Economic Systems 44 (2020) 100730

Contents lists available at ScienceDirect

Economic Systems

journal homepage: www.elsevier.com/locate/ecosys

# Growth and real business cycles in Vietnam and the Asean-5. Does the trend shock matter?

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ARTICLE INFO

*Keywords:*
Vietnam
ASEAN
DSGE
Real business cycles
Trend shock
Growth

ABSTRACT

We examine Vietnam's economy in comparison with its closest trade partners. We show that capital accumulation has been the primary growth engine since the start of its transition to the pro-market economy in 1986 – the Doi Moi. We also show that the cyclical behavior of its macro aggregates is similar to that of its ASEAN-5 peers and other developing countries. We extend the standard small open economy RBC model by considering habit persistence and government consumption, which allows a close match of the moments of the growth variables. At the business cycle frequency, transitory productivity shocks account for approximately one-half of Vietnam's output variance, while country risk and non-transitory productivity shocks account to close to one-fifth each. Regarding the Solow residual's volatility, we find that the trend component merely accounts for 12 % of this variance in Vietnam, while in Thailand it is only 6 %. These findings refute the "the cycle is the trend" hypothesis in Aguiar and Gopinath (2007) and align with the hypotheses in García-Cicco et al. (2010) and Rhee (2017), where the stationary component is overwhelmingly dominant. We claim that technological progress and productivity-enhancing measures are fundamental for Vietnam's economy to sustain high growth.

## 1. Introduction

This paper examines Vietnam's economy from a real business cycle (RBC) perspective and compares its performance to that of the ASEAN-5 economies (Indonesia, Malaysia, the Philippines, Singapore and Thailand). Vietnam provides a compelling case for analysis since it is a young and dynamic economy that has achieved many socio-economic successes over the past three decades.

The contribution of our research is threefold. First, we extend the RBC models developed by Aguiar and Gopinath (2007) (AG) and García-Cicco et al. (2010) (GPU). We show that our extended setup provides a better account of the facts. In particular, we are able to match the downward slope autocorrelation between output and trade balance (as percent of GDP), which was an unresolved issue in AG's model. Second, we account for sources of growth since the implementation of the Doi Moi in 1986 and show that the characteristics of Vietnam's business cycles are not much different from its ASEAN-5 peers or other emerging market economies. Third, we provide evidence that Vietnam's business cycles have mainly been driven by transitory total factor productivity (TFP) shocks rather than by trend innovations, as claimed by AG for a set of 13 emerging economies. Given that the non-stationary component of the TFP shock only accounts for 12 % of the Solow residual volatility (6 % in Thailand), for these cases we refute their view that "the cycle is the trend" in emerging markets.

Our first task is to account for the supply-side factors that have driven Vietnam's economic growth. We show that TFP accounts for

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https://doi.org/10.1016/j.ecosys.2019.100730
Received 12 July 2018; Received in revised form 24 November 2018; Accepted 22 December 2018
Available online 07 November 2019

one third of economic growth on average in 1981–2015, while it falls to less than one fifth in the 2000s. In addition, capital accumulation becomes the main driver of growth from 1992, with the contribution of human capital increasing. Then, we use well-known filtering techniques (see, *e.g.*, Prescott, 1986; King and Rebelo, 1999; Stock and Watson, 1999, to name a few) to account for Vietnam's business cycle fluctuations in the period from 1986 to 2015. We find that the business cycle characteristics of Vietnam's national account components are essentially consistent with those reported in the literature for other emerging markets. Within this context, a significant difference is that Vietnam's cyclical output fluctuation is less than half of the ASEAN-5's average.

Provided with this information, we aim to quantify the exogenous forces that have shaped the dynamics of Vietnam's growth aggregates. Hence, we develop and estimate a dynamic stochastic general equilibrium (DSGE) RBC model to test the power of RBC theory in explaining Vietnam's economic fluctuations.

In their highly influential work, AG asserted that the business cycle facts of a sample of thirteen developing countries can be adequately captured by a standard small open economy RBC model equipped with some real frictions. Specifically, they argued that the behaviours of consumption and the trade deficit depend on the nature of shocks to output growth. If a positive productivity shock is temporary, the resulting shift in consumption will lead to a proportional increase in output. In contrast, a permanent shock not only raises the current income but also gives rise to reduced savings (or investment), inducing a negative trade balance. AG showed that non-stationary shocks to productivity bear the main responsibility for output growth variations and suggested that in emerging markets the cycle is the trend.

Unsurprisingly, this view has received support and objections in equal measure in subsequent studies. For example, Suzuki (2018a,Suzuki, 2018b provided evidence that the business cycles of two emerging economies such as Serbia and South Africa are driven by the trend. Miyamoto and Nguyen (2017), employing a sample of seventeen small open economies, found that even if the trend role is not as important as in AG, the average contribution of trend productivity shocks to economic fluctuations is slightly above 30 %.

On the other side, GPU pointed out that AG-type models could neither replicate the downward slope of the trade balance-to-output autocorrelation function nor explain some crucial moments of the long-time series of Argentina and Mexico. GPU augmented AG's model with financial frictions and country risk shocks and claimed that non-stationary productivity shocks only contributed a small fraction to the output variance. Along the same line, Boz et al. (2011); Alvarez-Parra et al. (2013) and Rhee (2017), among others, provided support to GPU's view. For instance, by considering a recursive utility function and an endogenous risk premium channel, Rhee (2017) showed that transitory productivity shocks significantly drive Korea's economy.

Taking an intermediate viewpoint, Cao et al. (2016) postulate that two conditions for permanent shocks to dominate the cycle are an insensitive risk premium at any debt level and time-separable preferences. Whereas the second condition can be satisfied by a suitable utility function, the absence of a risk premium reaction causes the AG hypothesis to hold only in extremely unrealistic cases. In the same vein, after revising the literature, Durdu (2013) concludes that the explanatory power of trend and/or interest rate shocks are magnified if the model contains a rich friction structure.

Accordingly, departing from the models by AG and GPU, we enrich GPU's financial friction setting by incorporating into the period utility function: (1) internal habit persistence (Boldrin et al., 2001), and (2) government consumption (Christiano and Eichenbaum, 1992). In this way, our proposed RBC model outperforms GPU's financial friction specification when reproducing the moments of the growth variables at the time that matches the low value of the trade balance-to-output autocorrelation (0.18 after four lags). Although the presence of habit formation improves the moment matching capability, we acknowledge that the short-run observational dynamics cannot be emulated entirely. Nevertheless, neither the plain RBC nor the AG-type model is able to deliver a better performance.

The long-run variance decomposition reveals that transitory productivity shocks explain approximately 50 % of Vietnam's output volatility. Moreover, the transitory standard deviation is higher than its non-stationary counterpart by 25 %. Another 41 % of the variance is accounted for by the exogenous risk premium and trend shocks, which have a similar contribution, while shocks to consumption preferences absorb the remaining 9 % of the variance. These findings critically imply that AG's claim that the cycle is the trend cannot be sustained in the case of Vietnam, where non-stationary shocks to productivity only account for 12 % of the Solow residual's volatility.

In order to gain insight into policy implications, we perform counterfactual simulations under three scenarios. First, we simulate the trajectories of output growth and the trade balance-to-output ratio by sequentially turning off several exogenous processes: (i) by disabling trend (transitory) productivity shocks or interest rate shocks; (ii) by substituting the actual productivity shocks in 2008–2013 by the values they took in 1992–1997; and (iii) by assuming no interest rate shocks in 2008–2013 as if no (financial) crisis had taken place. The outcome of these counterfactual analyses is twofold. First, transitory productivity shocks have a significant impact on Vietnam's income growth but not on the trade balance. Neither trend nor productivity shocks can explain the huge trade deficits experienced in 2007–2010. Instead, interest rate shocks greatly govern the trade balance and help stabilize the growth path of Vietnamese output. This second result, which is striking since it unveils strong real effects of Vietnam's proactive monetary policy in the past two decades, is consistent with the findings in Anwar and Nguyen (2018). In addition, our simulations strengthen Huynh et al.'s (2017) claim that the Vietnamese monetary policy was unable to counterbalance the economic downturn through the managing of interest rates, and that loan supply should have been directed toward productivity generating sectors.

We conclude the analysis by examining Thailand, which is Vietnam's main competitor and has historically been the star economy within the ASEAN-5. The explanatory power of both the transitory and trend shocks is relatively weak because each productivity innovation accounts for only one-fourth of the output growth variance. In this context, the trend component of the shock only accounts for 6 % of the Solow residual's variance. In addition, in contrast to Vietnam, but similarly to Korea (Rhee, 2017), country risk premium innovations appear to govern Thailand's business cycles. We therefore argue that Thailand's economy was more

**Table 1**
Descriptive statistics of the main aggregate variables.

| | | VIETNAM | | | | | | | ASEAN-5 | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Year | Stats | *y* | *c* | *g* | *i* | *x* | *m* | *tby* | *y* | *c* | *g* | *i* | *x* | *m* | *tby* |
| 1986 - 2015 | Mean | 4.84 | 3.86 | 5.18 | 9.02 | 9.73 | 9.78 | −0.06 | 3.74 | 3.57 | 3.34 | 4.15 | 6.64 | 6.37 | 0.06 |
| | SD | 1.74 | 2.39 | 3.58 | 12.17 | 8.48 | 8.81 | 0.05 | 2.92 | 2.54 | 2.34 | 9.75 | 5.71 | 7.85 | 0.04 |
| | Min | 0.27 | −0.21 | −6.33 | −12.78 | −13.26 | −12.04 | −0.17 | −9.25 | −6.33 | −4.56 | −37.66 | −9.89 | −13.84 | −0.01 |
| | Max | 7.40 | 8.40 | 10.55 | 41.34 | 29.33 | 29.02 | −0.02 | 6.67 | 8.94 | 8.38 | 17.45 | 13.98 | 17.84 | 0.10 |
| 1986 - 2000 | Mean | 4.40 | 2.48 | 3.94 | 10.86 | 9.36 | 9.20 | −0.03 | 3.78 | 3.66 | 2.31 | 3.90 | 8.93 | 8.16 | 0.03 |
| | SD | 2.27 | 1.95 | 4.61 | 16.11 | 11.83 | 11.56 | 0.01 | 3.92 | 3.54 | 2.59 | 13.63 | 4.42 | 8.30 | 0.03 |
| | Min | 0.27 | −0.21 | −6.33 | −12.78 | −13.26 | −12.04 | −0.04 | −9.25 | −6.33 | −4.56 | −37.66 | −0.90 | −13.69 | −0.01 |
| | Max | 7.40 | 7.15 | 10.06 | 41.34 | 29.33 | 29.02 | −0.02 | 6.67 | 8.94 | 4.86 | 17.45 | 13.98 | 17.84 | 0.10 |
| 2001 - 2015 | Mean | 5.29 | 5.24 | 6.43 | 7.19 | 10.09 | 10.37 | −0.08 | 3.69 | 3.48 | 4.38 | 4.41 | 4.34 | 4.57 | 0.08 |
| | SD | 0.82 | 1.99 | 1.42 | 6.35 | 2.92 | 5.14 | 0.05 | 1.53 | 0.91 | 1.54 | 3.32 | 6.06 | 7.21 | 0.01 |
| | Min | 3.99 | 1.24 | 4.26 | −9.22 | 3.99 | 2.93 | −0.17 | 0.43 | 0.98 | 1.19 | −1.24 | −9.89 | −13.84 | 0.07 |
| | Max | 6.36 | 8.40 | 10.55 | 20.71 | 14.87 | 23.46 | −0.03 | 6.56 | 4.77 | 8.38 | 8.90 | 13.52 | 16.40 | 0.10 |

*Notes*: Growth rate (in percentage, %) of main aggregate variables: output (*y*), private consumption (*c*), government consumption (*g*), investment (*i*), export (*x*), and import (*m*), except for trade balance-to-output ratio (*tby*). ASEAN-5 countries: Indonesia, Malaysia, Philippines, Singapore and Thailand.

vulnerable to international externalities than Vietnam's since the latter has stricter capital flow controls and its economy is *de facto* a non-free market economy.

The rest of the paper is structured as follows. Section 2 is devoted to the stylized facts of Vietnam and the ASEAN-5 countries. Section 3 presents the small open RBC model. Sections 4 and 5 contain the estimation and counterfactual experiments for Vietnam. Section 6 briefly focuses on Thailand, and Section 7 concludes.

## 2. Business cycle analysis

### 2.1. Data and background

Macroeconomic research on Vietnam's economy is challenging due to the limited data availability. For output, the most recent quarterly data just covers the years after 2000, but not all the other main aggregates are provided publicly. We collect annual aggregate data from the United Nations Statistical Division (UNSD) because the time coverage is long enough to identify meaningful business cycles. To economize notation, we define a vector of main aggregate variables $J \equiv (y, c, i, g, x, m, h, tby)$ corresponding, respectively, to output, private consumption, investment, government consumption, exports, imports, and the trade balance-to-output ratio; in turn, *tby* expresses the ratio of net exports over output. Table 1 summarizes all main aggregate growth rates for Vietnam and the average of the ASEAN-5 economies.

Vietnam's real GDP per capita has steadily grown from a low of \$200 US in the 1980s (constant 2005 USD) to six times higher over the past three decades (in nominal terms it reached more than \$2100 US in 2015). This marked the success of the Doi Moi program, an economic renovation strategy that transformed Vietnam from one of the world's poorest countries in the beginning of the 1990s to a middle-income one in less than twenty-five years (World Bank, 2013, 2016).

Table 1 documents this success in comparison to the ASEAN-5 countries as Vietnam's economy experienced higher growth rates in GDP as well as in all demand components with respect to the ASEAN-5 average. Taking as reference the whole period, 1986–2015, Vietnam's real income growth sustained a higher level than the mean of the ASEAN-5 countries, at 4.84 % and 3.74 % respectively. This positive differential was small prior to 2000, when the economy was still suffering from economic sanctions due to the consequences of the Vietnam War. However, the Bilateral Trade Agreement (BTA) between Vietnam and the US signed in 2001 and the subsequent Free Trade Agreements (FTA) with other East Asian countries (*e.g.* Korea, Japan and China) in later years led to a flourishing economy. This explains the much larger differential between 2001 and 2015 (5.29 % and 3.69 %, respectively). In addition, Vietnam's economy was much less volatile in 2001–2015 than in 1986–2000, as the standard deviation of GDP growth went down by a third.

It is also interesting to compare the relative behaviour in the two five-year recession periods, 1997–2001 and 2008–2012, related to the Asian crisis and the global financial crisis (GFC). In the former, Vietnam's economy displayed an impressive growth rate of 5.06 % per annum (0.18 % in the ASEAN-5 economies), while in 2008–2012 it became more vulnerable with a loss of 0.7 percentage points with respect to the average in 2001–2015 (note that this is the highest loss within the ASEAN community). Note that in the post-Asian crisis years, Indonesia was the most stable economy together with Vietnam, while Singapore displayed the highest volatility.

The standard growth accounting framework (Solow, 1957) is often used to decompose output growth into parts due to input factors (*i.e.*, capital and labour in the canonical Cobb-Douglas production function) and the Solow residuals. The latter component, the so-called total productivity factor (TFP), summarizes all information about technological progress and other unexplained elements. Using the classical methodology in Bosworth and Collins (2003) and the Penn World Table (PWT) 9.0 datasets, in Table 2 we

**Table 2**
Growth rates and cyclical volatility of the Solow residuals.
Source: Penn World Table 9.0.

| | Solow residual's growth rate, % | | | | Standard deviation ($\sigma_{sr}$), % | | | |
|---|---|---|---|---|---|---|---|---|
| | 1970 - 2015 | 1986 - 2015 | 1986 - 2000 | 2001 - 2015 | 1970 - 2015 | 1986 - 2015 | 1986 - 2000 | 2001 - 2015 |
| Indonesia | 2.15 (0.53) | 2.80 (0.77) | 2.14 (1.53) | 3.44 (0.19) | 3.55 (1.27) | 3.94 (1.63) | 5.36 (2.11) | 0.97 (0.17) |
| Malaysia | 1.41 (0.55) | 0.99 (0.58) | 1.19 (1.00) | 0.78 (0.60) | 3.62 (0.69) | 3.01 (0.74) | 3.62 (1.15) | 2.19 (0.58) |
| Philippines | 1.32 (0.51) | 0.39 (0.38) | −0.29 (0.61) | 1.07 (0.39) | 3.41 (0.68) | 1.99 (0.21) | 2.29 (0.30) | 1.41 (0.23) |
| Singapore | 1.67 (0.53) | 1.87 (0.74) | 2.53 (0.97) | 1.21 (1.12) | 3.50 (0.42) | 3.89 (0.52) | 3.65 (0.86) | 4.00 (0.72) |
| Thailand | 1.37 (0.45) | 1.77 (0.66) | 1.35 (1.14) | 2.17 (0.68) | 2.96 (0.47) | 3.35 (0.66) | 3.98 (0.95) | 2.40 (0.30) |
| Vietnam | 1.96 (0.38) | 1.90 (0.21) | 1.85 (0.39) | 1.97 (0.16) | 2.54 (0.45) | 1.10 (0.15) | 1.42 (0.18) | 0.59 (0.11) |

*Notes*: Standard errors in parentheses are estimated by GMM estimator. Cyclical standard deviation obtained by regressing TFP over time trend $t$. That is, $sr_t = sr_0 + \lambda t + \varepsilon_t^{sr}$, where $sr_t$ is the estimated Solow residuals from the Cobb-Douglas production function; then $\sigma_{sr} = \sqrt{Var(\varepsilon_t^{sr})}$.

report the TFP growth rate and its volatility for Vietnam and the five ASEAN countries. The variability of the linearly detrended TFP behaves differently both across countries and time periods. Vietnam's TFP varied in a considerably narrower band than its ASEAN-5 peers, as its standard deviations were the smallest, especially in the 2000s when they were very modest (0.6 %). In the context of neoclassical economics, this explains why per capita GDP growth in Vietnam fluctuated much less than that of its peers in the last two decades. Conversely, Singapore appears to be the most fluctuating economy. Finally, note that Indonesia and Thailand experienced larger positive TFP growth rates than Vietnam in 2001–2015 (3.4 % and 2.2 % against 2.0 %). Later on, we briefly examine Thailand's economy, while the unique case of Indonesia has been analysed in Lee and Hong (2012).

### *2.2. Business cycle facts*

The growth accounting exercise so far provides the contribution of supply-side factors to output growth, but it cannot describe the variability of output in the short run nor its co-movements with other aggregate variables. Those fluctuations have been defined as temporary deviations from some secular growth path over time (or the so-called "trend"), which should be a fitted smooth curve (Prescott, 1986). Accordingly, one can separate the transitory part from the trend in several ways. In this chapter we consider four standard econometric techniques, which are first-differencing, linear regression, the Hodrick and Prescott (1997) (HP) filter and the one-sided variant of the HP filter (Stock and Watson, 1999). Note that elsewhere in the text the small case letter variables (*e.g.*, $z_t$) denote natural logarithms of the respective level, which are expressed in capital letter ($Z_t$). Hence, the cyclical component should be interpreted as the percentage deviation from its trend. Business cycle analysis involves computing the standard deviations of the cyclical main aggregate components ($\sigma_{j\in J}$) and their relative ratios with respect to output ($\sigma_j/\sigma_y$). We apply all four filters discussed above to extract cyclical signals from the trends. The outcome is presented in Table 3.

At first glance, the HP filters provide the smoother paths, which are close to the demeaned first-differencing series and contrast with the oscillatory pattern resulting from the linear-quadratic filter. For Vietnam the latter delivers output standard deviations roughly 1.5–2.5 times larger than the HP and first-differencing counterparts (see Fig. 1). According to the results from this estimation, Vietnam's output oscillates around the trend by 4.75 % per year on average so that two business cycles are clearly identified over the Doi Moi period, *i.e.*, the first over 1986–1997 and the second one over 1997-2008. Then, in the aftermath of the GFC, Vietnam's output evolved below the trend. Another relevant outcome is the different behaviour of the ASEAN-5 countries, where 10-year cycles involving the periods 1975–1986, 1987–1998 and 1998–2008 are addressed. Lastly, Vietnam has lower volatility in GDP but higher relative volatility in demand components with respect to the ASEAN-5 countries.

Given the estimated relative moments, Vietnam's economic regularities seem to be consistent with the RBC literature for emerging countries (*e.g.*, Uribe and Schmitt-Grohé, 2017). Investment and foreign trade activities are by far the most volatile components, coinciding with the growth accounting evidence reported before. This could be explained by Vietnam's intense demand of imported

**Table 3**
Business cycle's statistics for Vietnam and the ASEAN-5 countries 1986–2015.

| Filter | First differenced | | Linear quadratic | | Hodrick-Prescott | | One-sided HP | |
|---|---|---|---|---|---|---|---|---|
| StdDev | Vietnam | ASEAN-5 | Vietnam | ASEAN-5 | Vietnam | ASEAN-5 | Vietnam | ASEAN-5 |
| $\sigma_y$ | 1.74 | 3.56 | 4.75 | 6.89 | 2.05 | 4.28 | 1.74 | 4.15 |
| $\sigma_c/\sigma_y$ | 1.38 | 1.04 | 0.75 | 0.93 | 1.22 | 1.00 | 1.06 | 1.04 |
| $\sigma_i/\sigma_y$ | 7.01 | 3.21 | 5.56 | 3.63 | 6.42 | 3.74 | 5.52 | 3.71 |
| $\sigma_g/\sigma_y$ | 2.06 | 1.32 | 1.26 | 1.47 | 2.44 | 1.07 | 1.91 | 1.37 |
| $\sigma_x/\sigma_y$ | 4.88 | 2.31 | 2.84 | 2.05 | 3.52 | 1.37 | 3.05 | 1.97 |
| $\sigma_m/\sigma_y$ | 5.07 | 3.06 | 2.89 | 2.73 | 3.70 | 1.90 | 3.14 | 2.83 |
| $\sigma_{tby}/\sigma_y$ | 2.62 | 1.61 | 0.96 | 0.96 | 2.61 | 1.55 | 2.22 | 1.60 |

*Notes*: All ratios are estimated by GMM estimators. Standard errors are not reported.

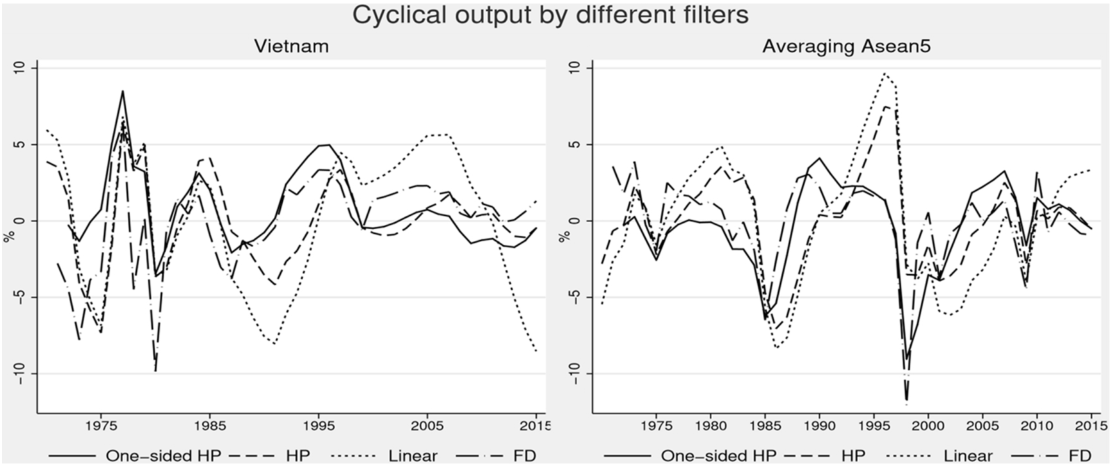



**Fig. 1.** Dynamics of Vietnam and the ASEAN-5 economies.

goods over the past decades because of the needs of high-tech manufacturing equipment, by-products for fabricating and assembling industries, electronic devices, automobiles, and sizable investments in public infrastructure.

Although the specific case of linear filtering would not support this conclusion, the consumption of Vietnamese households seems to fluctuate more than output, in line with the higher variance of consumption to output reported in the literature on emerging economies. One of the great successes of the Doi Moi was the subsequent increase in the living standard of the Vietnamese people as "more than 40 million people escaped poverty over the course of two decades" (World Bank, 2016). This fact, which is connected to the low starting base of household consumption, at least partly explains the higher consumption-to-output standard deviation ratios in Vietnam with respect to the ASEAN-5 economies.

Vietnam's export and import growth rates have sustained a notable pace of 10 % per year since 2001, twice that of the ASEAN-5 countries, leading the degree of trade openness to expand from 1.13 at the beginning of the 2000s to a high of 2.4 by the end of 2015. However, Vietnam's trade balance was negative over the whole thirty-year period, revealing weak competitiveness, over-consumption of imported goods, and vulnerability of the economy to adverse shocks, especially during the GFC in 2008–2012. Besides, the autocorrelation function of the trade balance-to-output ratio presents a monotonically downward trend approaching 0.18 beyond the fourth order (see Table 4). This phenomenon is commonly observed in emerging markets, but not all studies have succeeded in matching or reproducing it.

Looking at the results from the one-sided HP filter perspective, Vietnam's exports and imports seem to be experiencing procyclical movements, as their first lagged and contemporaneous correlations with output are at moderate levels, 0.57 and 0.65 respectively. These figures are close to the ASEAN-5's average as displayed in Table 4. Note that all autocorrelation coefficients almost die out after two years, yet they have three times more volatility than output. This is indicative of the fact that both Vietnam's and the ASEAN-5's

**Table 4**
Cyclical correlations.

| Correlation with $y_c$ | | Lag | | | | Lead | | | Autocorrelation | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Country | Variables | −3 | −2 | −1 | 0 | 1 | 2 | 3 | −1 | −2 | −3 | −4 |
| Vietnam | $y_c$ | 0.36 | 0.65 | 0.89 | 1.00 | 0.89 | 0.65 | 0.36 | 0.89 | 0.65 | 0.36 | 0.08 |
| | $c_c$ | −0.39 | −0.13 | 0.20 | 0.48 | 0.58 | 0.56 | 0.45 | 0.75 | 0.42 | 0.13 | −0.06 |
| | $ic$ | 0.64 | 0.77 | 0.74 | 0.63 | 0.40 | 0.12 | −0.20 | 0.68 | 0.56 | 0.30 | −0.01 |
| | $g_c$ | 0.24 | 0.34 | 0.36 | 0.27 | 0.06 | −0.28 | −0.54 | 0.78 | 0.47 | 0.16 | −0.14 |
| | $x_c$ | 0.35 | 0.46 | 0.57 | 0.57 | 0.48 | 0.44 | 0.26 | 0.49 | 0.37 | 0.20 | 0.27 |
| | $m_c$ | 0.32 | 0.46 | 0.60 | 0.65 | 0.57 | 0.53 | 0.35 | 0.48 | 0.33 | 0.14 | 0.16 |
| | $tby$(*) | 0.47 | 0.45 | 0.38 | 0.32 | 0.27 | 0.19 | 0.12 | 0.75 | 0.49 | 0.30 | 0.18 |
| ASEAN-5 | $y_c$ | 0.04 | 0.26 | 0.64 | 1.00 | 0.64 | 0.26 | 0.04 | 0.64 | 0.26 | 0.04 | −0.12 |
| | $c_c$ | −0.03 | 0.13 | 0.41 | 0.70 | 0.45 | 0.12 | −0.05 | 0.66 | 0.23 | −0.04 | −0.21 |
| | $ic$ | 0.03 | 0.22 | 0.54 | 0.88 | 0.64 | 0.28 | 0.01 | 0.69 | 0.30 | 0.04 | −0.12 |
| | $g_c$ | −0.19 | −0.11 | 0.10 | 0.28 | 0.17 | 0.09 | 0.00 | 0.70 | 0.41 | 0.14 | −0.09 |
| | $x_c$ | 0.17 | 0.36 | 0.52 | 0.66 | 0.38 | 0.07 | −0.01 | 0.59 | 0.31 | 0.15 | 0.01 |
| | $m_c$ | 0.08 | 0.32 | 0.55 | 0.76 | 0.42 | 0.05 | −0.10 | 0.62 | 0.27 | 0.04 | −0.11 |
| | $tby$(**) | 0.21 | 0.11 | −0.03 | −0.28 | −0.31 | −0.27 | −0.19 | 0.74 | 0.50 | 0.33 | 0.23 |

*Notes*: Correlation with contemporaneous $y_c$ and $y_c$, $c_c$, $i_c$, $g_c$, $x_c$ and $m_c$. One-sided HP filter with $\lambda = 100$. Period 1986–2015. Note: (*) Standard HP filter, linear-quadratic and first-differencing filters report $tby$ being acyclical with slightly negative correlation with output. (**) There are three (Malaysia, Philippines, Thailand) out of five countries showing negative contemporaneous correlation between $tby$ and $y_c$.

foreign trade was fairly sensitive to international economic conditions.

Household consumption and investment present the expected procyclical behaviour. Their contemporaneous correlations with output are in the range of [0.5, 0.6] and investment interestingly exhibits a "time-to-build" effect as its first and second-order correlations are, respectively, as high as 0.74 and 0.77. Correspondingly, lead relationships between investment and output are poor, and the second-order lead seems to be uncorrelated. On the contrary, the ASEAN-5's private consumption and investment are procyclical and moderately persistent, as their contemporaneous correlations and first-order auto-correlations are roughly 0.7.

In contrast, both Vietnam and the ASEAN-5's public demand for goods and services behave acyclically and display some persistence, as indicated by their first-order autocorrelations and contemporaneous correlation with output, which are, respectively, above 0.70 and below 0.28 (but positive). The same reading applies to the trade balance-to-output ratio variable.

To conclude, the above set of business cycle facts show that the characteristics of Vietnam's economy are similar to those of the ASEAN-5 countries and reconcilable with the findings of the RBC literature on small open emerging economies (Uribe and Schmitt-Grohé, 2017).

## 3. Model

The natural question arising from the previous analysis refers to the kind of technological shocks that are most relevant in driving the economic fluctuations of small open emerging economies like Vietnam and its ASEAN-5 peers. To respond to this question, we develop a small open economy DSGE-RBC model aiming to quantify the exogenous forces that shape the dynamics of Vietnamese growth aggregates (and, in a final comparative exercise, Thailand's).

### 3.1. Setting the economic environment

The model is an extension of GPU's financial friction specification, which adds internal habit persistence (Boldrin et al., 2001) and government consumption to the period utility function (Christiano and Eichenbaum, 1992). In what follows, we opt for the end-of-period notation since it is naturally compatible with Dynare's coding convention.

We assume a symmetric, single good economy endowed with a constant returns-to-scale Cobb-Douglas production technology. The production function is defined as

$$Y_t = A_t K_t^{\alpha} (X_t h_t)^{1-\alpha} \tag{1}$$

where $X_t$ represents labour-augmenting technological change, which has a cumulative effect, as noted in AG. Thus, $X_t = g_t X_{t-1}$; where $g_t$ is the productivity's gross rate of growth so that $\log\left(\frac{g_t}{h_t}\right) = \rho_g \log\left(\frac{g_{t-1}}{\mu_g}\right) + \varepsilon_t^g$, with $\mu_g$ being the gross long-run growth rate and $|\rho_g| < 1$. The single trend shock $\varepsilon_t^g$ is assumed to be normally distributed with variance $\sigma_g^2$, $\varepsilon_t^g \sim N(0, \sigma_g^2)$.

Variable $a_t \equiv \log(A_t)$ denotes the transitory productivity process following the usual AR(1) propagation mechanism such that $a_t = \rho_a a_{t-1} + \varepsilon_t^a$, with $|\rho_a| < 1$ and $\varepsilon_t^a \sim N(0, \sigma_a^2)$.

Capital stock accumulation is subject to the following law of motion:

$$K_t = (1-\delta)K_{t-1} + I_t - \frac{\phi}{2}\left(\frac{K_t}{K_{t-1}} - \mu_g\right)^2 K_{t-1} \tag{2}$$

where $\delta$ represents the rate of depreciation and $\phi$ is the parameter to be estimated. Note that the last term on the right-hand side governs the capital adjustment costs.

The instantaneous utility function takes the Greenwood et al. (1988) (hereafter GHH) form as in GPU because it is well-known that GHH preferences generate the excess volatility of consumption over output and counter-cyclical net exports (see Correia et al., 1995). Thus,

$$U(C_t^*, h_t) = \frac{[C_t^* - \theta\omega^{-1}X_{t-1}h_t^{\omega}]^{1-\eta} - 1}{1-\eta} \tag{3}$$

where $C_t^* = \nu_t C_{p,t} - \gamma C_{p,t} - 1 + \pi C_{g,t}$, with $C_p$ and $C_g$ being private and public consumption, respectively; $\eta > 0$ and $\eta \neq 1$; and $\nu_t$ denotes an exogenous and stochastic preference shock. The existence of $C_{p,t-1}$ and $\gamma > 0$ in the utility function accounts for internal habit persistence, implying that household consumption has a time-non-separable structure as past decisions affect the present utility. Two parameters, $\theta > 0$ and $\omega > 0$, control the disutility of work and need to be calibrated to the normalized long-run hours-work of the target country. The involvement of government consumption in the instantaneous utility (3) is determined by $\pi \in [0, 1]$, which represents the elasticity of substitution between private and public consumption. We model $C_{g,t} = \zeta_{cg}\xi_{cg,t}Y_t$, with the spending shock $\xi_{cg,t}$ and a constant ratio of $\zeta_{cg}$.

Specifying the utility function as in (3) has several non-trivial implications. First, habit formation generally improves the explanatory power of the DSGE models by allowing consumption smoothness (see *e.g.*, Fuhrer, 2000; Christiano et al., 2005; Ravn et al., 2006, to name a few). Second, the present internal setting for habit persistence has been advocated by Constantinides (1990) and Boldrin et al. (2001), as their studies suggest that "internal habit" is better than the "keeping-up-with-the-Joneses" counterpart in jointly explaining the risk premium puzzle and business fluctuations that small open economies likely encounter. Lastly, GHH preferences give rise to the complementarity between consumption and hours. In our case, by incorporating government spending into the period utility, we allow for partial ($0 < \pi < 1$) or complete ($\pi = 1$) complementarity between private and public spending. Note that this is in contrast to GPU's financial friction model, which implicitly imposes $\pi = 0$.

The two AR(1) processes, $\nu_t$ and $\xi_{cg,t}$, perturb the present household and government consumption as follows

$$\log(\nu_t) = \rho_\nu \log(\nu_{t-1}) + \varepsilon_{\nu,t} \log(\xi_{cg,t}) = \rho_{cg} \log(\xi_{cg,t-1}) + \varepsilon_{cg,t}$$

where $\varepsilon_{\nu,t} \sim N(0, \sigma_\nu^2)$ and $\xi_{cg,t} \sim N(0, \sigma_{cg}^2)$.

The household budget is

$$B_t(1 + r_t)^{-1} = C_{p,t} + C_{g,t} + I_t + B_{t-1} \tag{4}$$

where $B_t$ and $r_t$ are the stock of debt and domestic interest rate at time $t$, respectively. The trade balance-to-output ratio in the model is defined as $TB_t/Y_t = (B_{t-1} - B_t/(1+r_t)) / Y_t$, so that the negative value of $B_t$ represents the economy-wide indebtedness (*i.e.*, investment over saving). The market clearing condition is written as

$$(Y_t - C_{p,t} - C_{g,t}) - I_t = TB_t \tag{5}$$

Regarding the bond discount rate $r_t$, Rhee (2017) considers an endogenous risk premium channel – proposed by Neumeyer and Perri (2005) – as an alternative to GPU's configuration. The latter approach postulates a domestic interest rate that is the sum of the world interest rate $r^* > 0$ (assumed to be constant), the country's risk premium, and an exogenous shock to the country's premium $\mu_t$. Since we are using GPU's model as a benchmark, the rule for the exogenous risk premium channel is restated as

$$r_t = r^* + \psi(e^{\frac{B_t}{X_{t-1}} - \bar{b}} - 1) + e^{\mu - 1} - 1 \tag{6}$$

where $log(\mu_t) = \rho\mu log(\mu_{t-1}) + \varepsilon\mu_t$, and $-1 < \rho\mu < 1$ with $\varepsilon_{\mu,t} \sim N(0, \sigma_\mu^2)$ is a parameter governing debt elasticity, and $\bar{b}$ is the steady-state level of governmental outstanding debt.

Our model is richer than GPU's financial friction setup in the number of exogenously stochastic processes as it allows distinguishing shocks to the pure-time discount rate ($\xi_{p,t}$) from shocks to current consumption ($\nu_t$). The former is also known as a preference-shifter since it lets parameter $\beta$ vary across time, while the latter is the same as in GPU. The representative household thus seeks to maximize (7) subject to constraints (2) and (4)

$$E_0 \sum_{t=0}^{\infty} \xi_{p,t} \beta^t U(C_t, h_t) \tag{7}$$

As a consequence, the Lagrangian function is

$$- \Lambda_t \left[ \frac{B_t}{1 + r_t} + a_t K_{t-1}^\alpha (X_t h_t)^{1-\alpha} - C_{p,t} + C_{g,t} + K_t - (1 - \delta) K_{t-1} + \frac{\phi}{2}\left(\frac{K_t}{K_{t-1}} - \mu_g\right)^2 K_{t-1} + B_t \right]\Bigg\} \tag{8}$$

with Lagrangian multiplier $\Lambda_t = \lambda_t X_{t-1}^{-\eta}$.

Along with Eqs. (1), (2), (4) and (6), and the six AR(1) exogenous stochastic shocks, the first-order conditions of (8) give us the set of equilibrium conditions (9)–(12) for:

Consumption,

$$\xi_{p,t}\left[\frac{\nu_t}{(\nu_t C_{p,t} - \gamma C_{p,t-1} + \pi C_{g,t} - X_{h,t-1}^*)} - \frac{\lambda_t}{X_{t-1}^\eta}\right] = \frac{\gamma \beta \xi_{p,t+1}}{(\nu_{t+1} C_{p,t+1} - \gamma C_{p,t} + \pi C_{g,t+1} - X_{h,t}^*)^\eta} \tag{9}$$

with $X_{h,t-1}^* \equiv X_{t-1} \theta \omega^{-1} h_t^\omega$.

Hours-work,

$$\frac{\lambda_t (1 - \alpha) a_t (X_t h_t)^{1-\alpha} K_{t-1}^\alpha}{X_{t-1}^\eta h_t} = \frac{\theta X_{t-1} \nu_t h_t^{\omega - 1}}{(\nu_t C_{p,t} - \gamma C_{p,t} + \pi C_{g,t} - X_{h,t-1}^*)^\eta} \tag{10}$$

Capital,

$$\frac{\lambda_t \xi_{p,t}}{X_{t-1}^\eta}\left[\left(\frac{K_t}{K_{t-1}} - \mu_g\right)\phi + 1\right] = \frac{\lambda_{t+1} \xi_{p,t+1}}{X_t^\eta} \beta \left[\alpha a_{t+1}\left(\frac{X_{t+1} h_{t+1}}{K_t}\right)^{1-\alpha} - \frac{\phi}{2}\left(\frac{K_{t+1}}{K_t} - \mu_g\right)^2 + \frac{\phi K_{t+1}}{K_t}\left(\frac{K_{t+1}}{K_t} - \mu_g\right) + 1 - \delta\right] \tag{11}$$

Domestic bond,

$$1 + r_t = \frac{\lambda_t}{\lambda_{t+1}} \frac{\xi_{p,t}}{\xi_{p,t=1}} \frac{g_t^\eta}{\beta} \tag{12}$$

To see the difference between internal and external consumption persistence, replace $C_{p,t-1} \equiv \bar{C}_{p,t-1}$ (the aggregate private consumption at time $t - 1$) in the period utility function. As a consequence, the optimizing household decides its current consumption taking the previous aggregate private spending as given, so that (9) reduces to

$$\frac{\lambda_t}{X_t^{\eta}} = \frac{\nu_t}{(\nu_t C_{p,t} - \gamma C_{p,t} + \pi C_{g,t} - X_{h,t-1}^{*})^{\eta}} \tag{13}$$

It is obvious that the preference shifter $\xi_p$ does not perturb Lagrangian multiplier $\lambda_t$ in (13). This implies that an external habit specification is inefficient in our setting.[1]

The model above is non-stationary and will not converge to the balance-growth path because the output $Y_t$ increases over time by factor $X_t$ in Eq. (1). It is thus necessary to detrend all equilibrium conditions, which we do following AG.[2]

### 3.2. The long-run equilibrium

We characterize the long-term relationships among (detrended) national income identity variables by ruling out the presence of all exogenous shocks. Given that the domestic interest rate and the world interest rate are identical in equilibrium, one has

$$1 + r^{*} = \frac{\mu_g^{\eta}}{\beta} = 1 + r \tag{14}$$

and

$$\frac{h}{k} = \frac{1}{\mu_g^2}\left[\frac{\frac{\mu_g^{\eta}}{\beta} + \delta - 1}{\alpha}\right]^{1/(1-\alpha)} = \frac{1}{\mu_g^2}\left[\frac{r+\delta}{\alpha}\right]^{1/(1-\alpha)} \tag{15}$$

Eq. (15) expresses the labor-to-capital relationship. Since $0 < \alpha < 1$, the ratio $h/k$ would increase with $r$, ceteris paribus. There are three ways to lift up the domestic interest rate $r$ in equilibrium (all else being equal). First, by calibrating under a higher value of $\eta$; second, under a higher long-run labor-augmenting growth rate $\mu_g$; third, under a lower subjective discount factor $\beta$.

The feature that distinguishes our approach from AG-type models is the presence of habit formation ($\gamma$), which changes the way hours-work $h_t$ behaves

$$h = \left[\frac{\theta}{\mu_g(1-\alpha)\left(1 - \frac{\gamma\beta}{\mu_g^{\eta}}\right)}\left(\frac{r+\delta}{\alpha}\right)^{\alpha/(1-\alpha)}\right]^{1/(1-\omega)} \tag{16}$$

We calibrate $\omega$ to a value greater than unity, which is the standard in the RBC literature. Subsequently, an increase in the internal habit formation coefficient, $\gamma$, decreases the steady-state labor-supply to a lower value, underpinning the household resistance to unanticipated changes. Likewise, the lower value of the subjective discount factor $\beta$ will decrease the hours-work, since the household may become impatient. Accordingly, the ratios $(k/y)$, $(i/y)$ and $(c_p/y)$, which are functions of $(h/k)$ as defined in Eqs. (17)–(19), are directly affected by the deep parameters implied in Eqs. (15) and (16),

$$\frac{k}{y} = \mu_g^{1-2\alpha}\left(\frac{h}{k}\right)^{1-\alpha} \tag{17}$$

$$\frac{i}{y} = \frac{k}{y}\left(1 - \frac{1-\delta}{\mu_g}\right) \tag{18}$$

$$\frac{c_p}{y} = 1 - \zeta_{cg} - \zeta_b\left(\frac{1}{\mu_g} - \frac{1}{1-r}\right) - \frac{k}{y}\left(1 - \frac{1-\delta}{\mu_g}\right) \tag{19}$$

$$\frac{x-m}{y} = \frac{tb}{y} = \zeta_b\left(\frac{1}{\mu_g} - \frac{1}{1+r}\right) \tag{20}$$

where $\zeta_{cg}$ and $\zeta_b$ are, respectively, the shares of government consumption and the stock of bonds in the steady state.

The model is the collection of equilibrium Eqs. (9)–(12), six AR(1) processes, and Eqs. (1), (2), (4) and (5). It is then solved by the second-order solution algorithm of Schmitt-Grohé and Uribe (2004) before proceeding with the Bayesian estimation.

[1] The identification test (Iskrev, 2010) reports pairwise multi-collinearity for $\xi_{p,t}$ and $\nu_t$ if $\gamma$ is external.

[2] We follow the standard convention according to which a stationary variable – represented by a lowercase letter – is equivalent to the ratio of that variable (dividend) with respect to $X_{t-1}$ (divisor). Following the end-of-period convention, the variable determined at time $t$ will be adjusted by factor $X$ at time $t-1$, and so on.

**Table 5**
Calibrated parameters.

| Calibrated params | $\beta^*$ | $\pi$ | $\theta$ | $\omega$ | $\eta$ | $\alpha$ | $\delta$ | $\mu_g$ | $\zeta_{cg}$ | $\zeta_b$ |
|---|---|---|---|---|---|---|---|---|---|---|
| Value | 0.03-0.07 | 1 | 1.85 | 2.00 | 2.00 | 0.35 | 0.10 | 1.048 | 0.07 | −0.36 |

## 4. Estimation and discussions

We estimate the model using the standard Bayesian Markov Chain Monte Carlo (MCMC) estimator which – after Smets and Wouters (2003) – is used in most recent DSGE literature and is available in Dynare and similar packages. The Bayesian estimation lends itself to the Bayes rule for the conditional distribution of a set of estimating parameters $\theta \in \Theta$ given observational data $Y$. The core formula is

$$\Gamma(\theta Y) = \frac{f(Y\theta)\Gamma(\theta)}{f(Y)} \tag{21}$$

where $\Gamma(\theta|Y)$ is the so-called posterior probability distribution function of the parameter $\theta$ (or posterior distribution shortly) conditional on observational data $Y$. The prior distribution $\Gamma(\theta)$ is the unconditional probability distribution of $\theta$, whilst the likelihood function $f(Y|\theta)$ is defined as in classical econometric methods. The last component, $f(Y)$, is the marginal likelihood defined as $f(Y) = \int f(Y\theta)\Gamma(\theta)d\theta$.

Methodologically, Bayesian estimation requires three key elements: advanced knowledge of prior specifications, a suitable filter for the likelihood evaluation at every observational data point, and the Metropolis-Hastings Monte Carlo (MHMC) simulation algorithm (see also An and Schorfheide, 2007; Fernández-Villaverde, 2010; and Herbst and Schorfheide, 2015, for in-depth technical expositions).

### 4.1. Configuration

A time unit is meant to represent a year in the model. Econometrically, not all "deep" model parameters are estimated, as some of them will be calibrated to the commonly used values within the RBC literature (see *e.g.*, Cooley and Prescott, 1995; King and Rebelo, 1999; and Schmitt-Grohé and Uribe, 2003, among many others). For instance, the depreciation rate and labor-supply elasticity are micro-based parameters which are not intrinsically supported by the main aggregate information.

Table 5 reports the choices of a number of calibrated parameters based on GPU and Vietnam long-run data. First, we set the value of $\omega$ to 2.0, which results in $\theta = 1.85$, so that the normalized value of hours-work is approximately one-fourth of a unit time-endowment. For consistency with earlier growth accounting evidence, we set the capital income share to $\alpha = 0.35$. The depreciation rate, $\delta$, and the curvature of the period utility function, $\eta$, take the common values of 10 % and 2, respectively. Next, the ratios of government consumption, $\zeta_{cg}$, and stock of domestic bonds to output, $\zeta_b$, are calibrated to 0.07 and −0.36, respectively. The former value is simply the sample average, but the latter is determined from the long-run trade balance-to-output ratio of −2.6 % in years before 2000.

We neither calibrate the subjective discount factor $\beta$ nor the long-term growth rate $\mu_g$, as Rhee (2017) and GPU do in their studies. Given the absence of evidence on Vietnam's business cycles, they need to be estimated. To check for robustness and verify the result's sensitivity to these estimates, we consider a range of values such that $\mu_g \in [0.03, 0.05]$ and $\beta^* \in [0.03, 0.07]$. The presence of government consumption in the period utility causes government consumption to reduce the total household utility whenever $0 \leq \pi < 1$. In turn, fixing $\pi = 1$ entails full complementarity between private and public consumption.[3]

There is a total of sixteen structural parameters to be estimated, comprising six AR(1) coefficients, their corresponding exogenous stochastic disturbance variances, and $\beta$, $\psi$, $\gamma$, $\phi$. GPU suggested the sole use of uniform distributions to estimate the structural parameter space. Rhee (2017), however, estimates the AG-type model using a mixture of *Beta*, *Gamma* and $Inverse-Gamma$ distribution priors. As shown in Table 6 (first four columns), we adopt Rhee's (2017) prior specifications.

It is worth emphasizing that an advantage of directly modeling non-stationary variables in AG- and GPU-type models is that it allows for model-based detrending instead of an arbitrary selection amongst abundant filtering techniques used in data transformations for estimating deep parameters. Canova (2014) stresses the consistency between model and data when both permanent and transitory shocks coexist. Along the same lines, Canova and Ferroni (2011) and Ferroni (2011) show that structural estimates could be biased or distorted due to the wrong choice of time series filter or trend misspecification.

Our model has more shocks (six) than observed data (five), namely four growth rate aggregates and the trade balance-to-output series. The direct links among data and model variables, *i.e.*, the measurement equations, are defined as

[3] Even though it passes Iskrev's (2010) test, the identification strength of $\pi$ is weak (with a hardly distinguishable posterior from the prior distribution). Figs. A4 and A5 provide a sensitivity analysis to study how the variability of parameter $\pi$ affects the modelled variables. The habit persistence coefficient appears to be the strongest parameter, while the elasticity of public spending $\pi$ only has considerable effects on the responses of $g_y$ and $g_{cp}$ with respect to two shocks $\varepsilon_{cg}$ and $\varepsilon_p$. Hence, contrary to what could be expected, the presence of public spending in the utility function does not have a relevant impact. This is in line with the conclusions in Cantore et al. (2014).

**Table 6**
Estimated parameters for Vietnam (1981–2015).

| Estimated params | Prior | | | Posterior | | | |
|---|---|---|---|---|---|---|---|
| | Dist. | Mean | SD | Mean | 5% | 95% | p-val |
| $\mu_g$ | *Beta* | 1.05 | 0.02 | 1.045 | 1.039 | 1.052 | 0.981 |
| $\beta^* = 1/\beta$ - 1 | *Beta* | 0.04 | 0.02 | 0.064 | 0.030 | 0.098 | 0.507 |
| $\psi$ | *Gamma* | 0.50 | 0.50 | 0.322 | 0.000 | 0.629 | 0.794 |
| $\gamma$ | *Beta* | 0.50 | 0.20 | 0.268 | 0.093 | 0.435 | 0.932 |
| $\phi$ | *Gamma* | 5.00 | 2.00 | 3.844 | 1.907 | 5.844 | 0.793 |
| $\rho_a$ | *Beta* | 0.50 | 0.20 | 0.720 | 0.513 | 0.933 | 0.584 |
| $\rho_g$ | *Beta* | 0.50 | 0.20 | 0.608 | 0.372 | 0.851 | 0.751 |
| $\rho_p$ | *Beta* | 0.50 | 0.20 | 0.579 | 0.296 | 0.882 | 0.110 |
| $\rho_{cg}$ | *Beta* | 0.50 | 0.20 | 0.774 | 0.619 | 0.934 | 0.892 |
| $\rho_\mu$ | *Beta* | 0.50 | 0.20 | 0.781 | 0.598 | 0.953 | 0.711 |
| $\rho_\nu$ | *Beta* | 0.50 | 0.20 | 0.739 | 0.521 | 0.934 | 0.504 |
| $\sigma_a$ | *IGamma* | 0.01 | 0.20 | 0.007 | 0.003 | 0.001 | 0.088 |
| $\sigma_g$ | *IGamma* | 0.01 | 0.20 | 0.005 | 0.003 | 0.008 | 0.159 |
| $\sigma_p$ | *IGamma* | 0.05 | 0.20 | 0.043 | 0.012 | 0.080 | 0.127 |
| $\sigma_{cg}$ | *IGamma* | 0.05 | 0.20 | 0.027 | 0.016 | 0.037 | 0.978 |
| $\sigma_\mu$ | *IGamma* | 0.05 | 0.20 | 0.026 | 0.014 | 0.037 | 0.405 |
| $\sigma_\nu$ | *IGamma* | 0.05 | 0.20 | 0.039 | 0.014 | 0.064 | 0.021 |
| $\sigma_{gy}$ | *IGamma* | 0.01 | 0.20 | 0.009 | 0.004 | 0.014 | 0.437 |
| $\sigma_{gcp}$ | *IGamma* | 0.01 | 0.20 | 0.018 | 0.013 | 0.023 | 0.931 |
| $\sigma_{gcg}$ | *IGamma* | 0.01 | 0.20 | 0.012 | 0.003 | 0.025 | 0.874 |
| $\sigma_{ginv}$ | *IGamma* | 0.10 | 0.50 | 0.094 | 0.067 | 0.120 | 0.398 |
| $\sigma_{tby}$ | *IGamma* | 0.05 | 0.20 | 0.016 | 0.011 | 0.022 | 0.530 |

*Notes*: Beta denotes Beta distribution; (I) Gamma represents (inversed) gamma distribution. p-val is p-value (15 %-taper) of Geweke (1992) Convergence test (see Appendix A for additional trace plots). The estimation is based on 1,000,000 draws from the MHMC algorithm.

$$g_{y,t} = \Delta lnY_t = \log(y_t) - \ln(y_{t-1}) + g_{t-1} - 1$$
$$g_{cp,t} = \Delta lnC_{p,t} = \log(c_{p,t}) - \log(c_{p,t-1}) + g_{t-1} - 1$$
$$g_{cg,t} = \Delta lnC_{g,t} = \log(c_{g,t}) - \log(c_{g,t-1}) + g_{t-1} - 1$$
$$g_{inv,t} = \Delta lnI_t = \log(i_t) - \log(i_{t-1}) + g_{t-1} - 1$$
$$tby_t = TB_t/Y_t = (y_t - c_{p,t} - c_{g,t} - t_t)/y_t$$

where the vector of observational data ($g_y$, $g_{c\ p}$, $g_{cg}$, $g_{inv}$) is the per capita annual growth rate of (output, private consumption, government consumption, investment) and tby represents the trade balance-to-output ratio. These linkages arise naturally from the model, implying the model-based differencing data transformation. Following GPU, a measurement error is added to each observed variable to resolve filtering errors, data quality or even occasional model misspecification (Del Negro and Schorfheide, 2009). Besides, although the sample size could be taken to be a matter (we exercise Bayesian estimation over the sample period of 1981–2015), it has been shown that relatively small sample sizes can produce valid Bayesian inference within the DSGE context (Fernández-Villaverde and Rubio-Ramírez, 2004; Benchimol and Fourçans, 2017)

## 4.2. Results

### 4.2.1. Estimated parameters

All estimated parameters are reported in columns 5–9 of Table 6. The results indicate that the subjective discount rate is fairly moderate at approximately 6.4% per annum (equivalent to 1.6 % per quarter), implying that the stationary value of the real interest rate $r = \mu_g^\eta/\beta$ would be in the range of [12.2%, 16.2%], corresponding to $\eta$ = [1.2, 2.0] and $\mu_g$ = 1.045. As a sensitivity check, we run Bayesian estimations at each $\eta$ = (1.25, 1.50, 1.75, 2.00), which still deliver a consistent interval of $\beta^* \in$ [6.3%, 6.9 %]. The low value of $\beta$ for Vietnam (0.94), compared to that obtained for the US by King and Rebelo (1999), $\beta$ = 0.98, indicates that a Vietnamese household is qualitatively more impatient than one living in an advanced country. This is not unreasonable in view of the severe difficulties regarding high inflation experienced by Vietnam in the 1980s and 90 s, and even in the aftermath of the GFC, regarding high inflation. Unlike most small open emerging economies, however, Vietnam is on its way to transforming from a closed and centralized economy to a pro-business and pro-market one.

The estimated value of $\psi$ is 0.29 (much lower than GPU's estimate for Argentinian data), reflecting that the Vietnamese domestic interest rate is quite insensitive to the debt level. In turn, capital adjustment costs $\phi$ ~ 4.0 are remarkably close to those found in Latin America by GPU, suggesting that fixed capital formation in developing countries is, in general, a costly process to sustain growth targets.

Regarding the AR(1) processes, the six estimated coefficients $\rho_{(.)}$ are within the plausible range of [0.58, 0.78], in accordance with the use of annual data. The fluctuations of both stationary and non-stationary technological shocks are fairly small, 0.68 % and 0.55

**Table 7**
Vietnam's growth variable moments from Bayesian estimation, 1981–2015.

| | | | Model | | GPU | |
|---|---|---|---|---|---|---|
| Moments | Actual Data | | W. habit | W/o. habit | Basic | FFR |
| $\sigma_{gy}$ | 1.64 | (0.21) | 1.69 | 2.23 | 5.68 | 5.67 |
| $\sigma_{gcp}/\sigma_{gy}$ | 1.35 | (0.17) | 1.21 | 1.84 | 0.91 | 2.21 |
| $\sigma_{gcg}/\sigma_{gy}$ | 2.03 | (0.46) | 2.24 | 2.58 | – | – |
| $\sigma_{ginv}/\sigma_{gy}$ | 6.84 | (1.44) | 5.31 | 6.24 | 1.29 | 3.24 |
| $\sigma_{tby}/\sigma_{gy}$ | 2.61 | (0.59 | 2.82 | 3.91 | 3.22 | 5.70 |
| $corr(tby, g_y)$ | −0.12 | (0.17) | 0.04 | −0.13 | −0.01 | 0.01 |
| $corr(g_{cp}, g_y)$ | 0.61 | (0.14) | 0.07 | 0.15 | 0.90 | 0.30 |
| $corr(g_{cg}, g_y)$ | 0.36 | (0.16) | 0.45 | 0.39 | – | – |
| $corr(g_{inv}, g_y)$ | 0.35 | (0.16) | −0.20 | −0.18 | 0.77 | 0.09 |
| $AR1(tby)$ | 0.78 | (0.17) | 0.76 | 0.72 | 0.92 | 0.82 |
| $AR1(g_y)$ | 0.89 | (0.19) | 0.02 | 0.12 | 0.07 | −0.09 |
| $AR1(g_{cp})$ | 0.47 | (0.13) | 0.19 | −0.08 | 0.07 | −0.12 |
| $AR1(g_{cg})$ | 0.36 | (0.16) | −0.05 | −0.08 | – | – |
| $AR1(g_{inv})$ | −0.11 | (0.28) | −0.13 | −0.16 | 0.04 | −0.17 |

*Notes*: Standard errors in parentheses. AR1 denotes first-order autocorrelation. With (W.) habit is model with internal habit persistence, while W/o. is the one without habit coefficient.

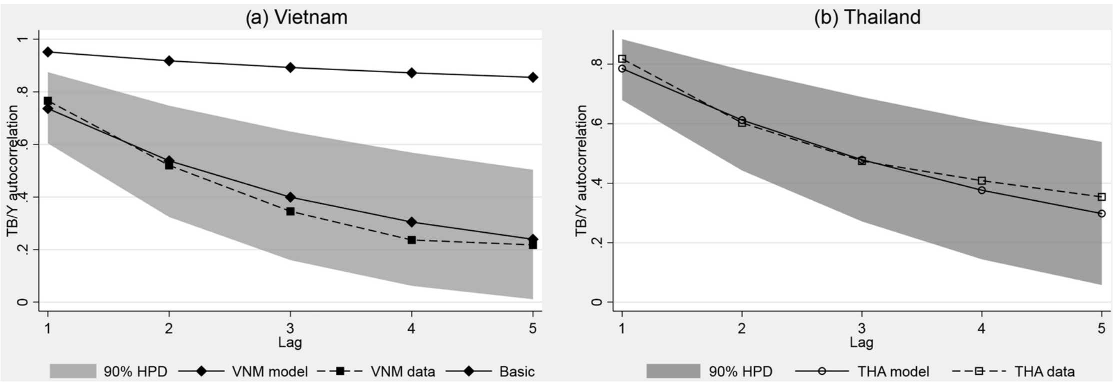



**Fig. 2.** Trade balance-to-output ratio (*tby*) autocorrelation of Vietnam (left panel) and Thailand (right panel). VNM (THA) model refers to our proposed RBC model; VNM basic is the standard RBC prototype in GPU.

% respectively, whereas the variations of the temporal preference and consumption taste are comparatively large, 4.4 % and 3.9 % respectively. These estimates are able to account for the excess volatility of Vietnamese household consumption compared to output.

### 4.2.2. Simulated results

The top block of Table 7 reports the striking performance of the model (column W. habit), as it can effectively reproduce output variability and other important relative standard deviations. In the absence of internal habit formation (W/o. habit), the predicted variance of output is higher than the actual one by 25 %; the ratios ($\sigma_{gcp}/\sigma_{gy}$, $\sigma_{ginv}/\sigma_{gy}$, $\sigma_{ginv}/\sigma_{gy}$, $\sigma_{tby}/\sigma_{gy}$) are higher than those of the full model, and neither the downward slope of the trade balance-to-output autocorrelation function (Fig. 2) nor the excess variation of household consumption to output can be reproduced in GPU's basic setting. When GPU's financial friction (FFR) is added, the output growth variance is overpredicted, while $\sigma_{ginv}/\sigma_{gy}$ is underestimated. Besides, the basic model has a tendency of generating strong procyclical growth rates of consumption and investment.

Failure to model consumption memory would worsen the model's moment matching and the short-run dynamic behaviour. The estimated habit persistence γ is close to the mean value reported in the meta-analysis of Havranek et al. (2017, Table 1), 0.27 and 0.30, respectively. This indicates that the strength of habit persistence of Vietnamese households is as strong as those living in advanced countries.[4] This strength underlies the gradual responses of Vietnamese private consumption and inflation to all policy shocks (Fuhrer, 2000).

We need to acknowledge that neither the current model nor GPU's can wholly recreate the short-run dynamics of Vietnam's growth observables, as displayed in the two bottom blocks of Table 7. This can be explained by the relatively small sample timeframe that we used in comparison to the 100-year datasets of Mexico and Argentina in GPU's study. On the other hand, Table 8

[4] Havranek et al. (2017) explore 81 studies covering Australia, New Zealand, the G7 group and many EU countries.

**Table 8**
Simulated (auto-) correlations of level variables.

| Moment | Actual | W. habit | Moment | Actual | W. habit |
|---|---|---|---|---|---|
| $corr(y, y)$ | 1.00 | 1.00 | $AR1(y)$ | 0.77 | 0.86 |
| $corr(y, c_p)$ | 0.69 | 0.54 | $AR1(c_p)$ | 0.61 | 0.85 |
| $corr(y, c_g)$ | 0.53 | 0.65 | $AR1(c_g)$ | 0.73 | 0.78 |
| $corr(y, i)$ | 0.55 | 0.41 | $AR1(i)$ | 0.52 | 0.72 |

*Notes*: Actual expresses HP filtered data ($\lambda = 100$). *corr* and *AR1* are contemporaneous and 1st-order autocorrelation, respectively.

**Table 9**
Unconditional variance decomposition (in percent, %).

| Shock | Model | Output growth | Cons. growth | Gov. spending growth | Investment growth | Trade balance to GDP ratio |
|---|---|---|---|---|---|---|
| Nonstationary tech | W. habit | 18.67 | 7.90 | 8.65 | 8.53 | 5.84 |
| | W/o habit | 26.60 | 7.00 | 3.97 | 11.77 | 9.10 |
| | GPU-FFR | [3.81] | [1.20] | | [1.35] | [1.15] |
| Stationary tech | | 49.72 | 10.64 | 22.69 | 2.46 | 1.62 |
| | | 51.25 | 7.12 | 7.49 | 1.03 | 1.70 |
| | | [88.71] | [10.18] | | [3.63] | [5.16] |
| Preference shift | | 0.07 | 4.88 | 0.03 | 0.16 | 0.63 |
| | | 0.13 | 12.39 | 0.02 | 0.21 | 2.94 |
| Cons. preference | | 8.77 | 60.93 | 4.23 | 7.66 | 19.85 |
| | | 0.65 | 30.56 | 0.10 | 0.82 | 12.14 |
| | | [0.85] | [83.08] | | [2.09] | [67.86] |
| Gov. spending | | 0.00 | 0.58 | 53.39 | 0.00 | 0.01 |
| | | 0.00 | 1.50 | 85.00 | 0.00 | 0.00 |
| | | [0.00] | [0.01] | | [0.00] | [0.02] |
| Country premium | | 22.78 | 15.07 | 11.00 | 81.19 | 72.05 |
| | | 21.38 | 41.43 | 3.43 | 86.16 | 74.12 |
| | | [6.62] | [5.54] | | [92.92] | [25.81] |

*Notes*: The estimates are based on 800,000 draws from the posterior distribution. GPU-FFR estimates – GPU's financial friction model estimated using the mix of Beta and Gamma priors – are in squared brackets.

demonstrates that the model is sound in capturing the contemporaneous and first-order autocorrelations of the four aggregate levels, namely $y$, $c_p$, $c_g$ and $i$.

Finally, we highlight the failure of GPU-type models in replicating Vietnam's cyclical moments due to (i) the time-invariant subjective discount factor $\beta$, which implies that optimizing agents do not adjust their forward-looking expectations over time; (ii) the absence of a mechanism that permits inter-temporal consumption smoothing; and crucially (iii) the co-existence of habits in consumption and capital adjustment costs, which markedly reduces volatility in investments and output (Khorunzhina, 2015). Our estimation is eloquent in reporting increasing $\sigma_{gy}$ and other relative ratios when optimizing agents have no consumption memory.

### *4.3. Variance decomposition*

Using Kalman's filter, we compute the long-term (unconditional) variance decomposition of the variables due to orthogonal shocks and compare them across models (Table 9). The results indicate that the contribution of a transitory TFP shock to the output growth variance amounts to 50 % if we use our extended model (89 % when using GPU's). Accordingly, it seems safe to conclude that Vietnam's business cycles in 1981–2015 were not driven by trend innovations. This is in clear contrast to the results of AG for their set of 13 emerging market economies.

Shocks to trend productivity and the country risk premium each account for about 20 % of the remaining output variability. This finding uncovers a novel relevant impact of risk premium shocks on economic growth. Moreover, the results suggest that the long-term fluctuations of investments and the trade balance-to-output ratio are mainly driven by the country risk premium. Surprisingly, GPU's financial friction contradicts this key finding by predicting a critical impact of the consumption preference shock on the trade balance. To illustrate the significant impact of risk premium $\varepsilon\mu$ on the observable *tby*, we decompose its variance over time, as displayed in Fig. 3. This figure shows that most of the variation in the trade balance-to-output ratio in 1999–2015 was due to the exogenous interest rate disturbances $\varepsilon\mu$.

### *4.4. The importance of the trend shock*

AG use the Beveridge-Nelson (BN) decomposition to derive the relative importance of the trend component with respect to the transitory component of the productivity shock in shaping the dynamics of Solow's residual. The variance ratio reads

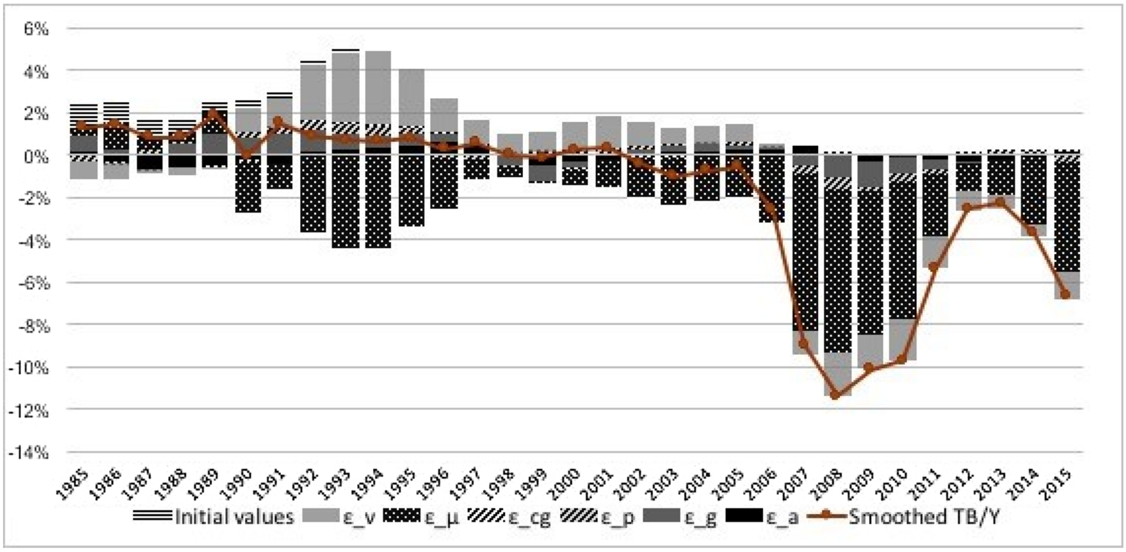



**Fig. 3.** Historical variance decomposition of *tby* (Vietnam).

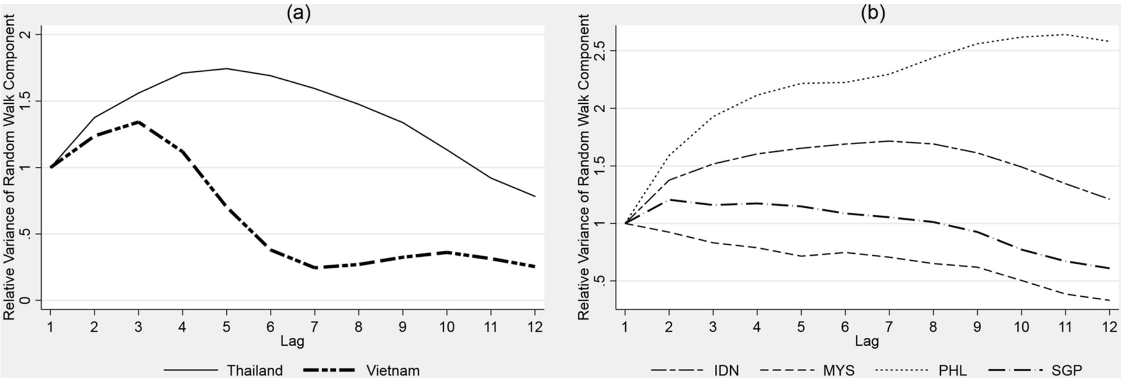



**Fig. 4.** Relative variance of the Solow residual's components using Cochrane's (1988) formula.[9] Panel (4a): Vietnam and Thailand; Panel (4b): Indonesia, Malaysia, the Philippines and Singapore.

$$\frac{\sigma^2_{\Delta\tau}}{\sigma^2_{\Delta sr}} = \frac{\alpha^2\sigma^2_g}{(1-\rho_g)^2\sigma^2_{\Delta sr}}$$

where $sr_t$ denotes the conventional Solow residuals and $\tau_t$ is the trend part of the BN decomposition, such that $sr_t = \tau_t + s_t$. In Fig. 4, we report all relative variance ratios up to lag 12.[5] It appears that, with the exclusion of the Philippines, the trend does not seem to be as important as AG predicted when the lag-length increases infinitely. The above decomposition, however, is not fully convenient since it does not directly contain $\sigma_a$ – the estimated standard deviation of a transitory shock. Since $a_t$ and $g_t$ are AR(1) processes, their variances would be $\sigma_a^2/(1-\rho_a^2)$ and $\sigma_g^2/(1-\rho_g^2)$, respectively. Recall that $\sigma_a$ and $\sigma_g$ are the standard deviations of the corresponding $\varepsilon_{a,t}$ and $\varepsilon_{g,t}$, respectively. Thus, the importance of the trend shock can be computed as[6] :

$$\frac{Var(g_t)}{Var(\Delta sr_t)} = \frac{\sigma_g^2}{(1-\alpha^2)(1+\rho_a)\sigma_g^2 + 2(1-\rho_g^2)\sigma_a^2}$$

with $\sigma_a = 0.68$ %, $\sigma_g = 0.54$ %, $\rho_a = 0.72$ and $\rho_g = 0.61$. Hence, the relative variance of the trend process $g_t$ is 28.5 %. As $g_t$ appears in the Cobb-Douglas function with factor $(1-\alpha)$, the model predicts that the non-stationary component only explains $(1–0.35)^2 \times 28.5\ \% \approx 12\ \%$ of output movements. This implies that output growth in Vietnam's economy has been quite stable in response to the non-stationary component of productivity shocks driving Solow's residual. This estimation is in accordance with the empirical relevance as the ratios displayed in Fig. 4 show.

[5] We find negligible differences between Cochrane (1988) and the AG-modified formula, but the former is easier to compute.
[6] The fact is that $Var[(1-\alpha)g_t] = (1-\alpha^2)\sigma_g^2/(1-\rho_g^2)$ and $Var(a_t - a_{t-1}) = 2Var(a_t) - 2Cov(a_t,\ a_{t-1})$.

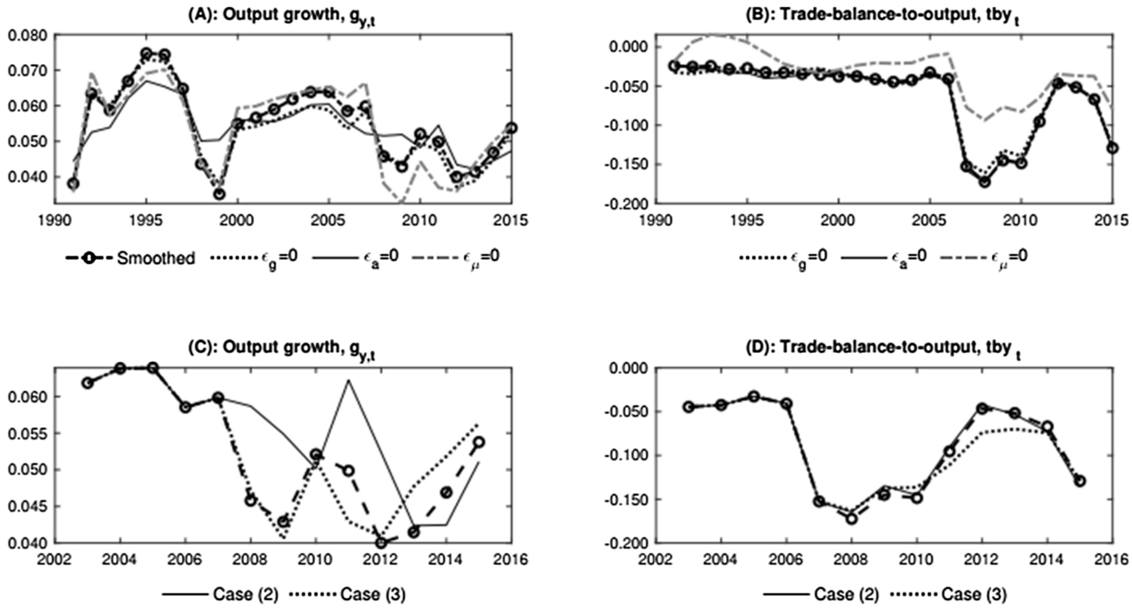



**Fig. 5.** Counterfactual simulation.

Notes: Counterfactual simulation. First row, Panel A and B express case (1) of setting by turn $\varepsilon_a = 0$ (solid line), $\varepsilon_g = 0$ (dotted line), and $\varepsilon\mu = 0$ (gray dashed line). Second row, Panel C and D correspond to, respectively, the scenarios of setting $\varepsilon_a 2008_{,g} - 2013 = \varepsilon_a{}^{1992}{}_{,g}{}^{-199}$[10] (Case 2: solid line), and $\varepsilon\mu_{,t} = 0$ for period 2008–2013 (Case 3: dotted line). Smoothed series (dashed circle line) depict actual data. Note that the dashed circle line displays actual data in all panels.

## 5. Counterfactual simulations

In this section, we investigate the trajectories of several observables under different loaded shocks. There are some interesting policy implications that arise from analysing fictitious scenarios and their consequences for Vietnam's economy. In Scenario 1, transitory productivity shocks, trend shocks and monetary policy shocks are sequentially muted, *i.e.* $\varepsilon_{l\,\in\,\{a,g,\mu\}} = 0$. In Scenario 2, we apply the TFP shocks experienced by Vietnam in 1992–1997 to the years 2008–2013. In Scenario 3, shocks to the country risk premium become silent ($\varepsilon\mu_{,t} = 0$) in 2008–2013.

The first row of Fig. 5 reports the simulated paths of output growth $g_{y,t}$ (Panel A) and trade balance-to-output $tby_t$ (Panel B) in the first scenario when we set $\varepsilon_{a,t} = 0$ (solid line), $\varepsilon_{g,t} = 0$ (dotted line), and $\varepsilon\mu_{,t} = 0$ (gray dashed line) in every data point in the sample. The smoothed line (dashed circle line) is the actual data reconstructed using Kalman's filter. Panels C and D supply analogous information for Scenarios 2 and 3.

The counterfactuals in Panels A and B show that in the absence of transitory productivity shocks the trajectory of trade balance would have remained unchanged, while that of income growth would have become more stable. This implies that these kinds of shocks help stabilize the business cycle but have little effect on the external sector. This is in contrast to the impact of trend shocks $\varepsilon_{g,t}$, which have little impact on both income growth and trade balance, as the gap between the actual and simulated lines is small and even non-existent in many data points. With respect to the monetary policy shocks, note that in Panel A the deviation of the gray dashed line becomes more significant in the 2000s, reflecting real effects of the proactive Vietnamese monetary policy in the twenty-first century, as also found in Anwar and Nguyen (2018).[7] Panel B, however, shows that the monetary policy itself was responsible for the huge trade deficits in the period 2007–2010 despite the fact that Vietnam stabilized over the GFC (note that in Panel A the positive income growth gap between actual data and the gray dashed line is about 0.7 % in 2008–2013). All in all, the analysis in Scenario 1 contradicts, for the case of Vietnam, AG's claim regarding the leading role of the trend shock in explaining a sudden drop in the trade deficit along with large contractions in connected aggregates.

Panel C shows that if Vietnam's economy had experienced the same technological improvements as in 1992–1997 – as tested in Scenario 2 –, the simulated path of $g_{y,t}$ (solid line) would have evolved above the actual path during the GFC and subsequently. The average gain per year, over 2008–2013, amounts to 0.8 percentage points. On the contrary, in the absence of risk premium shocks, the net effect on income growth appears to be small in the same period. These shocks, however, exert a significant influence on the trade deficit, as shown by the substantially smaller deficits in 2011–2013 in the absence of such shocks (dotted line in Panel D). The main conclusion we draw from these exercises goes back to the first result. In terms of high growth and a sustainable pace, it seems that technological progress is the critical condition to achieve the best possible path for the Vietnamese economy.

Finally, let us note that the counterfactual exercises to some degree strengthen the findings in Huynh et al. (2017), according to which the monetary policy exerted through interest rates management was insufficient to face the economic downturn. Given our analysis and their evidence based on a calibrated DSGE model, it seems that a strategy more focused on fostering loan supply and targeting productivity generating sectors would have been more successful. We hope this lesson may be helpful for the future design

**Table 10**
Thailand and Vietnam 1976–2015.

| Period | Thailand | | Vietnam | |
|---|---|---|---|---|
| | rGDP | rGDPpc | rGDP | rGDPpc |
| 1970 – 1986, Pre-boom (*) | 6.08 | 3.82 | 4.67 | 2.39 |
| 1987 – 1996, Boom (**) | 9.00 | 7.78 | 6.54 | 4.53 |
| 1997 – 1998, Asian-crisis | −5.38 | −6.51 | 6.72 | 5.37 |
| 1999 – 2007, Post-crisis | 5.07 | 4.19 | 6.67 | 5.67 |
| 2008 – 2009, Global-crisis | 0.48 | 0.34 | 5.38 | 4.41 |
| 2010 – 2015, Post-crisis | 3.55 | 3.20 | 5.82 | 4.72 |

*Notes*: (*) and (**) show Thailand's pre-boom and boom periods as documented in Warr (2005, p.5). rGDP (rGDPpc) is the rate of growth of real GDP (per capita).

**Table 11**
Bayesian estimation for Thailand 1976–2015.

| Params | $\sigma_{gy}$ | $\sigma_{gcp}/\sigma_{gy}$ | $\sigma_{gcg}/\sigma_{gy}$ | $\sigma_{ginv}/\sigma_{gy}$ | $\sigma_{tby}/\sigma_{gy}$ |
|---|---|---|---|---|---|
| Model | 3.55 | 1.39 | 1.78 | 6.95 | 4.60 |
| Data | 3.62 | 1.12 | 1.07 | 3.71 | 2.15 |
| | | (0.088) | (0.238) | (0.419) | (0.366) |
| $\sigma_a$ | $\sigma_g$ | $\sigma_p$ | $\sigma_{cg}$ | $\sigma_\mu$ | $\sigma_\nu$ |
| 0.017 | 0.005 | 0.031 | 0.037 | 0.025 | 0.026 |
| $\rho_a$ | $\rho_g$ | $\rho_p$ | $\rho_{cg}$ | $\rho_\mu$ | $\rho_\nu$ |
| 0.855 | 0.512 | 0.545 | 0.831 | 0.864 | 0.680 |
| | $\psi$ | $\gamma$ | $\phi$ | $\beta^*$ | $\mu_g$ |
| | 0.095 | 0.128 | 3.766 | 0.032 | 0.038 |

*Notes*: Standard error in parentheses; $\sigma$ denotes the standard deviation. The $g_y$, $g_{cp}$, $g_{cg}$, $g_{inv}$ and *tby* variables are defined as in Section 5.

**Table 12**
Long-run variance decomposition for Thailand (in percent, %) 1976–2015.

| Variable | Output growth | Consumption growth | Gov. spending growth | Investment growth | Trade balance-to-GDP ratio |
|---|---|---|---|---|---|
| Nonstationary tech | 17.59 | 1.97 | 4.84 | 5.24 | 2.55 |
| Stationary tech | 26.74 | 4.86 | 7.25 | 0.69 | 0.78 |
| Preference shift | 0.05 | 5.80 | 0.01 | 0.05 | 0.51 |
| Cons. preference | 1.01 | 21.83 | 0.28 | 0.41 | 3.48 |
| Gov. spending | 0.00 | 4.82 | 71.31 | 0.00 | 0.01 |
| Country premium | 54.62 | 60.71 | 16.30 | 93.60 | 92.67 |

*Note*: The estimates are based on 800,000 draws from the posterior distribution.

of economic policy in Vietnam.

## 6. Application to Thailand

We extend our investigation by applying the analysis to Thailand, which has been, and still is, Vietnam's key business rival within the ASEAN-5 countries. Before the Asian financial crisis, Thailand was widely recognized as a reference case of an oil-importing emerging economy. For almost 40 years (1958–1996), Thailand sustained positive GDP growth rates and achieved "a combination of rapid growth, macroeconomic stability, and steadily declining poverty incidence" (Warr, 2005, p. 4). As explained by the ADB (2015), the success of Thailand was supported by "political stability, a business-friendly regulatory environment, a large domestic market, open access to foreign investment, and greater participation in regional value chains".

Having faced a debt crisis in 1983–1985, Thailand experienced a major boom-bust cycle in the second half of the 1980s and 90 s, as illustrated in Table 10. The years 1987–1996 were characterized by prosperity, with an average of 9 % GDP growth per annum that ceased when the region was hit by the 1997 crisis. Thailand was again affected by a series of political and financial shocks,[7] but its growth rate kept a pace of 3.97 % per year and it was classified as an upper-middle-income country in 2011 (World Bank, 2011). As Table 10 shows, in spite of this success, Vietnam's economy has outperformed Thailand's by approximately two percentage points since 2000.

Except for $\omega$ and $\theta$, the same calibration set is used to adapt the model to Thailand's data. Since Thailand is calibrated in AG, we

[7] These were a *coup d'état* in 2006 (political shock), the global 2008 crisis (financial shock) and the flood in 2011 (economic shock).

reuse $\omega = 1.6$, so that the elasticity of labor supply, $1/(\omega - 1)$, is set at 1.7 and $\theta = 1.4$. Table 11 demonstrates that the excess volatility of the growth variables with respect to output is essentially matched. The predicted parameter $\sigma_{gy}$ is close to the actual data, and the downward slope of *tby*'s autocorrelation function has been exactly reproduced, as plotted in Fig. 2. Nevertheless, the relative standard deviations are overly predicted because the standard model could not properly handle the two structural breaks (or debt crises) experienced by Thailand's economy in 1982–1985 and 1997-1999. As in the case of Vietnam, several growth variables' contemporaneous and first-order autocorrelation coefficients are understated by the model.[8]

The bottom block of Table 11 shows that the long-term growth rate of Thailand is lower than that of Vietnam by 0.8 percentage points, but the subjective discount rate of Thailand is just one-half of the Vietnamese one. The latter implies that Thailand's economy is better structured even though the capital adjustment costs, $\varphi$, are similar. The lower temporal discount rate reflects the stronger attachment of Thailand's households to their lifetime income. The low estimate of the habit persistence coefficient, $\gamma = 0.13$, supports this claim. In addition, the estimated parameter $\psi$, which controls for the sensitivity of the country risk premium, is one-third of that of Vietnam. This may help to understand Thailand's long-lasting trade balance surplus in the aftermath of the 1997 crisis.

The long-term variance decomposition Table 12 proves that the explanatory power of both the transitory and trend shocks is relatively weak since neither of them can account for more than 27 % of the output growth variance. In particular, the relevance of the non-stationary component is as low as 6.18 %. On the other hand, the country risk premium is vital to explaining Thailand's economic developments, as it accounts for 54.6 % of the unconditional variance over the sample range.

## 7. Concluding remarks

This paper performed a detailed analysis of Vietnam's economy in connection to its ASEAN-5 peers. In the first part, we provided information on two complementary sides of Vietnam's performance in terms of growth sources and business cycle drivers. In the second part, we developed a DSGE-RBC model. Our model departs from those in AG and GPU, further incorporating habit formation and government consumption in utility. In this way it provides a close match of the facts in which our analysis of Vietnam, and the comparison with Thailand, can be conducted.

We show that the contribution of TFP to Vietnam's economic growth is approximately one-third on average in 1986–2015, although it drops to less than 20 % in the 2000s. In turn, capital accumulation has driven Vietnam's economy since 1992, in parallel to the acceleration in the opening and deregulation processes that started the same year. One of the main consequences of such processes was the development of a structural trade deficit.

In terms of business cycle characteristics, we document similar behavior in Vietnam to the ASEAN-5 economies, which does not differ significantly from that in other emerging market economies. The intrinsic difficulties of RBC models in replicating short-term observational dynamics are well-known. With this caveat in mind, it is important to note that the proposed model provides a better account of the facts in Vietnam and Thailand than other reference models in the emerging markets literature – for example, the ones by AG and GPU, which provide the departure point of our modelling strategy.

Provided with this improved setting, the variance decomposition analysis reveals that transitory productivity shocks account for around 50 % of Vietnam's output growth fluctuations. Country-risk premium shocks are also relevant, although they are far more critical in the case of Thailand, as they totally dominate the impact of the productivity shock throughout the whole period (1976–2015). Given these findings, we conclude that Thailand's economy is more vulnerable to international externalities than Vietnam's, which has stricter capital flow controls and is still *de facto* a non-free market economy. It is probably on account of Vietnam's incomplete transition that technological progress and productivity-enhancing measures are shown to be fundamental to securing a sustainable high growth path. All other examined growth drivers seem to be of secondary order.

Another crucial finding is the scarce contribution of non-stationary TFP shocks to Solow's residual volatility at 12 % in Vietnam and 6 % in Thailand. These results, which refute AG's claim that "the cycle is the trend" in emerging market economies, line up with recent literature with similar results for Argentina, Mexico and Korea.


## Acknowledgments

We are grateful for the valuable comments and suggestions made by two anonymous referees. Pham Thai Binh gratefully acknowledges the financial support from the Universitat Autònoma de Barcelona; Hector Sala is grateful to the Spanish Ministry of Economy and Competitiveness for financial support through grant ECO2016-75623-R; José I. Silva is grateful to the Spanish Ministry of Economy and Competitiveness for financial support through grant ECO2017-82350R and Generalitat de Catalunya2017SGR0558.


## Appendix A

See Figs. A1–A3.

[8] Not reported in the main text since the estimates are similar to those for Vietnam.

[9] $\frac{\sigma^2_{\Delta\tau}}{\sigma^2_{\Delta sr}} = \frac{\alpha^2\sigma^2_g}{(1-\rho^2_g)}\sigma^2_{\Delta sr}$, where $\lim_{K\to\infty} K^{-1}Var(sr_t - sr_{t-K} = \sigma^2_{\Delta\tau}$ . We find negligible differences between Cochrane (1988) and the AG-modified formula, but the former is easier to compute.

[10] In contrast to most western and emerging countries, Vietnam did not have a truly independent central bank for decades (Anwar and Nguyen, 2018). Besides, the State Bank of Vietnam regulated both deposit and lending rates until the beginning of the 2000s (Camen et al., 2006).

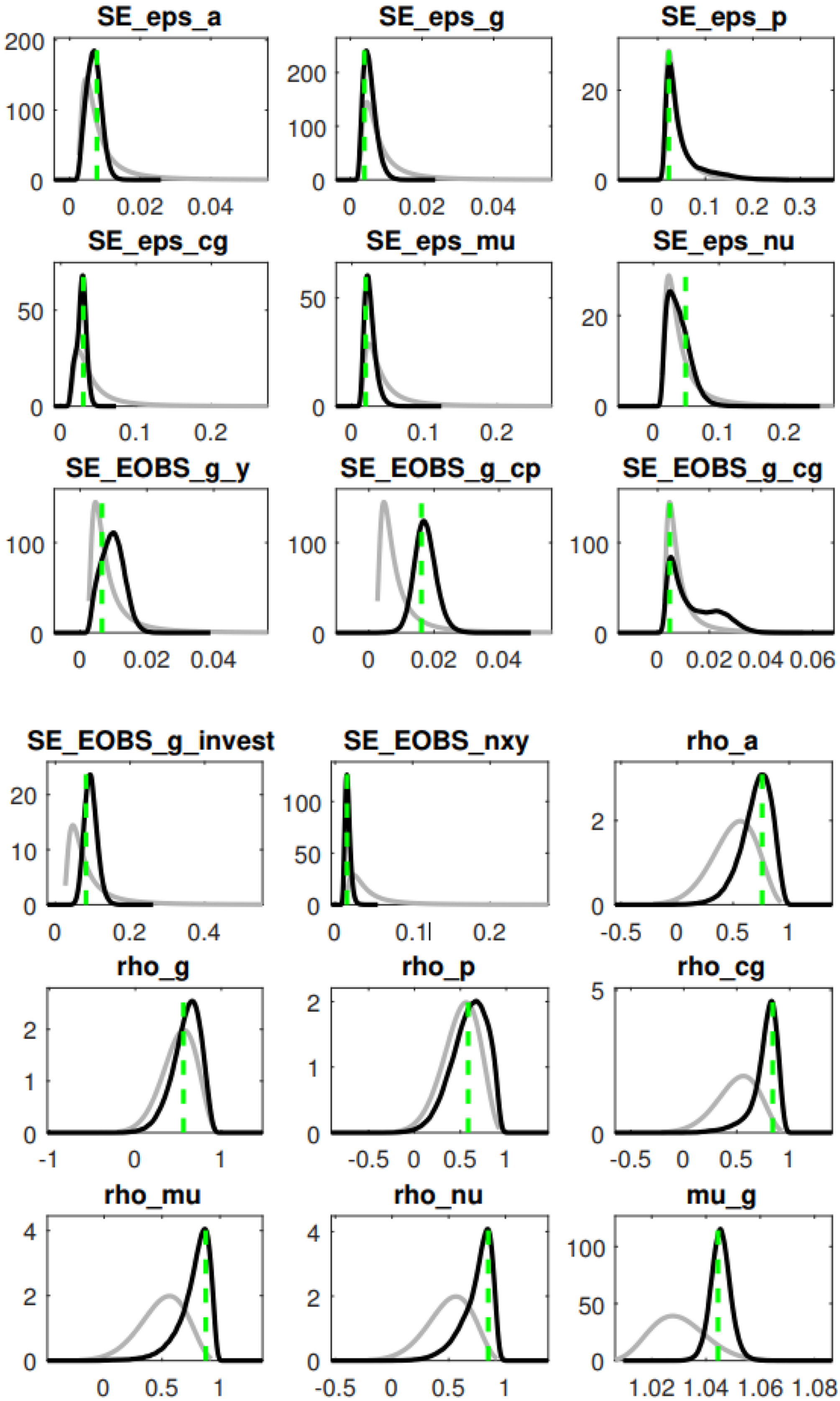



**Fig. A1.** Priors and posteriors (in case of estimated $\pi$ with Beta prior).

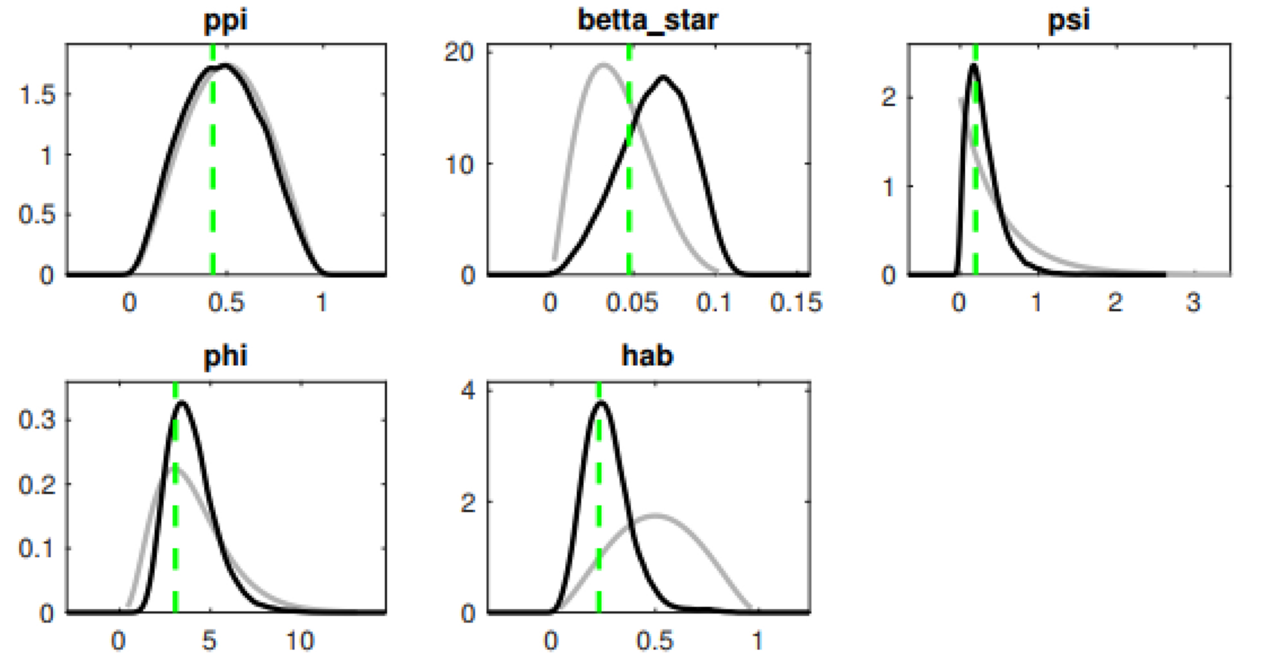



**Fig. A2.** Priors and posteriors (in case of estimated $\pi$ with Beta (0, 1) prior).

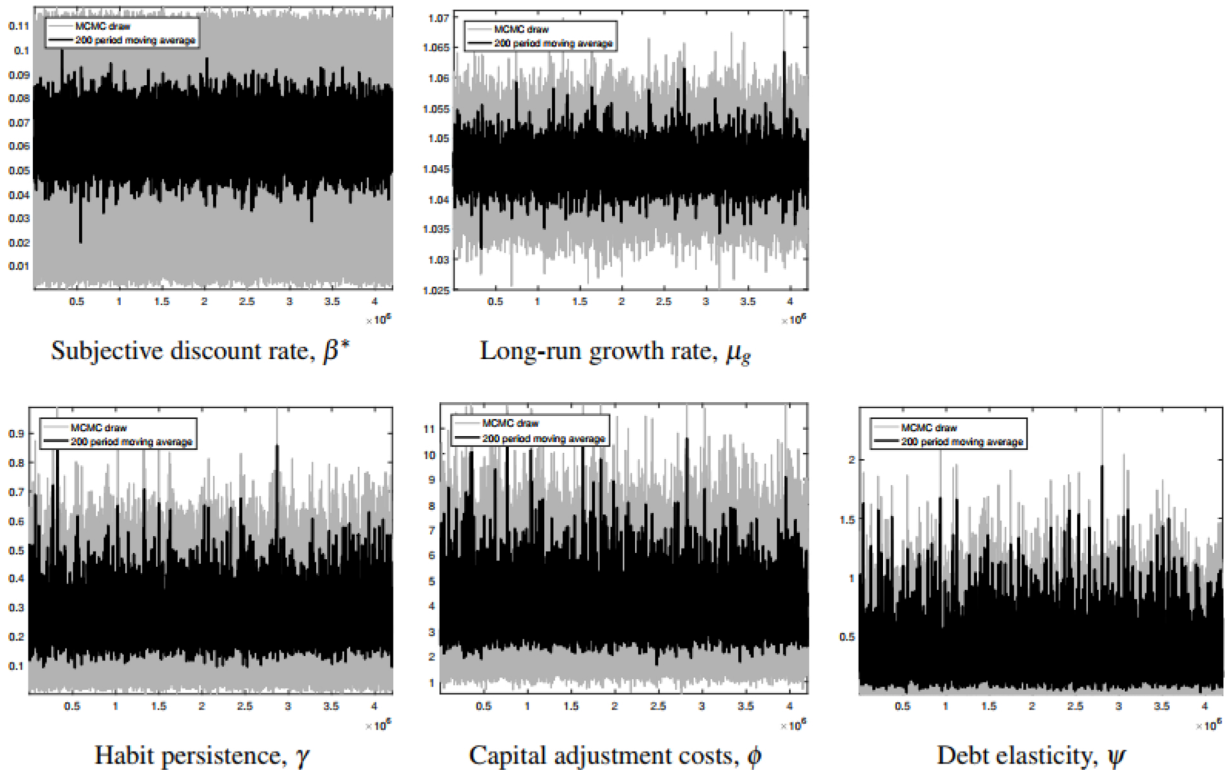



**Fig. A3.** Selected trace plots.

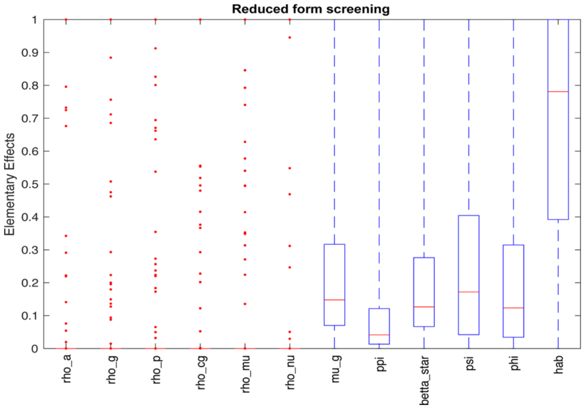
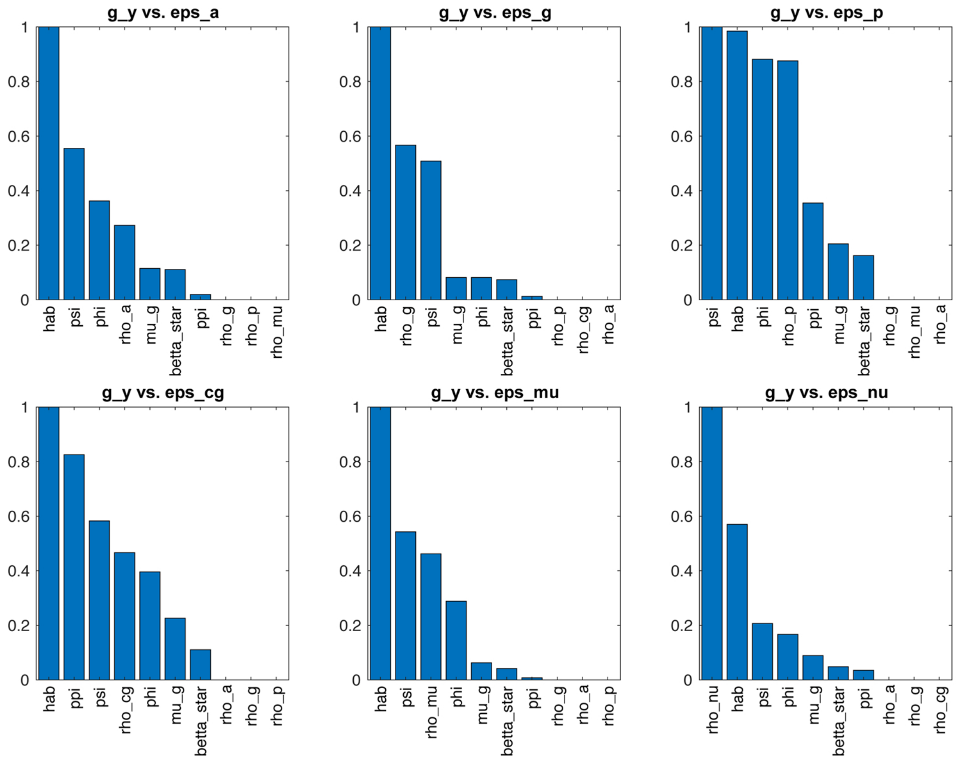
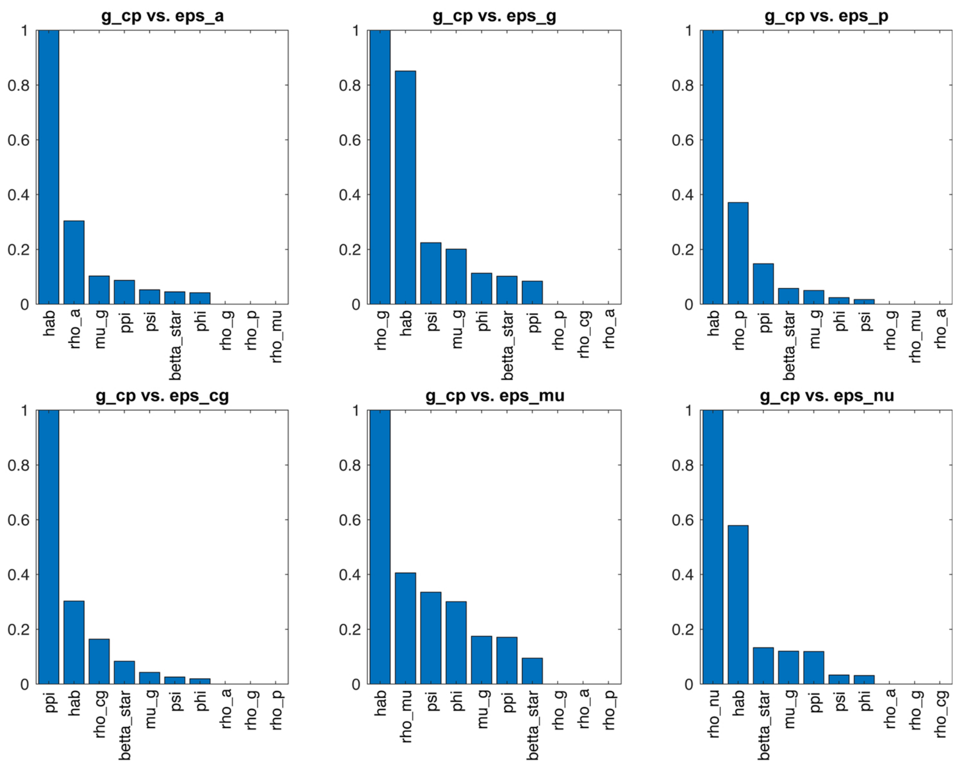



**Fig. A4.** Elementary effects analysis (in case of estimated $\pi$ with Beta (0, 1) prior).




**1**




**2**

**Fig. A5.** 1) Sensitivity analysis w.r.t $g_y$ (in the case of estimated $\pi$). 2) Sensitivity analysis w.r.t $g_{cp}$ (in the case of estimated $\pi$).

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