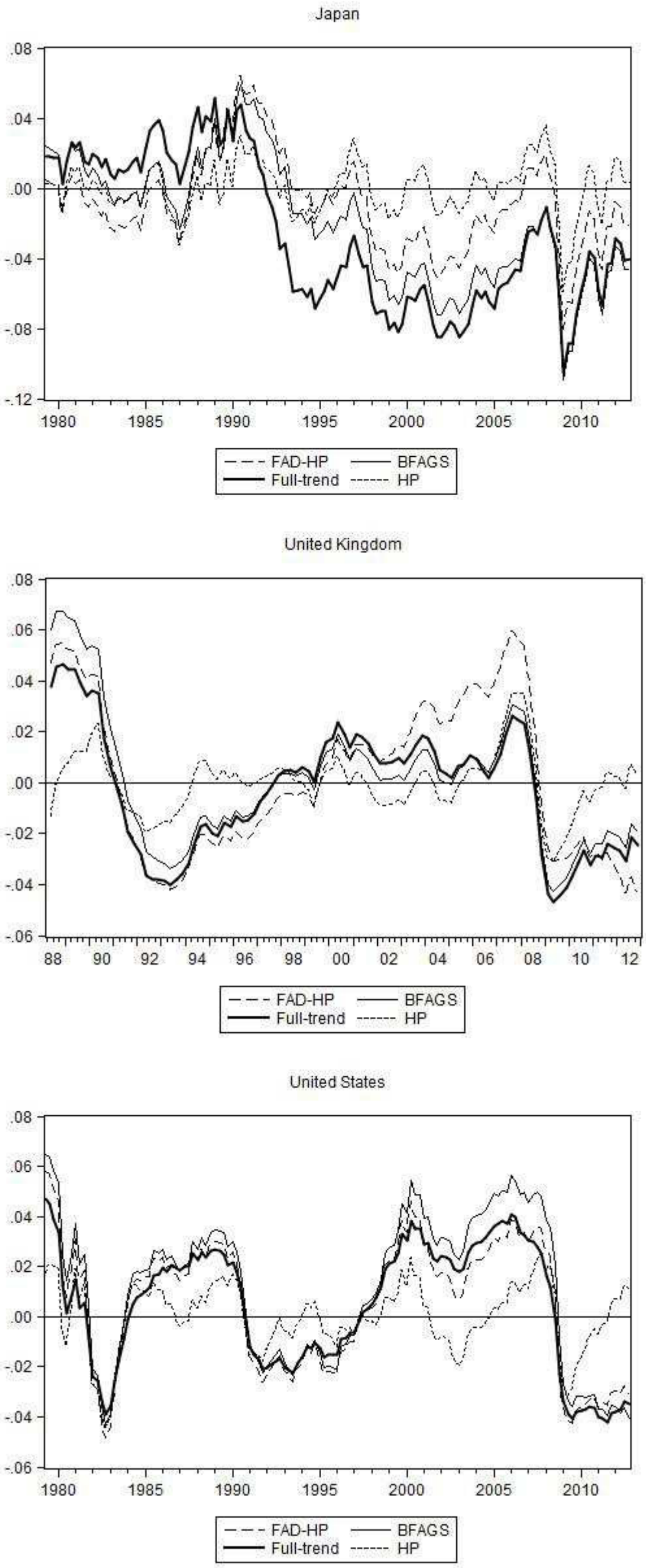

Figure 1: Estimated cyclical components of GDP

and FAD-HP models in the end of the sample. In particular, the estimates using the FAD-HP and the full-trend models show smoother deviations from the potential than the BFAGS during the expansion in the 2000s as a result of a more volatile trend. Thus, the full-trend model estimate converges to the FAD-HP filter in the case of the US GDP.

4. Discussion

The policy implications of this short paper can broadly be summarized in two main issues. First, in line with earlier studies, we confirm a considerable impact of financial cycle variables on business cycle estimates, as illustratively shown for Japan, the UK, and the US. Credit constraints and house prices affect income fluctuations and should therefore be considered in the estimation of cyclical estimates. The considerable difference in the magnitude of output gaps and the higher persistence of corresponding finance-augmented estimates highlight the need for a modification of monetary policy beyond inflation targeting (Borio 2014). Second, our study might also serve as a cautionary note for the use of finance-augmented business cycle estimates in the context of economic policy-making. Our results underline the sensitivity of some properties of the output gap estimates to the definition of the trend in the filter used. Even though the dynamics of estimated cyclical components do not depend much on the corresponding trend specification, both the amplitude and the persistence of the cycle show a strong sensitivity to underlying assumptions of the trend model. We propose the use of a general specification for the trend in order to avoid imposing a priori constraints on the estimation of output gaps. Further research may consider generalized stochastic cyclical processes in spirit of Harvey and Trimbur (2003) and Trimbur (2005). These generalized stochastic cyclical components have an autoregressive moving-average (ARMA) reduced form and may be used in unobserved components models to extract smoother cycles in economic series. Thus, further research seems necessary not only to highlight the importance of the financial cycle for business cycles but also to simultaneously ensure the robustness of corresponding estimates for the purpose of effective policy-making.

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