Universitat Autònoma de Barcelona

MACROECONOMIC ESSAYS ON VIETNAM ECONOMY

PHAM THAI BINH

MACROECONOMIC ESSAYS ON VIETNAM ECONOMY

by

PHAM THAI BINH

A Dissertation Submitted in Partial Fulfilment of the Requirements for the Degree of Doctor of Philosophy

at the

Department of Applied Economics

Faculty of Economics and Business Studies

Universitat Autònoma de Barcelona

December 2019

DECLARATION

This dissertation is the result of my own work and includes nothing, which is the outcome of work done in collaboration except where specifically indicated in the text. It has not been previously submitted, in part or whole, to any university of institution for any degree, diploma, or other qualification.

Signed:____

Date:____

Pham Thai Binh

Barcelona

(This declaration should also make clear any parts of the work that have been submitted for any other qualification.)

Supervisor	Co-Supervisor
(signed)	(signed)
Prof. Héctor Sala Lorda	Prof. José Ignacio Silva

ACKNOWLEDGEMENTS

I am deeply grateful to

Prof. Héctor Sala Lorda

&

Prof. José Ignacio Silva

for their guidance towards the end of the Doctoral degree.

I gratefully acknowledge the financial support from Department of Applied Economics – Universitat Autonòma de Barcelona. I also appreciate for support and advice from

Prof. Anna Matas

Prof. Emilio Padilla

Prof. Javier Asensio

Prof. Rosella Nicolini

Imma Ruano Moyano

María Pilar Bernabé Conde

Mireia Buj Gelonch

and my dear friends.

Dedicated to

My family

CONTENTS

CHAPTER 1: Introduction

1.1 Forewords

This research project is a compendium of four essays that have required me to study the main theoretical and applied tools used today in the broadly defined academic field of macroeconomics. Within the neoclassical framework, I have learnt to work with dynamic stochastic general equilibrium models (DSGE), which first evolved to explain the main assumptions and equations characterising the behaviour of economic agents: individuals/consumers/households, firms, and the public sector. I researched how to manage these models, how to replicate the results of some reference articles and learnt to conduct similar analyses following my own judgement on what required further investigation. Closer to the empirical side of the analysis, I discovered how to perform simulations and conduct counterfactual exercises, which I believe are a very useful outcome of the DSGE framework of analysis. Regarding the use of quantitative methods, I have become familiar with the estimation of Vector Auto Regression (VAR) Models, and Vector Error Correction Models (VECM). I am also now fluent in the setup of *state-of-the-art* Structural VAR (SVAR) models, their estimation using Bayesian methods, and the reading and interpretation of economic results through the resulting impulse-response-functions (IRFs) and variance decomposition.

This steep learning process, in terms of economic tools and methods, has allowed me to conduct a comprehensive analysis of Vietnam. The reason for choosing Vietnam as the main subject for my analysis (I have made frequent comparisons with other economies) was not only due to my familiarity with my home country, but also as it is a largely unexplored economy, which has experienced major structural transformations since the reunification in 1975. The reason why the existing literature has not conducted many studies on Vietnam is due to data availability. While it is true that the World Bank and the International Monetary Fund provide data for Vietnam, this data often lacks information on several key variables, is supplied in a short time-series, or is provided in annual records with some discontinuities between years. Internal major institutional bodies in Vietnam provide more rich data but accessibility to this data is difficult for foreign researchers. Hence, an important part of the empirical work conducted in this thesis has been to assemble coherent databases that are suitable for use in the research planned for each essay.

Regarding the structural transformation, Vietnam is a compelling case of analysis for at least two reasons. First, because it is a transition economy, moving from a Communist regime to a market-based economic system. This transition necessitates a new design for all institutional and regulatory settings, with far-reaching implications for all economic agents operating inside Vietnam and from outside the country (trade partners, for example). It is, thus, appropriate and necessary to gain further understanding on how public intervention (public expenditures, revenues, investment, firms, and regulations) and its changes have influenced the economy, both in the short and long term. Second, alongside this transition, there has been an increasing acceptance of international trade and capital markets. This implies that key macroeconomic variables, such as the balance of payments, exchange rates, oil prices, and interest rates, have become progressively more influential in determining Vietnam's economic performance. This second block of variables, related to the growing integration of Vietnam in global markets, merits further analysis and understanding.

Consequently, all the investigation into the economic tools and methods undertaken during the development of this thesis has been applied to the two major structural processes of transition and internationalisation experienced by this rapidly evolving country. I hope these essays will be a steppingstone towards a better comprehension of Vietnam's performance in the recent decades and, hopefully, will be found helpful in the debate on how to keep Vietnam moving and growing.

1.2 Research structure

The dissertation is structured into four chapters (Chapters 2 to 5) corresponding to four independently empirical papers. Together, these provide a panoramic view of Vietnam's economy from various standpoints. The first and last of these empirical works investigate the cyclicality of the Vietnamese economy, using dynamic stochastic general equilibrium (DSGE) models with neoclassical economic theory at their cores. The second study, Chapter 3, aims to explain how global demand and oil price shocks impact on Vietnam's macroeconomic conditions and monetary policy responses. Chapter 4 acts as a complement to the DSGE analysis of public expenditures conducted in chapter 5. This chapter also validates Aschauer's (1989a, b) hypothesis on the crowding-in effect of public investment and public capital.

1.2.1 Growth and Real Business Cycles in Vietnam and the Asean-5

Chapter 2 examines Vietnam's economy from a real business cycle perspective and compares its performance to that of ASEAN-5's economies. My first objective was to account for the supply-side factors that have driven Vietnam's economic growth over the period 1976 – 2015. It was shown that total factor productivity (TFP) could be attributed to one-third of income growth on average in 1981-2015, while it fell to less than one-fifth in the 2000s. In addition, capital accumulation became the main driver of growth since 1992, with human capital increasing its contribution. Using well-known detrending techniques, I also discovered that that the cyclical behaviours of the Vietnamese macroeconomic aggregates are similar to those of its ASEAN-5 peers and other emerging countries.

In my second task, I methodologically extended the small-open-economy RBC models developed by Aguiar and Gopinath (2007) (AG) and García-Cicco, Pancrazi, and Uribe (2010) (GPU), and demonstrated that the estimated DSGE-RBC model performed better than AG and GPU in several dimensions, capturing Vietnam's economic regularities over the sample period between 1981 – 2015 (35 years). The model appeared to flawlessly recreate the downward slope autocorrelation between output and trade balance (as a percent of GDP), which was an unresolved issue in AG's model. Secondly, due to the presence of habit persistence and government consumption in the period utility function, it outperformed the GPU's financial friction setting by simultaneously reproducing the moments of growth variables while matching the low value of trade-balance-to-output autocorrelation. The long-run variance decomposition revealed that transitory productivity shocks explain approximately one-half of Vietnam's output volatility, whilst exogenous risk premium and trend shocks each account for only one-fifth of that volatility. This critically indicates that AG's claim that "the cycle is the trend" cannot be sustained in the case of Vietnam as non-stationary shocks to TFP only accounted for 12% of the Vietnamese Solow residual's variance. My empirics align to those used in GPU and Rhee (2017), in which the stationary component is overwhelmingly dominant.

In order to gain insight into policy implications, I simulated the trajectories of output growth and the trade-balance-to-output ratio by sequentially turning off several exogenous processes. The outcome of these counterfactual analyses is twofold. First, transitory productivity shocks have a significant impact on Vietnam's income growth but not on trade balance. Neither trend nor productivity shocks could explain the huge trade deficits experienced in 2007 – 2010. Instead, interest rate shocks greatly affected the trade balance and helped to stabilise the growth path of Vietnamese output. This second result, which was striking as it reveals the strong effects of Vietnam's proactive monetary policy in the past two decades, was consistent with the findings in Anwar and Nguyen (2018) and supported Huynh et al.'s (2017) results. Therefore, I would argue that technological progress and productivity-enhancing measures are vital for Vietnam's economy to sustain high growth.

I concluded the chapter with an examination of Thailand, which is the main business rival of Vietnam and has historically been the star economy within the ASEAN-5. The explanatory power of both transitory and trend shocks was relatively weak as each productivity innovation accounted for only one-quarter of the output growth variance. In this context, the trend component of the shock only accounted for 6% of Solow residual's variance. In contrast to Vietnam, and similar to Korea (Rhee, 2017), country risk premium innovations appear to govern Thailand's business cycles, implying that Thailand's economy was more vulnerable to international externalities than Vietnam's, as the latter country has stricter capital flow controls and its economy is 'de-facto' a non-free-market economy.

1.2.2 The macroeconomic effects of oil price and risk premium shocks on Vietnam

One of main findings in Chapter 2 was that the role of interest rate (or risk premium) shocks in governing and stabilising the Vietnamese trade balance and output growth, respectively, signified the high sensitiveness of the Vietnamese economy in respect to international shocks, such as cross-border demand of goods and services, commodity prices, etc.

Accordingly, Chapter 3 researches the macroeconomic impact of oil price shocks, based on Kilian (2009), on Vietnam's typical macroeconomic indicators, i.e. trade balance, three-month Interbank interest rates, inflation rate, and real (effective) exchange rates (RER or REER). The analysis, which was performed on a unique dataset with variables defined at a monthly frequency running from 1998:01 to 2018:12, provided an advanced understanding of the mechanisms through which such shocks may shape Vietnam's economic performance.

To conduct the analysis, I constructed the Kilian-based (2009) small, open economy structural autoregression (SVAR) models and examined the macroeconomic consequences of different oil price shocks, including (i) oil supply shocks; (ii) oil demand shocks reflecting changes in the level of global economic activity (also called global demand oil shocks); (iii) oil-specific demand shocks, which are also referred to as precautionary, speculative or non-fundamental demand shocks; and (iv) international risk premium shocks proxied by the United States Federal Fund rate innovations. The baseline setting yielded the first important insight into Vietnam's economy, namely, that its inflation rate and real exchange rate are responsive to both types of oil demand shocks (and not to the supply-side shock).

Three other augmented models allowed for a deeper evaluation of how oil price and risk premium shocks affect Vietnam's competitiveness and monetary policy. Several of the findings drawn from these models should be considered valuable.

The first of these findings was that the insignificant influence of oil supply disruptions was confirmed by both the impulse response functions (IRFs) of the RER and the REER. Secondly, the harm caused by international price-competitiveness was also confirmed when oil price shocks arise either from global demand or speculative activities. Impulses, in both cases, strengthened the VND for at least a year, thereby leading to cheaper foreign goods for VN households. Thirdly, the research uncovered the remarkable influence of the U.S. Federal Fund rate, due to the strong tie between the two respective currencies. The policy instrument, i.e. 3-month Interbank interest rate (and/or short-term interest rates), appeared to be greatly sensitive to both types of oil demand shocks and to changes in international financial risk. It has been shown, however, that Vietnam's authorities acted quite conservatively in their reaction to international demand shocks and failed to counteract the inflationary pressures brought on by those shocks. As such, Bhattacharya's (2014) call for forward-looking monetary policy in Vietnam is endorsed by our analysis.

Further variance decomposition shows that both types of oil demand shock play an essential role in explaining the long-run variations of several VN macroeconomic indicators. Oil demand shocks, in particular, affect the trade balance, whereas short-term interest rates are strongly influenced by oil-specific demand innovations. In addition, both types of oil price shock are equally significant to the inflation rate.

Finally, I considered the two inflationary periods between 2007 – 2009 and 2010 – 2012. It is likely that Vietnam's monetary policy was, to some extent, inefficient in the former inflationary period. In the latter period, however, domestic aggregate demand and oil price declines due to non-fundamental innovations decreased Vietnam's inflation rate and counterbalanced the impact of global demand oil price shocks due to precautionary reasons. As shown by Lorusso and Pieroni (2018), socio-economic and political tensions around the world, between 2011 and 2014, resulted in a rising precautionary demand for oil. This was the only significant foreign driver of inflation in Vietnam in those years and provides a key example of how global shocks may affect domestic macroeconomic performance, requiring an appropriate policy response.

1.2.3 Crowding-in or Crowding-out macroeconomic effect of public investment in Vietnam

The claim that "technological progress and productivity-enhancing measures are fundamental for Vietnam's economy to sustain a high growth", from Chapter 2, relates to the assumption that government spending is an endogenous factor in the growth function (Ortigueira and Santos, 1997; Romer, 1994). The nexus between government spending on capital stock (public investment), private investment, and economic growth has become one of the most important topics on the macroeconomic research agenda since the highly influential works of Aschauer (1989a, b).

Chapter 4 aims to address a void in the existing literature by examining the responses of output and private investment (capital) to an injection of public capital investment from the Vietnamese state. Two aspects were found to be worthy of study. First, I tested the 'Granger causalities' between public (private) capital accumulation and output and hours-worked. Second, I investigated the macroeconomic transmission mechanism of public investment in Vietnam during its transitional phase, validating Aschauer's hypothesis. To this end, I established and estimated two cointegrated vector autoregression systems, considering structural breaks in Vietnam's economy over the sample period 1976 – 2015.

The findings robustly document the positive impacts of public (investment) capital on private (investment) capital and output growth, highlighting the crowding-in effect of public (investment) capital investment in the transitional economy of Vietnam. Interestingly, our analysis showed that variations in employment could be largely explained by private capital accumulation. As a result, an appropriate public investment strategy is recommended for Vietnam's economy, with a focus on 'core' infrastructure to promote economic development and private sector.

1.2.4 On the implications of public expenditures on Vietnam's business cycles

The above analysis of the role of public investment in Vietnam's economy naturally provokes the question, "how does public expenditure affect the economy in the sense of real business cycle theory?" The last chapter (Chapter 5) offers a deeper exploration into how the mechanism of government expenditure (investment and consumption) affects the Vietnamese output fluctuations. To achieve this, I constructed a DSGE model, which was heavily influenced by the work of Leeper et al. (2010a, b). This research uncovered several empirical contributions.

Firstly, the estimated DSGE model captured the Vietnamese cyclical aggregate moments. It was shown that public investment positively affected Vietnamese economic fluctuations in the early-to-mid 1990s. During the seven-year period between 1998 and 2006, the output was stable, although public capital spending seemed to result in less productivity as productivity shocks, historically, negatively impacted on output variations. Secondly, it was demonstrated that government expenditures accounted for about 29% and 18% of the cyclical output variance in the long and short run, respectively. The study of impulse – response functions suggested that public investment caused a noticeable beneficial effect of around 0.91% GDP of contemporaneous private income response or 3% real GDP of five-year accumulative of that response to 1% real GDP of public investment impulse. Lastly, the deterministic simulation shows that in the event of an expansionary policy design matters and the efficiency of public projects is of paramount importance.

CHAPTER 2: Growth and Real Business Cycles in Vietnam and the Asean-5. Does the trend shock matter?[♥]

Abstract

I examine Vietnam's economy together with its closest trade partners. I 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. I also show that the cyclical behavior of its macro aggregates is similar to the one of its ASEAN-5 peers and other developing countries. I 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 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 cycle is the trend" hypothesis in Aguiar and Gopinath (2007), and align to those in García-Cicco, Pancrazi, and Uribe (2010) and Rhee (2017), in which the stationary component is overwhelmingly dominant. I therefore argue that technological progress and productivity-enhancing measures are fundamental for Vietnam's economy to sustain a high growth.

Keywords: Vietnam, ASEAN, DSGE, real business cycles, trend shock, growth
JEL codes: E32, F43, 053

[♥] This chapter has been published under the same title to Economic Systems Journal (forthcoming in 2020).

2.1 Introduction

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

The contribution of the research is threefold. First, I extend the RBC models developed by Aguiar and Gopinath (2007) (AG) and García-Cicco, Pancrazi, and Uribe (2010) (GPU). I show that my extended setup provides a better account of the facts. In particular, I am able to match the downward slope autocorrelation between output and the trade balance (as percent of GDP), which was an unresolved issue in AG's model. Second, I account for the 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 nor other emerging market economies. Third, I provide evidence that Vietnam's business cycles have been mainly 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's residual volatility (6% in Thailand), for these cases I refute their view that "the cycle is the trend" in emerging markets.

My first task is to account for the supply-side factors that have driven Vietnam's economic growth. I show that TFP accounts for one third of economic growth on average in 1981-2015, while it falls in the 2000s to less than one fifth. In addition, capital accumulation becomes the main driver of growth since 1992, with human capital rising its contribution. Then, I use well-known filtering techniques (e.g., Prescott 1986; King and Rebelo 1999; and Stock and Watson 1999, to name a few) to account for Vietnam's business cycle fluctuations in the period from 1986 to 2015. I find that the business cycle characteristics of Vietnam's national account components are essentially consistent with those reported by 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, I aim at quantifying the exogenous forces that have shaped the dynamics of Vietnam's growth aggregates. Hence, I 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 as many supports as objections in subsequent studies. For example, Suzuki (2018a, b) 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 by 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, L'Huillier, and Yoo (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 risk premium reaction causes the AG hypothesis to hold only in extremely unrealistic cases. In the same vein, Durdu (2013) concludes, after revising the literature, 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, I 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, my proposed RBC model outperforms GPU's financial friction specification when reproducing the moments of the growth variables at the same time that matches the low value of trade-balance-to-output autocorrelation (0.18 after four lags). Although the presence of habit formation improves the moment matching capability, I 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 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, I do counterfactual simulations under three scenarios. First, I 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 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, my simulations strengthen Huynh et al. (2017)'s 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.

I conclude the analysis by examining Thailand, which is the main competitor of Vietnam and has been the star economy within the ASEAN-5 historically. The explanatory power of both the transitory and trend shocks is relatively weak because each productivity innovation accounts only for one-fourth of the output growth variance. In this context, the trend component of the shock only accounts for 6% of 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. I therefore argue that Thailand's economy was more vulnerable to international externalities than Vietnam 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 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. Section 7 concludes.

2.2 Business cycle analysis

2.2.1 Data and Background

Macroeconomic research on Vietnam's economy is challenging due to limited data availability. For output, the most recent quarterly data just covers years after 2000, but not all the other main aggregates are publicly provided. I collect annual aggregate data from the United Nations Statistical Division (UNSD) because the time coverage is long enough to identify the meaningful business cycles. To economize notation, I define a vector of main aggregate variables J ≡ (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, which transformed Vietnam from one of the world's poorest countries in the beginning of 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 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, 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 Vietnam War consequences. 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 in 2001 – 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.

		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

Note: 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.

Table 1: Descriptive statistics of main aggregate variables

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% 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 I 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 was much less fluctuating than its peers in the last two decades. Conversely, Singapore appears to be the most fluctuating economy. Note, finally, that Indonesia and Thailand experienced larger TFP growth rates positive than Vietnam in 2001–2015 (3.4% and 2.2% against 2.0%). Later on, I briefly examine Thailand's economy, while the unique case of Indonesia has been analysed in Lee and Hong (2012).

	Solow residual's growth rate, %				Standard deviation (σ_{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)

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 $\sigma_{sr} = \sqrt{Var(\varepsilon_t^{sr})}$.

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

2.2.2 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 out of the trend in several ways. In this chapter I consider four standard econometric techniques which are first-differencing, linear regression, Hodrick and Prescott (1997) (HP) filter and the one-sided variant of 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 (σ_j/σ_y). I apply all four different filters discussed above to extract cyclical signals out of the trends. The outcome is presented in Table 3.

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

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

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

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 to 2.5 times larger than the HP and first-differencing counterparts (see Figure 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 one in 1986 – 1997 and the second one in 1997 – 2008. Then, in the aftermath of the GFC, Vietnam's output has evolved below the trend. Another relevant outcome is the different behaviour of the ASEAN-5 countries, where 10-year cycles involving 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 ASEAN-5 countries.

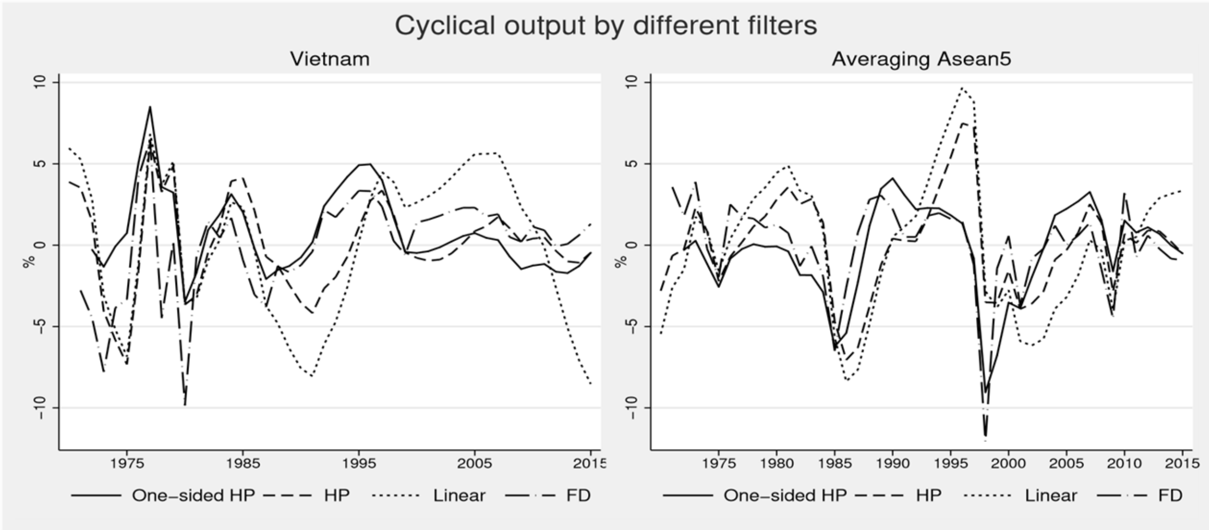


Figure 1: Dynamics of Vietnam and ASEAN-5 economies

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 the Vietnam's intense demand of imported 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, explains at least partially the higher consumption-to-output standard deviation ratios in Vietnam with respect to the ASEAN-5's 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 the height 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 Vietnam and ASEAN-5's foreign trades were 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, 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 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).

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

Note: 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.

Table 4: Cyclical Correlations

2.3 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 likes Vietnam and the ASEAN-5 peers. To respond to this question, I develop a small open economy DSGE-RBC model aiming to quantify the exogenous forces that shape the dynamics of Vietnamese growth aggregates (and Thailand's ones in a final comparative exercise).

2.3.1 Setting the economic environment

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

I 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 μ_g being the gross long-run growth rate and $\left|\rho_g\right| < 1$. The single trend shock ε_t^g is assumed normally distributed with variance σ_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 δ represents the rate of depreciation and ϕ 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, Hercowitz, and Huffman (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,

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

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 ν_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. I model $C_{g,t} = \zeta_{cg}\xi_{cg,t}Y_t$, with the spending shock $\xi_{cg,t}$ and a constant ratio of ζ_{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; and 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 my case, by incorporating government spending into the period utility I 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, ν_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

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

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

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

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 μ_t. Since I am using the GPU's model as benchmark, the rule for the exogenous risk premium channel is restated as

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

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 the debt elasticity; and $\bar{b}$ is the steady-state level of governmental outstanding debt.

My model is richer than GPU's financial friction setup in the number of exogenously stochastic processes as it allows to distinguish shocks to the pure-time discount rate (ξ_p) from shocks to the current consumption (ν_t). The former is also known as a preference-shifter since it lets parameter β 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

$$E_0 \sum_{t=0}^{\infty} \xi_{p,t} \beta^t \left\{ \frac{\left[\left[\nu_t C_{p,t} - \gamma C_{p,t} + \pi C_{g,t} - \theta \omega^{-1} X_{t-1} h_t^{\omega}\right]^{1-\eta}}{1-\eta} - \right. \tag{8}$$

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

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

Along with equations (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,

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

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

Hours-work,

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

Capital,

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

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 her current consumption taking the previous aggregate private spending as given so that (9) reduces to

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

It is obvious that the preference shifter ξ_p does not perturb Lagrangian multiplier λ_t in (13). This implies that an external habit specification is inefficient in my 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 the factor X_t in equation (1). It is thus necessary to detrend all equilibrium conditions, which I do follow AG.[2]

2.3.2 The long-run equilibrium

I 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

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

Equation (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 the value of η; secondly, under a higher long-run labor-augmenting growth rate μ_g; third, under a lower subjective discount factor β.

---

[1] Identification test (Iskrev, 2010) reports $\xi_{p,t}$ and ν_t are pairwise multi-collinearity if γ is external.

[2] I 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 the factor X at time $t-1$, and so on.

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

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

I calibrate ω to a value greater than unity, which is the standard in RBC literature. Subsequently, an increase in the internal habit formation coefficient, γ, 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 β 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 equations (17) – (19) , are directly affected by the deep parameters implied in equations (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 ζ_{cg} and ζ_b are, respectively, the shares of government consumption and the stock of bonds in the steady state.

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

2.4 Estimation and discussions

I 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 θ (or posterior distribution shortly) conditional on observational data Y. The prior distribution $\Gamma(\theta)$ is the unconditional probability distribution of θ, 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, a Bayesian estimation requires three key elements: advanced knowledge of prior specifications, a suitable filter for 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).

2.4.1 Configuration

A time unit is meant to represent a year in the model. Econometrically, not all model "deep" 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, I set the value of ω to 2.0, which results in θ = 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, I set the capital income share to α = 0.35. The depreciation rate, δ, and the curvature of the period utility function, η, take the common values of 10% and 2, respectively. Next, the ratio of government consumption, ζ_{cg}, and stock of domestic bond to output, ζ_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.

I neither calibrate the subjective discount factor β nor the long-term growth rate μ_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, I 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]

Calibrated params	β^*	π	θ	ω	η	α	δ	μ_g	ζ_{cg}	ζ_b
Value	0.03-0.07	1	1.85	2.00	2.00	0.35	0.10	1.048	0.07	-0.36

Table 5: Calibrated parameters

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), I adopt Rhee (2017)'s prior specifications.

It is worth emphasizing that an advantage of directly modeling non-stationary variables in the 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.

[3] Even though it passes the Iskrev (2010)'s test, the identification strength of π is weak (with a hardly distinguishable posterior from the prior distribution). Figure A4 and A5 provides a sensitivity analysis to study how the variability of parameter πaffects the modelled variables. The habit persistence coefficient appears to be the strongest parameter, while the elasticity of public spending π only has considerable effects on the responses of g_y and g_{cp} with respect to two shocks ε_{cg} and ε_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, Levine, and Melina (2014).

Canova (2014) stresses on the consistency between model and data when both permanent and transitory shocks coexist. Along the same line, 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.

My 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

$$g_{y,t} = \Delta lnY_t \quad = \log(y_t) - \ln(y_{t-1}) + g_{t-1} - 1$$

$$g_{cp,t} = \Delta lnC_{p,t} = \log(c_{p,t}) - \log(c_{p,t-1}) + g_{t-1} - 1$$

$$g_{cg,t} = \Delta lnC_{g,t} = \log(c_{g,t}) - \log(c_{g,t-1}) + g_{t-1} - 1$$

$$g_{inv,t} = \Delta lnI_t \quad = \log(i_t) - \log(i_{t-1}) + g_{t-1} - 1$$

$$tby_t = TB_t/Y_t = (y_t - c_{p,t} - c_{g,t} - t_t)/y_t$$

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 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 (I 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; and Benchimol and Fourçans 2017).

2.4.2 Results

2.4.2.1 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, of approximately 6.4% per annum (equivalent to 1.6% per quarter), implying that the stationary value of real interest rate $r = \mu_g^\eta/\beta$ would be in the range of [12.2%, 16.2%] corresponding to η = [1.2, 2.0] and μ_g = 1.045. As a sensitivity check, I run Bayesian estimations at each η = (1.25, 1.50, 1.75, 2.00), which still deliver the consistent interval of $\beta^* \in$ [6.3%, 6.9%]. The low value of β for Vietnam (0.94), when compared to the one obtained for the US by King and Rebelo (1999), β = 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 1990s, 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.

Estimated params	Prior			Posterior			
	Dist.	Mean	SD	Mean	5%	95%	p-val
μ_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
ψ	*Gamma*	0.50	0.50	0.322	0.000	0.629	0.794
γ	*Beta*	0.50	0.20	0.268	0.093	0.435	0.932
ϕ	*Gamma*	5.00	2.00	3.844	1.907	5.844	0.793
ρ_a	*Beta*	0.50	0.20	0.720	0.513	0.933	0.584
ρ_g	*Beta*	0.50	0.20	0.608	0.372	0.851	0.751
ρ_p	*Beta*	0.50	0.20	0.579	0.296	0.882	0.110
ρ_{cg}	*Beta*	0.50	0.20	0.774	0.619	0.934	0.892
ρ_μ	*Beta*	0.50	0.20	0.781	0.598	0.953	0.711
ρ_v	*Beta*	0.50	0.20	0.739	0.521	0.934	0.504
σ_a	*IGamma*	0.01	0.20	0.007	0.003	0.001	0.088
σ_g	*IGamma*	0.01	0.20	0.005	0.003	0.008	0.159
σ_p	*IGamma*	0.05	0.20	0.043	0.012	0.080	0.127
σ_{cg}	*IGamma*	0.05	0.20	0.027	0.016	0.037	0.978
σ_μ	*IGamma*	0.05	0.20	0.026	0.014	0.037	0.405
σ_v	*IGamma*	0.05	0.20	0.039	0.014	0.064	0.021
σ_{gy}	*IGamma*	0.01	0.20	0.009	0.004	0.014	0.437
σ_{gcp}	*IGamma*	0.01	0.20	0.018	0.013	0.023	0.931
σ_{gcg}	*IGamma*	0.01	0.20	0.012	0.003	0.025	0.874
σ_{ginv}	*IGamma*	0.10	0.50	0.094	0.067	0.120	0.398
σ_{tby}	*IGamma*	0.05	0.20	0.016	0.011	0.022	0.530

Note: Beta denotes Beta distribution; (I)Gamma represents (inversed) gamma distribution. p-val is p-value (15%-taper) of Geweke (1992) Convergence Test (see appendix for additional trace plots). The estimation is based on 1000,000 draws from the MHMC algorithm.

Table 6: Estimated parameters for Vietnam (1981 – 2015)

The estimated value of ψ 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 $\phi \sim 4.0$ 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%, 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.

2.4.2.2 Simulated results

The top block of Table 7 reports the striking performance of the model (column W. habit) , as it can effectively reproduce the 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 (σ_{gcp}/σ_{gy}, $\sigma_{ginv}/\sigma_{gy}$, $\sigma_{ginv}/\sigma_{gy}$, σ_{tby}/σ_{gy}) are higher than those of the full model; and neither the downward slope of the trade-balance-to-output autocorrelation function (Figure 2) nor the excess variation of household consumption to output can be reproduced in the GPU's basic setting. When GPU's financial friction (FFR) is added, the output growth variance is over predicted while the $\sigma_{ginv}/\sigma_{gy}$ 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 behaviors. 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 habit persistence strength 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).

			Model		GPU	
Moments	Data		W. habit	W/o. habit	Basic	FFR
σ_{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

Note: 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.

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

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

Actual expresses HP filtered data (λ = 100). *corr* and *AR1* are contemporaneous and 1st-order autocorrelation, respectively.

Table 8: Simulated (auto) correlations of level variables.

I need to acknowledge that neither the current model nor GPU's one 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 I used in comparisons to the 100-year datasets of Mexico and Argentina in GPU's study. On the other hand, Table 8 demonstrates that the

[4] Havranek et al. (2017) explore 81 studies covering Australia, New Zealand, G-7 group, and many EU countries.

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

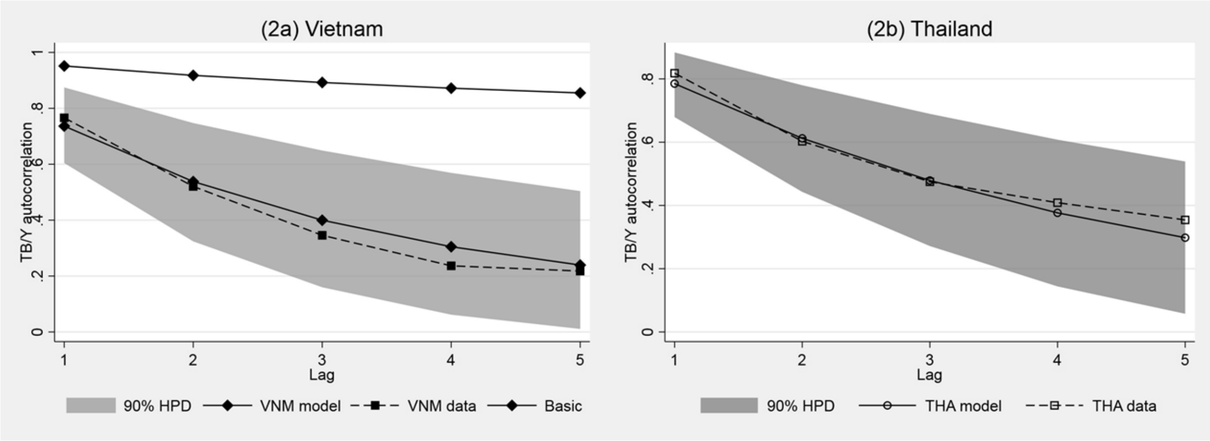


Figure 2: Trade-balance-to-output ratio (*tby*) autocorrelation of Vietnam (left panel) and Thailand (right panel). VNM (THA) model refers to my proposed RBC model; VNM basic is the standard RBC prototype in GPU.

2.4.3 Variance decomposition

Using Kalman's filter, I 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 I use my extended model (89% when using GPU's one). 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 account for about 20% each 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 ε_μ on the observable *tby*, I decompose its variance over time, as displayed in Figure 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, ε_μ.

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]

Note: 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.

Table 9: Unconditional variance decomposition (in percentage, %)

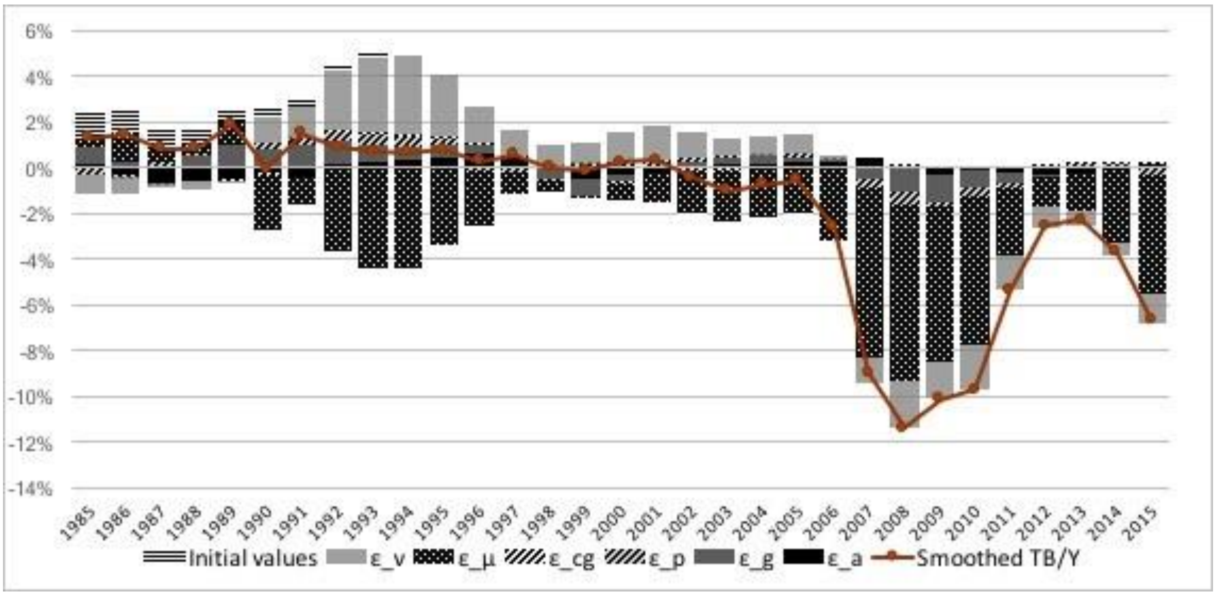


Figure 3: Historical variance decomposition of *tby* (Vietnam)

2.4.4 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

$$\frac{\sigma_{\Delta\tau}^2}{\sigma_{\Delta sr}^2} = \frac{\alpha^2 \sigma_g^2}{\left(1-\rho_g\right)^2 \sigma_{\Delta sr}^2}$$

where sr_t denotes the conventional Solow residuals, and τ_t is the trend part of the BN decomposition, such that $sr_t = \tau_t + s_t$. In Figure 4, I report all relative variance ratios up to lag 12.[5] It appears that, with the exclusion of the Philippines, the trend seems not 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 σ_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 σ_a and σ_g

[5] I find negligible differences between Cochrane (1988) and AG-modified formula, but the former is easier to compute.

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]

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

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 Figure 4.

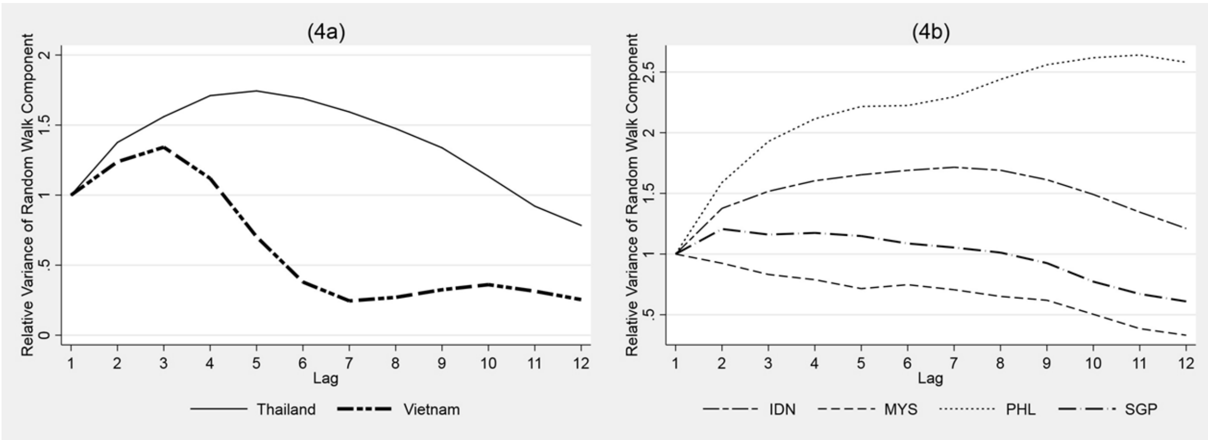


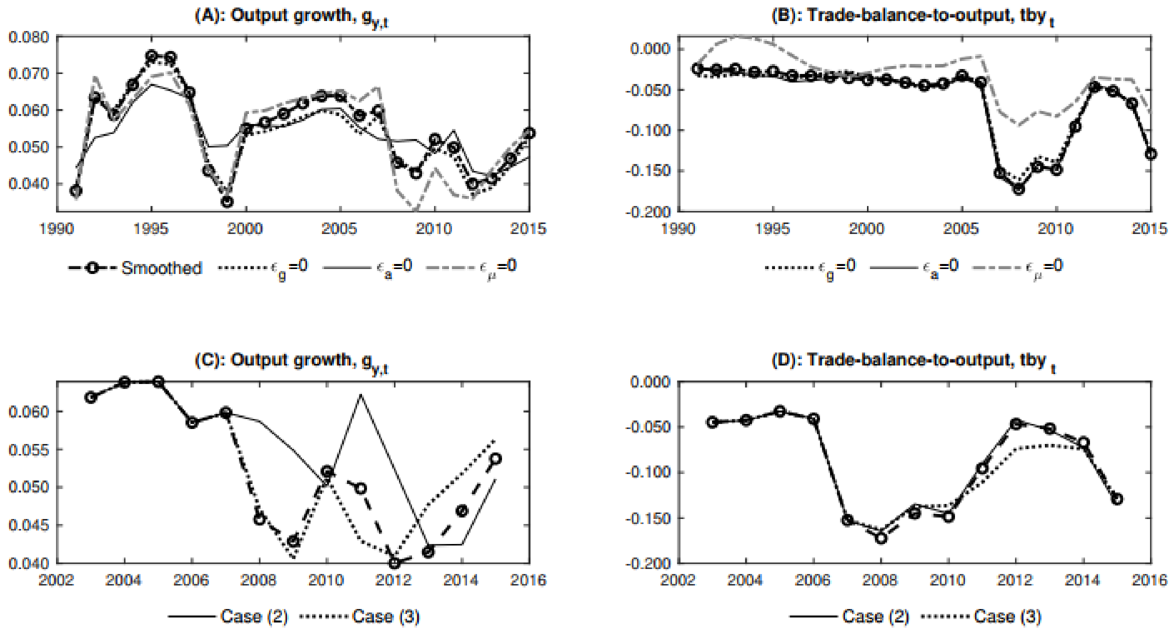
Figure 4: Relative variance of Solow residual's components using Cochrane's (1988) formula. Panel (4a): Vietnam and Thailand; Panel (4b): Indonesia, Malaysia, the Philippines and Singapore.

2.5 Counterfactual simulations

In this section, I investigate the trajectories of several observables under different loaded shocks. There are some interesting policy implications that arise from analysing fictitious scenarios and its consequences for Vietnam's economy. In Scenario 1, transitory productivity shocks, trend shocks and

[6] The fact is that $Var[(1-\alpha)g_t] = (1-\alpha^2)\sigma_g^2/(1-\rho_g^2)$ and $Var(a_t - a_{t-1}) = 2Var(a_t) - 2Cov(a_t, a_{t-1})$.

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 years 2008 – 2013. In Scenario 3, shocks to the country risk premium become silent ($\varepsilon_{\mu,t} = 0$) in 2008 – 2013.



Note: 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 $\varepsilon_a 2008_{,g} -2013 = \varepsilon_a{}^{1992}{}_{,g}{}^{-1997}$ (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.

Figure 5: Counterfactual simulation

The first row of Figure 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 I 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 lines (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 the trade balance would have stayed unchanged, while the one of income growth would

[7] 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).

have become more stable. This implies that this kind of shocks help stabilizing 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 the 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 (remark that in Panel A the positive income growth gap between actual data and the grey 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 on the leading role of the trend shock in explaining a sudden drop in 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 I draw from these exercises goes back to the first result. In terms of a high growth and sustainable pace, it seems that technological progress is the critical condition to achieve the best possible path for the Vietnamese economy.

Finally, let me note that the counterfactual exercises strengthen to some degree 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 my 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. I hope this lesson may be helpful in the future design of economic policy in Vietnam.

2.6 Application to Thailand

I extend my investigation by applying the analysis to Thailand, which has been, and still is, the key business rival of Vietnam 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".

	Thailand		Vietnam	
Period	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

(*) and (**) show Thailand 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 10: Thailand and Vietnam, 1976 – 2015

Having faced a debt crisis in 1983 – 1985, Thailand experienced a major boom-bust cycle in the second half of the 1980s and 1990s, as illustrated in Table 10. 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, [8] but its growth rate kept a pace of 3.97% per year and was classified as an upper-middle-income country in 2011

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

(World Bank, 2011). As Table 10 shows, in spite of this success, Vietnam's economy has outperformed Thailand's one by approximately two percentage points since 2000.

Except for ω and θ, the same calibration set is used to bring the model to Thailand data. Since Thailand is calibrated in AG, I 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 σ_{gy} is close to the actual data, and the downward slope of *tby*'s autocorrelation function has been exactly reproduced, as plotted in Figure 2. Nevertheless, the relative standard deviations are overly predicted because the standard model could not handle properly 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.[9]

Params	σ_{gy}	σ_{gcp}/σ_{gy}	σ_{gcg}/σ_{gy}	$\sigma_{ginv}/\sigma_{gy}$	σ_{tby}/σ_{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)
σ_a	σ_g	σ_p	σ_{cg}	σ_μ	σ_v
0.017	0.005	0.031	0.037	0.025	0.026
ρ_a	ρ_g	ρ_p	ρ_{cg}	ρ_μ	ρ_v
0.855	0.512	0.545	0.831	0.864	0.680
	ψ	γ	ϕ	β^*	μ_g
	0.095	0.128	3.766	0.032	0.038

Standard error in parentheses; σ denotes standard deviation. g_y, g_{cp}, g_{cg}, g_{inv} and *tby* variables are defined as in section 5.

Table 11: Bayesian estimation for Thailand, 1976 – 2015

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 counterpart. The latter implies that Thailand's economy is better structured even though capital adjustment costs, φ, are similar. The lower temporal discount rate reflects the stronger attachment of Thailand's households to their lifetime income. In support to this claim is the low estimate

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

of the habit persistence coefficient, $\gamma = 0.13$. In addition, the estimated parameter ψ, which controls for the sensitivity of the country risk premium, is one-third of Vietnam's counterpart. 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 explain Thailand's economy developments, as it accounts for 54.6% of the unconditional variance over the sample range.

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.

Table 12: Long-run variance decomposition for Thailand (in percentage, %), 1976 – 2015

2.7 Concluding remarks

This paper provides a detailed analysis of Vietnam's economy in connection to its ASEAN-5 peers. In the first part, I provide information on two complementary sides of Vietnam's performance in terms of growth sources and business cycle drivers. In the second part, I develop a DSGE-RBC model. My model departs from those in AG and GPU, but further incorporates habit formation. In this way it provides a close match of the facts in which my analysis of Vietnam, and the comparison with Thailand, can be conducted.

I 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 that year. One of the main consequences of such processes was the development of a structural trade deficit.

In terms of business cycle characteristics, I document a similar behavior in Vietnam than in the ASEAN-5 economies, which does not differ significantly from the one in other emerging market economies. The intrinsic difficulties of RBC models in replicating short-term observational dynamics is 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 my 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, I conclude that Thailand's economy is more vulnerable to international externalities than Vietnam's one, which has stricter capital flow controls and is still, *de-facto*, a non-free market economy. It is probably on account of Vietnam's uncompleted transition that technological progress and productivity-enhancing measures come out as fundamental to secure 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, 12% in Vietnam and 6% in Thailand. These results, which refute AG's claim that "the cycle is the trend" in emerging market economies, add-up to recent literature with similar results for Argentina, Mexico and Korea.

References

ADB, 2015. THAILAND Industrialization and Economic Catch-up. Asian Development Bank.

Aguiar, M., Gopinath, G., 2007. Emerging market business cycles: The cycle is the trend. Journal of political Economy 115 (1), 69–102.

Alvarez-Parra, F., Brandao-Marques, L., Toledo, M., 2013. Durable goods, financial frictions, and business cycles in emerging economies. Journal of Monetary Economics 60 (6), 720–736.

An, S., Schorfheide, F., 2007. Bayesian analysis of dsge models. Econometric reviews 26 (2-4), 113–172.

Anwar, S., Nguyen, L. P., 2018. Channels of monetary policy transmission in vietnam. Journal of Policy Modeling 40 (4), 709 – 729.

URL http://www.sciencedirect.com/science/article/pii/S0161893818300292

Benchimol, J., Fourçans, A., 2017. Money and monetary policy in the eurozone: an empirical analysis during crises. Macroeconomic Dynamics 21 (3), 677–707.

Boldrin, M., Christiano, L. J., Fisher, J. D., 2001. Habit persistence, asset returns, and the business cycle. American Economic Review, 149–166.

Bosworth, B., Collins, S. M., 2003. The empirics of growth: An update. Brookings papers on economic activity 2003 (2), 113–206.

Boz, E., Daude, C., Durdu, C. B., 2011. Emerging market business cycles: Learning about the trend. Journal of Monetary Economics 58 (6-8), 616–631.

Camen, U., et al., 2006. Monetary policy in vietnam: the case of a transition country. Press & Communications CH 4002 Basel, Switzerland 232.

Canova, F., 2014. Bridging dsge models and the raw data. Journal of Monetary Economics 67, 1–15.

Canova, F., Ferroni, F., 2011. Multiple filtering devices for the estimation of cyclical dsge models. Quantitative Economics 2 (1), 73–98.

Cantore, C., Levine, P., Melina, G., 2014. On habit and utility-enhancing government consumption. Tech. rep., Discussion Paper Series 14/06, City University London.

Cao, D., L'Huillier, J.-P., Yoo, D., 2016. When is the trend the cycle. Tech. rep., Mimeo.

Christiano, L. J., Eichenbaum, M., 1992. Current real-business-cycle theories and aggregate labor-market fluctuations. The American Economic Review, 430–450.

Christiano, L. J., Eichenbaum, M., Evans, C. L., 2005. Nominal rigidities and the dynamic effects of a shock to monetary policy. Journal of political Economy 113 (1), 1–45.

Cochrane, J. H., 1988. How big is the random walk in gnp? Journal of political economy 96 (5), 893–920.

Constantinides, G. M., 1990. Habit formation: A resolution of the equity premium puzzle. Journal of political Economy 98 (3), 519–543.

Cooley, T. F., Prescott, E. C., 1995. Economic growth and business cycles. Frontiers of business cycle research 1.

Correia, I., Neves, J. C., Rebelo, S., 1995. Business cycles in a small open economy. European Economic Review 39 (6), 1089–1113.

Del Negro, M., Schorfheide, F., 2009. Monetary policy analysis with potentially misspecified models. American Economic Review 99 (4), 1415–50.

Durdu, C. B., 2013. Emerging market business cycles: Recent advances. International Economic Journal 27 (2), 183–199.

Fernández-Villaverde, J., 2010. The econometrics of dsge models. SERIEs 1 (1-2), 3–49.

Fernández-Villaverde, J., Rubio-Ramírez, J. F., 2004. Comparing dynamic equilibrium models to data: a bayesian approach. Journal of Econometrics 123 (1), 153–187.

Ferroni, F., 2011. Trend agnostic one-step estimation of dsge models. The BE Journal of Macroeconomics 11 (1).

Fuhrer, J. C., 2000. Habit formation in consumption and its implications for monetary-policy models. American Economic Review 90 (3), 367–390. García-Cicco, J., Pancrazi, R., Uribe, M., 2010. Real business cycles in emerging countries? American Economic Review 100 (5), 2510–31.

Greenwood, J., Hercowitz, Z., Huffman, G. W., 1988. Investment, capacity utilization, and the real business cycle. The American Economic Review, 402–417.

Havranek, T., Rusnak, M., Sokolova, A., 2017. Habit formation in consumption: A meta-analysis. European Economic Review 95, 142–167.

Herbst, E. P., Schorfheide, F., 2015. Bayesian Estimation of DSGE Models. Princeton University Press.

Hodrick, R. J., Prescott, E. C., 1997. Postwar us business cycles: an empirical investigation. Journal of Money, credit, and Banking, 1–16.

Huynh, P., Nguyen, T., Duong, T., Pham, D., 2017. Leaning against the wind policies on vietnam's economy with dsge model. Economies 5 (1), 3.

Iskrev, N., 2010. Local identification in dsge models. Journal of Monetary Economics 57 (2), 189–202.

Khorunzhina, N., 2015. Real business-cycle model with habits: Empirical investigation. Economic Modelling 46, 61–69.

King, R. G., Rebelo, S. T., 1999. Resuscitating real business cycles. Handbook of macroeconomics 1, 927–1007.

Lee, J.-W., Hong, K., 2012. Economic growth in asia: Determinants and prospects. Japan and the World Economy 24 (2), 101–113.

Miyamoto, W., Nguyen, T. L., 2017. Business cycles in small open economies: Evidence from panel data between 1900 and 2013. International Economic Review 58 (3), 1007–1044.

Neumeyer, P. A., Perri, F., 2005. Business cycles in emerging economies: the role of interest rates. Journal of monetary Economics 52 (2), 345–380.

Prescott, E. C., 1986. Theory ahead of business-cycle measurement. Carnegie-Rochester Conference Series on Public Policy 25, 11 – 44.

Ravn, M., Schmitt-Grohé, S., Uribe, M., 2006. Deep habits. The Review of Economic Studies 73 (1), 195–218.

Rhee, W., 2017. Can rbc models explain business cycles in korea? Macroeconomic Dynamics 21 (3), 599–623.

Schmitt-Grohé, S., Uribe, M., 2003. Closing small open economy models. Journal of international Economics 61 (1), 163–185.

Schmitt-Grohé, S., Uribe, M., 2004. Solving dynamic general equilibrium models using a second-order approximation to the policy function. Journal of economic dynamics and control 28 (4), 755–775.

Smets, F., Wouters, R., 2003. An estimated dynamic stochastic general equilibrium model of the euro area. Journal of the European economic association 1 (5), 1123–1175.

Solow, R. M., 1957. Technical change and the aggregate production function. The review of Economics and Statistics 39 (3), 312–320.

Stock, J. H., Watson, M. W., 1999. Forecasting inflation. Journal of Monetary Economics 44 (2), 293–335.

Suzuki, T., 2018a. Business cycles in post-conflict serbia: The cycle is the trend. Eastern European Economics, 1–15.

Suzuki, T., 2018b. Permanent productivity shocks, migration and the labour wedge: business cycles in south africa. Macroeconomics and Finance in Emerging Market Economies, 1–14.

Uribe, M., Schmitt-Grohé, S., 2017. Open Economy Macroeconomics. Princeton University Press.

Warr, P., 2005. Thailand Beyond the Crisis. Routledge Curzon.

World Bank, Aug. 2011. Thailand now an upper middle income economy.

URL http://www.worldbank.org/en/news/press-release/2011/08/02/thailand-now-upper-middle-income-economy World Bank, Apr. 2013. Vietnam: Achieving success as a middle-income country.

URL http://www.worldbank.org/en/results/2013/04/12/vietnam-achieving-success-as-a-middle-income-country

World Bank, July 2016. Vietnam 2035 - Toward Prosperity, Creativity, Equity and Democracy. No. 23724 in World Bank Publications. The World Bank.

URL https://ideas.repec.org/b/wbk/wbpubs/23724.html

Appendix

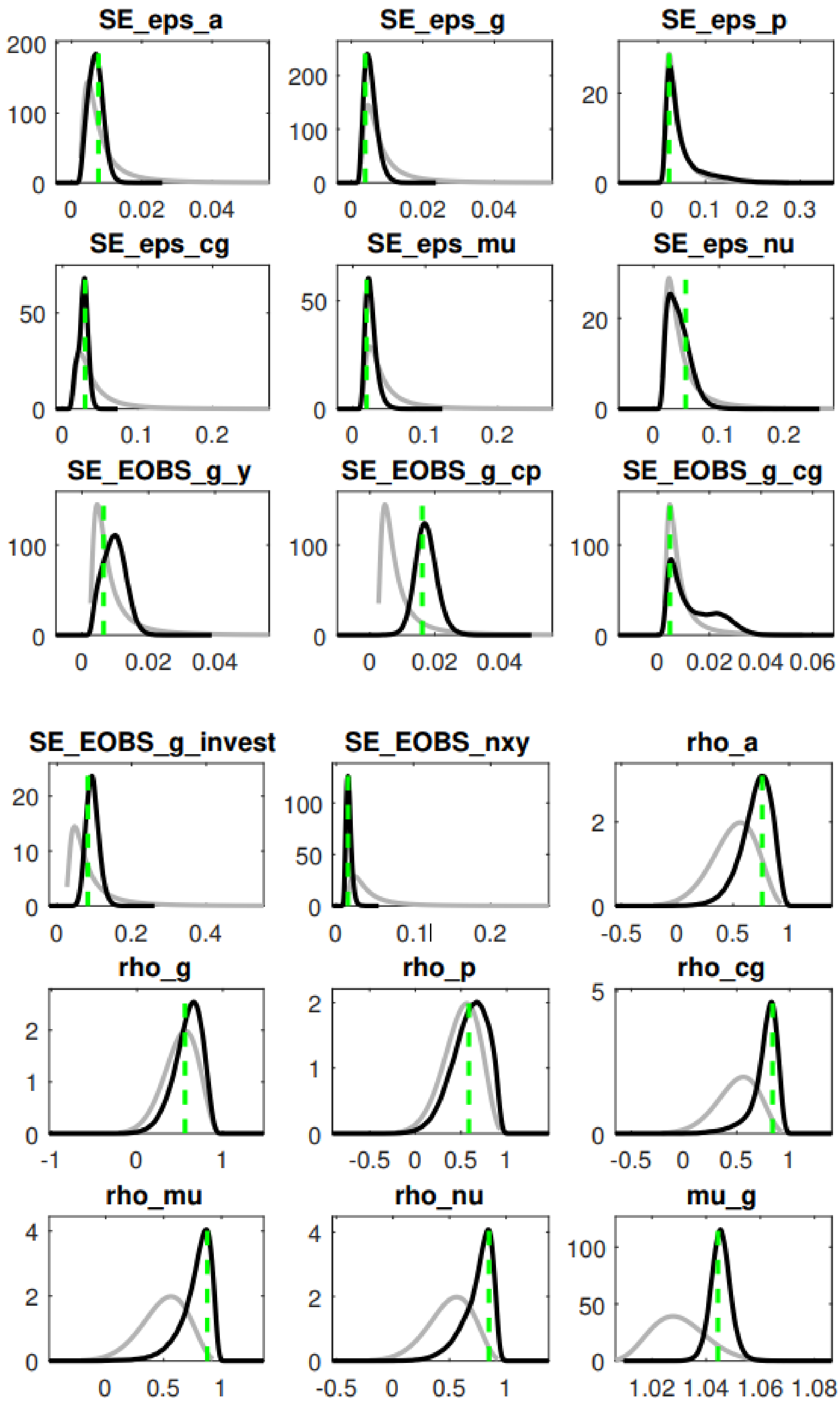


Figure A1: Priors and Posteriors (in case of estimated π with Beta prior)

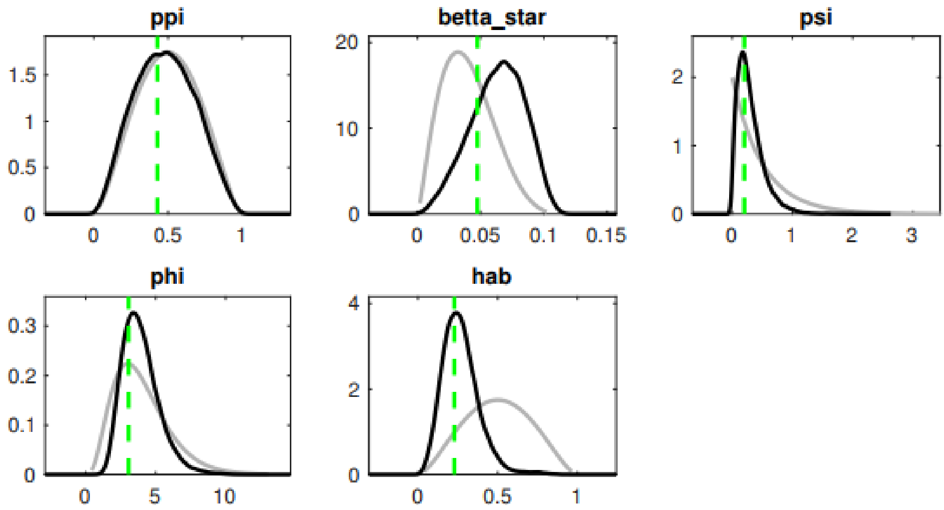
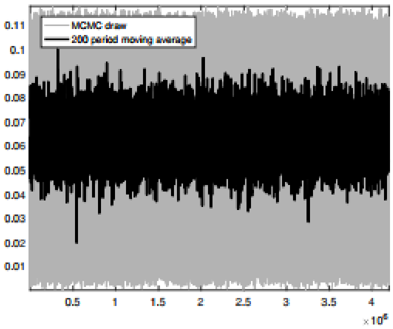
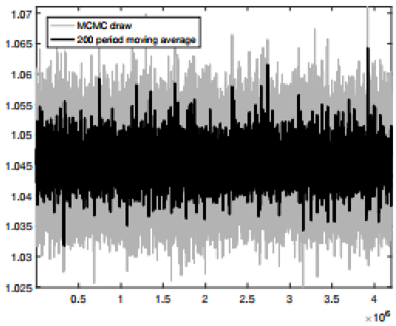


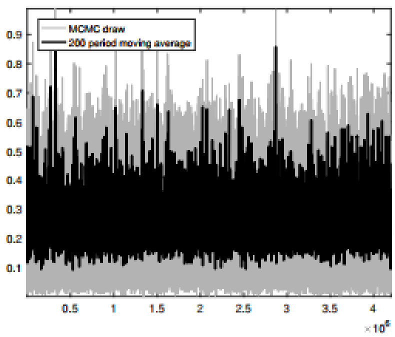
Figure A2: Priors and Posteriors (in case of estimated π with Beta(0, 1) prior)



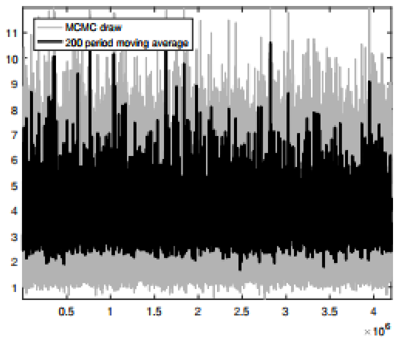
Subjective discount rate, β^*



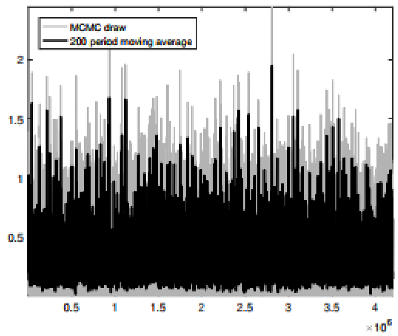
Long-run growth rate, μ_g



Habit persistence, γ



Capital adjustment costs, ϕ



Debt elasticity, ψ

Figure A3: Selected trace plots

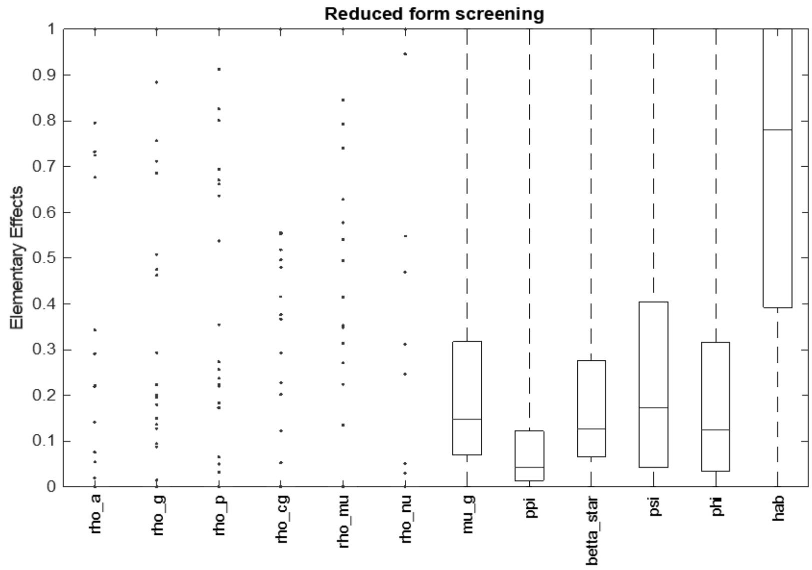


Figure A4: Elementary Effects analysis (in case of estimated π with Beta (0, 1) prior)

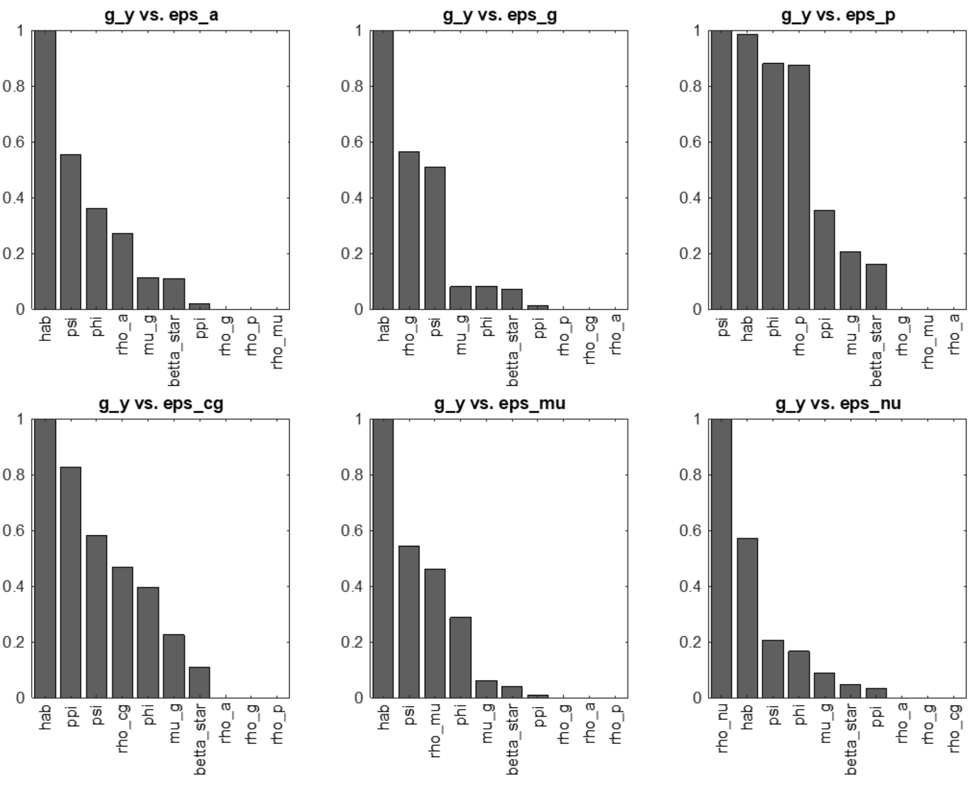


Figure A5.1: Sensitivity analysis w.r.t g_y (in the case of estimated π)

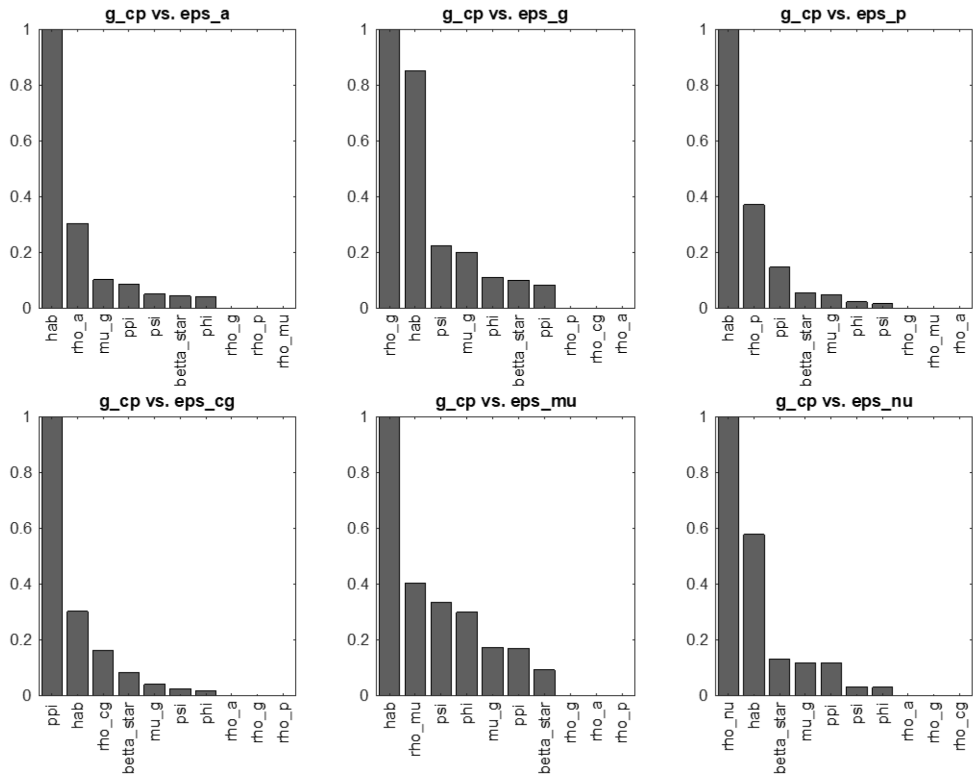


Figure A5.2: Sensitivity analysis w.r.t g_{cp} (in the case of estimated π)

CHAPTER 3: The macroeconomic effects of oil price and risk premium shocks on Vietnam – Evidence from an over-identifying SVAR analysis[♦]

Abstract

This chapter studies the macroeconomic effects of oil price shocks in Vietnam. It expands Kilian's (2009) framework to simultaneously consider risk-premium shocks and comprehensively assess their consequences on international competitiveness and the State Bank management of the monetary policy. Methodologically, this implies dealing with an over-identified structural vector autoregression (SVAR) model. Data wise, the analysis is performed on a unique dataset with variables defined at a monthly frequency running from 1998:01 to 2018:12. Demand-side, global-, and specific-oil price shocks determine inflation and international competitiveness, and play an essential role in explaining the long-run variations of several Vietnamese macroeconomic indicators (mainly the trade balance, three-month interest rates, and the inflation rate). Vietnam's Dong pegging to the US Dollar results in a stronger impact of these shocks when real exchange rates and the rate of exports are modelled, than when real effective exchange rates and the trade balance are modelled. In the latter case, shock absorption is quicker given the multilateral trade context in which no single pegging holds. In association to the strong tie between Vietnam's Dong and the U.S. dollar, I also uncover remarkable effects of risk-premium (or U.S. Federal Fund rate) shocks. Supply-side oil price shocks have little impact on inflation and international competitiveness but condition the monetary policy. Neglecting such influence in the past may have resulted in an excessively conservative monetary policy.

Keywords: Oil price shocks; risk-premium shocks; SVAR; international trade; Vietnam.

JEL codes: Q41; Q43; F41; F62.

[♦] This chapter has been submitted to Journal of International Trade & Economic Development.

3.1 Introduction

The strong upswing in crude oil price in the 2000s, and the subsequent Global Financial Crisis (GFC) in 2008 sparked an extensive research agenda on comprehending the causes and macroeconomic effects of oil price innovations.[1] In a seminal publication, Kilian (2009) claimed that the transmission of oil price increases may be different than traditionally understood, due not only to their fundamental underlying shocks (i.e. fundamental supply versus demand shocks), but also to their interactions with specific investment activities and with the biggest oil consumption country – the U.S. economy. It was also stressed that temporary oil production disruptions owing to political tensions were less important than fundamental shocks in explaining oil price movements and their consequences.

Taking an oil importer perspective, cost-push shocks due to oil and/or commodity price rises push up domestic prices of goods and services, thus triggering indirectly monetary policy measures to stabilise macroeconomic conditions. The transmission of the upsurge in oil prices into national main aggregates can be modelled using the traditional Cobb-Douglas production function (Bohi, 1991), which predicts that income will decline if energy (input) prices increase because of the higher cost-share of energy in the manufacturing process for the oil-importing economy. As a result, the macroeconomic performance will be worse in episodes of high oil prices (Hamilton, 1983, 2003), with negative impacts on living standards (Considine, 1988). For example, the rise of oil prices jeopardises GDP and increases the consumer price index (CPI) and unemployment in many OECD and emerging countries (Katircioglu et al., 2015; Choi et al., 2018). High oil prices may also weaken the country's competitiveness in exporting raw materials and intermediates (Cavalcanti and Jalles, 2013; Korhonen and Ledyaeva, 2010; Lee and Chiu, 2011), but the effect is unlikely to be symmetric in case of low oil prices (An et al., 2014; Tatom, 1988). Moreover, the local currency exchange rate against the US dollar

[1] According to Scopus database 78.1% of oil- and commodity-related articles in the "economics", "business" and "social sciences" literature were published in 2008 – 2018. Articles in English containing "oil price" or "oil shock" in the title or keywords, or "commodity price" in the title were searched for a period above 100 years, 1911-2018 (search conducted in April 2019). If the category "Energy" is included, this proportion remains constant (78.6%).

often depreciates when oil prices soar up (Lizardo and Mollick, 2010), affecting the terms of trade unfavourably (Kilian et al., 2009; Le and Chang, 2013; Rafiq et al., 2016; Salvatore and Winczewski, 1990).

This paper studies the impact of oil price shocks on the macroeconomic dynamics of Vietnam over the two decades since the 1997 Asian financial crisis. Vietnam's economy is an appealing case of analysis for a twofold reason. First, it has been a 'de-facto' small open economy characterised by exporting crude oil since 1992. Second, it imports up to 70% of its domestic gasoline demand and most of cracked petroleum products despite the operation of its first and only oil refinery Dung Quat since 2009.[2] Vietnam was actually a net oil exporter until 2009 as depicted in Figure 1 (panel B), but rapid growth alongside the steady expansion of exports and private vehicles have made the Vietnamese (VN henceforth) economy more oil-dependent in the recent years. The situation may even worsen in the near future given the current VN oil production, reserve levels and exploration difficulties due to political tensions in Vietnam territorial waters. It should be noted, however, (i) that the VN state-owned enterprises have exerted control in all oil-related activities such as oil exploration, production, and distribution; and (ii) that the government budget systematically benefited both from exporting crude oil and importing petrol products by levying different taxes (for example, export/import tax, excise tax, consumption tax, among others). As a consequence, the transmission of world oil price changes on the prices of domestic goods and services is not so obvious due to a heavily regulated retail gasoline market.

Figure 1 depicts some crucial macroeconomic information for Vietnam's economy. Panels A and C disclose two negative relationships, one between the real oil price and the trade-balance-over-output (TBY), and another one between global real economic activity, the so-called Kilian index, and the real effective exchange rate (REER). Higher world demand may induce better trade competitiveness in Vietnam, as expressed by the lower REER, because it has long exported low added-value goods and agricultural products. Given this information, and the corresponding higher anticipated income, VN

[2] The second oil refinery, Nghi Son, just began operation at the end of 2018, after several years delay.

households may demand and import more sophisticated goods whose prices are also inflated, leading to a worsened trade balance due to the appreciation of Vietnam's Dong (VND hereafter). In turn, real oil prices tend to rise in parallel to increases in world demand. This reinforces the deterioration of the TBY while simultaneously pushing up inflation. Conversely, it seems that changes in oil production do not exhibit significant correlations with VN macroeconomic aggregates.[3]

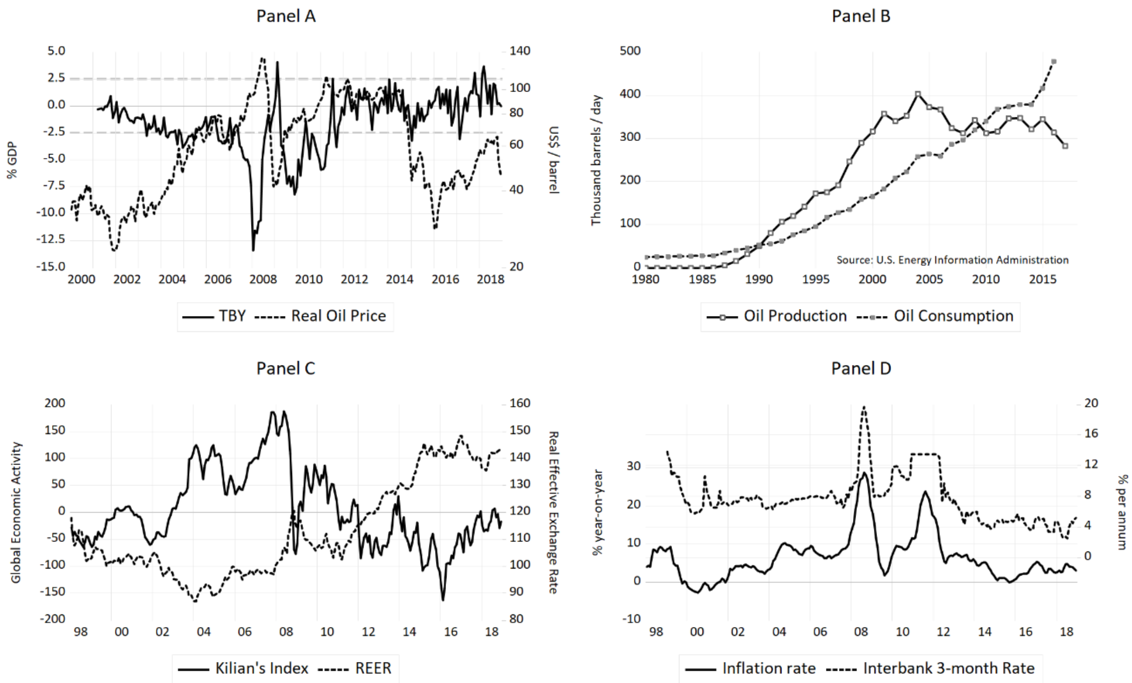


Figure 1: Panel (A): Vietnamese trade-balance-over-output (TBY) and real oil price in 2001 – 2018; Panel (B): Vietnamese oil production versus oil consumption over 1980 – 2017; Panel C: Kilian's (2009) global economic activity index versus VN real effective exchange rate; Panel D: VN inflation rate (year-on-year percentage) and 3-month interbank interest rate (percentage per annum). All series are on a monthly basis except for the VN oil production and consumption which are annual based.

Over the period 2008 – 2012, the VN economy experienced economic turmoil because of the GFC. Its year-on-year inflation rate (blue line in panel D, Figure 1) climbed to peaks of 28.6% and 23.7% in the third quarters of 2008 and 2009, respectively. Meanwhile, the trade balance deteriorated severely, by -10% of GDP on average. This notable macroeconomic instability triggered a battery of monetary

[3] The correlation coefficients between changes in oil production and VN inflation rate, interest rate, real export growth, changes in real (effective) exchange rate, and trade-balance-over-output, respectively, are basically within the range [-0.1, 0.1].

measures that helped the rebalancing of the VN economy in 2013 – 2015. The State Bank of Vietnam (henceforth SBV) raised the short-term interest rate (red line in panel D, Figure 1) to a record high of 19.7% in the first inflationary episode, 2008:01 – 2009:03. In the second one, instead, the SBV maintained a flat interest rate of 13.5% over sixteen months, 2010:12 – 2012:03, accompanying a steady decline in headline inflation to below 5% in 2014:04. Moreover, the sizable current account deficits over four years 2007 – 2011 stimulated the SBV to devalue the VND by 30% from 2008:04 to 2011:04. These policies actually brought about the recovery of the trade deficit in years later.

These developments provide a first rough evidence that global demand shocks may be harmful for the VN economy and thus require an appropriate design of the monetary policy. Nevertheless, the existing macroeconomic literature on Vietnam's economy is relatively scarce. Pham et al. (2019) put forward a small open economy real business cycle model showing that VN trade-balance-over-output has been sensitive to international shocks in the past three decades, 1986 – 2015. No explicit attention is paid, however, to the role of oil price shocks. Trang et al. (2017), using a VAR model with Cholesky decomposition, claim that oil price shocks inflate domestic prices while diminishing government budgets. Anwar and Nguyen (2018) consider oil prices in their two-block SVAR study of the transmission of oil price rises to the VN monetary policy between 1995 and 2010. Their results imply that "the monetary policy of Vietnam was quite sensitive and vulnerable to fluctuations in the world price of oil". Still, none of these VAR analyses is able to disentangle structural oil supply and demand shocks in Kilian's (2009) sense, as the research thereafter has done (for example, Ahmed and Wadud, 2011; Cuñado et al., 2015; Iwaisako and Nakata, 2017; Lorusso and Pieroni, 2018, among many others).[4] Bhattacharya (2014) and Narayan (2013) are rare papers from international scholars considering Vietnam. Bhattacharya examined the movements of inflation in the short-run and connected the effectiveness of the monetary policy to changes in the nominal effective exchange rate, and to the

[4] In addition, their results may be liable to some inaccuracy from the linear extrapolation used to convert annual data to quarterly frequency observations in order to be able to expand backwards their sample period of analysis.

credit expansion policy over the period 2004 – 2012. In turn, Narayan suggested that hikes in oil prices tend to depreciate the VND in the near future.

The drawback of these studies is that the triangle inflation-oil prices-monetary policy has been examined by components and not holistically. Hence, departing from the foregoing literature, I fill a void in the literature and comprehensively assess the pass-through of oil price fluctuations into the VN economy in two dimensions. First, I question how different oil price shocks affect the VN inflation rate, the real (effective) exchange rate, and foreign trade. Then, I analyse the responses to these shocks of VN monetary instruments, which comprise the three-month Interbank interest rate and the nominal exchange rate of VND against the U.S. Dollar. In all the analysis, we use monthly data covering the period 1998:01-2018:12.

The methodological contribution is threefold. First, although I depart from Kilian's (2009) framework of analysis, I work with an extended SVAR model with expanded foreign and domestic blocks. Second, I deal, as a consequence, with an over identified SVAR model whose predictive reliability requires accepting the over-identifying restrictions. Third, in using monthly data I am able to focus on a shorter sample period without falling short of degrees of freedom. This allows me to avoid any noise from a potential need of a low-to-high frequency data conversion.

The paper is structured as follows. Section 2 presents a summary of recent related literature on major Asian economies. Section 3 outlines the SVAR methodology, model specifications and data description. Sections 4 and 5 deal with the empirical results and variance analysis. Section 6 concludes.

3.2 Recent evidence on the macroeconomic effects of oil price shocks in Asia

A number of studies have focused on the macroeconomic effects of oil price and global demand shocks in several major Asian countries – namely, Japan, China, India, and other ASEAN economies. With the exception of Malaysia, these are all twenty-first century net oil importers.

Although not unanimous, the evidence for Japan points to harmful macroeconomic effects of oil price shocks. For the shocks in the mid-1970s and 1979-1980, Lee et al. (2001) showed not only direct negative impacts, but also indirect effects caused by the subsequent tightening of the monetary policy that resulted in a higher money call rate. Zhang (2008) studied the non-linearity of these impacts and disclosed asymmetric effects on the Japanese macroeconomic performance with oil price increases causing larger negative impacts on income growth than the positive ones from equivalent price cuts. Fukunaga et al. (2011) used Kilian's (2009) framework to show negligible effects of oil price shocks on the Japanese economy, but Jiménez-Rodríguez and Sánchez (2012) found evidence of the reverse for the early 1980s. Against the feeling that Japanese industrial production seems immune to oil price increases these days, the most recent evidence claims that global demand shocks and oil market-specific shocks not only are relevant but should be considered as chief stimulants of dynamism in the Japanese aggregates (Rahman and Zoundi, 2018). Notably, Iwaisako and Nakata (2017) assert that positive non-fundamental oil price shocks supported Japanese exports in the 2000s.

China, the second biggest economy and oil consumer in the world, endogenously affects the world oil price due to its enormous size and export expansion strategy (Faria et al., 2009). Kim et al. (2017) use different SVAR estimation techniques to find evidence of a price stabilisation policy of Chinese policymakers to counteract the inflationary effect of oil price shocks between 2001 and 2014. Nevertheless, the cumulative impact of China's broad money supply is responsible for the strong recovery of oil prices during 2009, as noted by Ratti and Vespignani (2013). Taking a microeconomic-based approach, Zhao et al. (2016) propose a calibrated open economy DSGE model proving that oil supply shocks driven by non-political events, aggregate shocks to the demand of industrial commodities, and oil-specific demand shocks have long-term impacts on China's output and inflation fluctuations. Interestingly, Osorio and Unsal (2013) find that inflation in China has spillover effects on economies in the ASEAN community and India owing to their huge demand of commodity goods.

Likewise, India – the second most populous country in the world – imports as much as 80% of its fuel demand, thus rendering its economy exposed to oil and commodity price shocks. Holtemöller and

Mallick (2016) show that Indian consumer prices are highly sensitive to inflationary supply shocks (oil price, food price, and other cost-pushes), but question policy measures such as raising interest rates because of the harm that a monetary contraction would cause on output growth. It has also been observed that the oil price does not Granger cause the USD/INR exchange rate (Inumula and Solanki, 2017), implying that a policy of stabilising and strengthening the Indian Rupee would contribute to brake the pass-through of global shocks on domestic inflation.

Finally, there are several studies on the ASEAN-5 countries, namely, Malaysia, Indonesia, Philippines, Singapore, and Thailand. The battery co-integration tests by Kisswani (2016) reports a two-way relationship between real oil prices and real exchange rates in the long run, but Basnet and Upadhyaya (2015) claim that oil price shocks have only temporary effects on the ASEAN-5 markets. In particular, they show that inflation reflects oil price rises in all countries in the first two quarters after the shock, but restrict the positive impact of such rises on real output to Indonesia, Malaysia, and Singapore. Sultonov (2017) studies the negative side of the oil price shocks from 2014 for the ASEAN-5 countries. He shows that crude oil price statistically affects exchange rates, and that the oil price volatility spills over from the crude oil market to the foreign exchange market.

3.3 Empirical methodology and data description

3.3.1 Methodology

The small open economy SVAR model used in Kilian's (2009) framework is typically set up in two blocks with a foreign (or exogeneous) block consisting of several variables accounting for oil price and/or other international shocks. For a small open economy, the second block includes domestically endogenous variables supposed to have negligible influence on their foreign block counterparts. Specifically, the two-block SVAR has a form[5]

[5] I have suppressed deterministic terms to simplify the exposition.

$$\underbrace{\begin{bmatrix} A_0^{11} A_0^{12} \\ A_0^{21} A_0^{22} \end{bmatrix}}_{A_0} \underbrace{\begin{bmatrix} y_t^f \\ y_t^d \end{bmatrix}}_{y_t} = \sum_{p=1}^{\mathcal{P}} \left(\underbrace{\begin{bmatrix} A_p^{11} & A_p^{12} \\ A_p^{21} & A_p^{22} \end{bmatrix}}_{A_p} \underbrace{\begin{bmatrix} y_{t-p}^f \\ y_{t-p}^d \end{bmatrix}}_{y_{t-p}} \right) + \underbrace{\begin{bmatrix} B_0^{11} B_0^{12} \\ B_0^{12} B_0^{22} \end{bmatrix}}_{B_0} \underbrace{\begin{bmatrix} u_t^f \\ u_t^d \end{bmatrix}}_{u_t} \tag{1}$$

where y_t^f and $y_{,t}^d$ are vectors of k_f foreign and k_d domestic variables, respectively; $p = 1 \dots \mathcal{P}$ denotes lagged index of the time series; A_0, $A_{1..\mathcal{P}}$ and B_0 are structural coefficient matrices that cannot be directly estimated; and u_t is therefore the so-called vector of structural residuals assumed to be independently identically distributed such that $E(u_t u_t') = I_{K \equiv k_f + k_d}$.

Pre-multiplying both sides of equation (1) by A_0^{-1} (assuming A_0 is invertible), one obtains the unrestricted reduced form as

$$y_t = \sum_{p=1}^{\mathcal{P}} A_0^{-1} A_p y_{t-p} + A_0^{-1} B_0 u_t$$

which can be also rewritten in the more compact form $B(L)y_t = \epsilon_t$, with the lag polynomial $B(L) = I + B_1 L + \cdots + B_p L^p$ and the vector of reduced residuals $\epsilon_t = A_0^{-1} B_0 u_t$, so that $E(\epsilon_t \epsilon_t') = \Omega_\epsilon$ is diagonal, i.e. $\epsilon_t \overset{iid}{\sim} N(0, \Omega_\epsilon)$. Furthermore, block exogeneity due to the small open economy assumption postulates that all elements of matrix B_p^{12} are restricted to zeros. This results in an unbalanced VAR model which behaves similarly to seemingly unrelated regression (SUR) models.

Explicitly, the relationship between ϵ_t and u_t in the sense of Amisano and Giannini (1997) is

$$A_0 \epsilon_t = B_0 u_t \tag{2}$$

To recover either A_0 or B_0 or both from the consistent estimate of ϵ_t, some restrictions need to be imposed on elements of A_0 and/or B_0 because of the symmetry of $\widehat{\Omega}_\epsilon$. For example, if $u_t \overset{iid}{\sim} N(0, I_K)$ and the diagonal elements of A_0 are normalised to unity, then the just-identified identification requires a total of $K^2 + K(K-1)/2$ restrictions on both A_0 and B_0.

I consider several SVAR specifications. The structure, common to the proposed models in the literature, is based on three standard variables in the exogenous block that emerged from Kilian's study (2009) – oil production (Oilpd), a global economic activity index (Globix), and the real oil price (Oilpr). Restrictions on the A_0 and/or B_0's elements are a fundamental matter in any SVAR analysis. Following Kilian (2009), the A_0^{11} block should have a recursive structure like

$$A_0^{11} \equiv \begin{bmatrix} a_{11} & a_{12} & a_{13} \\ a_{21} & a_{22} & a_{23} \\ a_{31} & a_{32} & a_{33} \end{bmatrix} \equiv \begin{bmatrix} * & 0 & 0 \\ * & * & 0 \\ * & * & * \end{bmatrix} \tag{3}$$

where the asterisks (*) represent free parameters to be estimated. The ordering (oil production, global economic activity, real oil price) implies that contemporaneous impacts on the oil price may originate from oil supply disruption, oil consumption demand, or oil-specific market demand such as precautionary or non-fundamental shocks. Elements of the A_0^{21} block are free in the baseline setting, but I do impose additional zero-restrictions on the A_0^{21} matrix in the augmented models, as will become clear below. In regard to the domestic block, A_0^{22}, the recursive structure is again applied to the small set of two variables, say the VN consumer price index (CPI) and the bilateral VND/USD real exchange rate (RER), so that the normalised matrices A_0 and B_0 of the baseline model are

$$A_0^{base} \equiv \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ * & 1 & 0 & 0 & 0 \\ * & * & 1 & 0 & 0 \\ * & * & * & 1 & 0 \\ * & * & * & * & 1 \end{bmatrix} \quad B_0^{base} \equiv \begin{bmatrix} * & 0 & 0 & 0 & 0 \\ 0 & * & 0 & 0 & 0 \\ 0 & 0 & * & 0 & 0 \\ 0 & 0 & 0 & * & 0 \\ 0 & 0 & 0 & 0 & * \end{bmatrix} \tag{4}$$

Abstracting from the matrices structure just described, the baseline model by Kilian (2009) can be summarized as[6]

Baseline model: $$\left(\overbrace{\text{DOilpd} \quad \text{Globix} \quad \text{LnOilpr}}^{\textit{Foreign block}} \quad \overbrace{\text{DLnCPI} \quad \text{DLnRER}}^{\textit{Domestic block}} \right)$$

[6] Prefixes D, Ln, and DLn refer respectively to logarithm, first-difference, and log-difference operators, where the last is $DLn(y_t) = \log(y_t) - \log\,(y_{t-1})$.

3.3.2 Augmented models

I take a step forward with respect to Kilian's (2009) model and consider the possibility of risk-premium shocks reflected in interest rate changes. I thus add the U.S. Federal Fund rate (Fedfunds) to the foreign block such that the effects of world risk premium shocks can be taken into account in the analysis.

In addition, the domestic block is augmented by taking into account real exports (REXP). This yields:

Model I $$\left(\overbrace{\text{DOilpd}\quad \text{Globix}\quad \text{LnOilpr}\quad \textbf{Fedfunds}}^{\textit{Foreign block}}\quad \overbrace{\text{DLnCPI}\quad \text{DLnRER}\quad \textbf{LnREXP}}^{\textit{Domestic block}}\right)$$

In connection to the real exchange rate, the addition of real exports allows us to model more precisely the impact of the oil price and risk-premium shocks on the external demand of VN goods and services. This can be further refined, however, by considering a wider approach to Vietnam's international competitiveness. In this widen setting, the real exchange rate is substituted by the real effective exchange rate (REER) while real exports are replaced by the trade balance (TBY). This delivers:

Model II $$\left(\overbrace{\text{DOilpd}\quad \text{Globix}\quad \text{LnOilpr}\quad \text{Fedfunds}}^{\textit{Foreign block}}\quad \overbrace{\text{DLnCPI}\quad \textbf{DLnREER}\quad \textbf{TBY}}^{\textit{Domestic block}}\right)$$

Finally, I are also interested in assessing how the SBV has dealt with the impact of oil price and risk-premium shocks in order to maintain inflation under control. This is certainly crucial for a small open economy such as VN. Consequently, I replace the real effective exchange rate and the trade balance by the interbank 3-month interest rate (Rate3M) and the nominal exchange rate Dong/$ (FX):

Model III $$\left(\overbrace{\text{DOilpd}\quad \text{Globix}\quad \text{LnOilpr}\quad \text{Fedfunds}}^{\textit{Foreign block}}\quad \overbrace{\text{DLnCPI}\quad \textbf{Rate3M}\quad \textbf{LnFX}}^{\textit{Domestic block}}\right)$$

In this way, the domestic block is made of the main variables of interest of a central bank so that its monetary policy response to oil price and risk-premium shocks can be properly assessed.

The reasons for augmenting Kilian's (2009) set up by recruiting different proxies for Vietnam's external trade and monetary policy can be found in Pham *et al.* (2019). They showed that country risk premium shocks account for about one-fifth of Vietnam's output growth variability, on the one hand; while, on the other, have a large explanatory power of the variations in the trade-balance-over-output ratio over the past two decades.

The first three columns of Table 1 summarise my model's variable choices. In the three models I, II and III, I restrict (i) the oil supply shocks to have no contemporaneous effects on Vietnam's domestic variables (i.e., in the month impact); and (ii) the risk-premium shocks to have no contemporaneous effects on Vietnam's inflation either. This implies that my three SVAR specifications are over-identified and should then pass the Likelihood-Ratio test for over-identifying restrictions. This is indeed the case, as shown in the last two columns of Table 1.

Model	Foreign Block	Domestic Block	VAR(p)	Largest Root	Over-identification Test
Baseline	Kilian (2009)	DLnCPI, DLnRER	6	0.9704	-
A1	Kilian (2009), U.S. Fedfunds	DLnCPI, DLnRER, DLnREXP	3	0. 9962	0.699
A2	Kilian (2009), U.S. Fedfunds	DLnCPI, DLnREER, TBY	3	0.9552	0.717
B	Kilian (2009), U.S. Fedfunds	DLnCPI, Rate3M, LnFX	3	0.9914	0.634

Notes: Kilian's (2009) variables are oil production (OilProd), the global economic activity index (Globix), and the real oil price (Oilpr). Prefixes D, Ln, and DLn refer to logarithm, first-difference, and log-difference operators, with $DLn(y_t) = \log(y_t) - \log(y_{t-1})$; FX, RER, and REER are, respectively, the nominal VND/USD exchange rate, the bilateral VND/USD real exchange rate, and the real effective exchange rate; REXP denotes real exports; TBY represents trade-balance-over-output; Rate3M is the three-month Interbank interest rate; and CPI denotes the consumer price index. Zero-restrictions on A_0^{21} matrix of models I, II, and III are as $\begin{bmatrix} 0 & * & * & 0 \\ 0 & * & * & * \\ 0 & * & * & * \end{bmatrix}$. Numbers in the last columns are p-values.

Table 1: Model specifications

According to Akaike's information criteria and the rule-of-thumb in VAR order selection, I pick up the suitable order, VAR(p), of 3 for models I, II and III. In contrast, the baseline model is intentionally estimated with a six-month lag to entirely capture the effects of oil price shocks on the VN CPI and

RER dynamics.[7] The fifth column reports that all VAR models are stable since their largest inverse roots of the AR characteristic polynomial lie inside the unit circle.

3.3.3 Data description

Variables	Oilprod	Globix	Oilpr	DLnCPI	DLnFX	DLnRER	REER	Rate3M	TBY	DLnREXP
Mean	73945	5.76	59.18	0.11	0.24	-0.08	112.97	7.73	-1.62	1.05
Median	73931	-9.07	54.42	0.09	0.05	-0.10	106.77	7.37	-1.13	0.99
Max	84225	188.00	132.97	0.79	8.77	6.94	148.36	19.69	4.06	45.80
Min	64307	-163.00	13.98	-0.28	-0.54	-3.09	87.01	2.47	-13.37	-24.99
Std.Dev	4908	71.35	27.49	0.15	0.87	1.01	17.54	3.14	2.72	8.63
Obs	252	252	252	252	252	252	252	240	214	252

Notes: Oilprod (thousand barrels / day), Oilpr (US$ / barrel), DLnCPI (% m-o-m), DLnFX (% m-o-m), DLnRER (% m-o-m), Rate3M (% pa), TBY (% GDP), DLnREXP (% m-o-m).

Table 2: Descriptive statistics, 1998:01 – 2018:12

It should be stressed that the constructed dataset is on a monthly basis. This is important because it allows my analysis to focus on a recent period, 1999-2018, without running out of degrees of freedom in the estimation. In addition, it is important to remark that no yearly-quarterly data interpolation has been needed, as it is often the case in studies on close emerging economies. I obtained the VN CPI, the nominal exchange rate of the VND against the U.S. Dollar (FX), and export (EXP) data from the Vietnamese General Statistical Office (GSO) via Datastream. The interbank 3-month interest rate (Rate3M) was taken from the SBV. Real effective exchange rate (REER) running up to 2018:12 was extracted from Darvas (2012), since the IMF does not officially provide this series for Vietnam. Kilian's studies (2009, 2019) supplied the corrected global economic activity index (Globix), which is a proxy for the world demand of goods. It should be noted that Hamilton (2018) criticises Kilian's (2009) index for failing to account for global consumption demand. Nonetheless, Kilian (2019) adds a corrigendum justifying that the gap between the old and the new index is highly unlikely to bias any related studies.

The U.S. Energy Information Administration (EIA) provided oil production (Oilprod) data in terms of a monthly average in thousands of barrels per day. For real oil price (Oilpr), I compute the average

[7] See Table A.1 in the appendix for further details.

of West Texas Intermediate (WTI) and Brent oil prices, also obtained from EIA, after adjusting by the U.S. consumer price index. All data series cover the timespan 1998:01 – 2018:12 except for the Rate3M series starting in 1999:01. Trading-balance-to-output (TBY) is calculated as $\frac{RealExports_t - RealImports_t}{Output_t}$. Since output data for Vietnam in terms of US Dollar currency[8] is only available at quarterly frequency from 2001:01, the Chow-Lin interpolation method is used for the low-to-high frequency conversion. Figure A.1 in the Appendix, shows the dynamics of the considered time series, whereas Tables 2 and A.2 in the Appendix report, respectively, descriptive statistics and the corresponding correlation matrix.

3.4 Empirical results and discussion

In this section, I present and then discuss the estimated impulse-response functions (IRFs) of four models, from which the transmission of foreign shocks to the VN economy is quantitatively assessed over the full sample 1998 – 2012 period.[9] These are Kilian's (2009) baseline specification and its subsequent expanded versions – Models I, II, and III.

I interpret a positive (negative) shock to any variable in the foreign block as causing increases (decreases) either in oil production, the global demand of goods, or in the level of speculation in the oil market. Similarly, a rise (fall) in the U.S. Federal Fund rate tightens (loosens) the monetary policy pushing up the cost of borrowing. Therefore, in the rest of this paper I use positive or negative shock interchangeably depending upon the context, but the IRFs are always computed and plotted as positive impacts.

3.4.1. The baseline model

Figure 2 shows the baseline IRFs, with one-standard-deviation error bands in red dotted lines. The first row of panel A shows that oil supply surprises only have a short-lived impact on both domestic

[8] Downloaded from CEIC data provider.

[9] Beyond the direct estimation of equation (1), I have also estimated the system including dummy variables that control for the possible inflection point experienced by the VN economy in 2009, when it became a net oil importer. All the reported results in this paper hold in the presence of these dummy variables, which take value 1 in 2009-2018 and zero otherwise. I interpret this robustness as evidence that 2009 did not cause a structural break in the economic relationships under scrutiny. These additional results are available upon request.

endogenous variables (CPI, RER), as their responses vanish within two quarters. On the contrary, oil demand shocks induced by global economic activity or speculation/innovation induce highly persistent responses in the VN inflation and real exchange rates. Specifically, the responses of the inflation rate to oil demand and oil-specific demand shocks reach their peaks in two and eight months, respectively, and then asynchronously revert to equilibrium. However, the recovery of the inflation rate under non-fundamental oil price shocks is much faster than under shocks influenced by global demand, with the former clearly dying down within a one-year horizon.

From the second row of panel B, it is observed that one standard deviation of an oil demand shock (about 18 index points) raises the VN CPI by an annualised percentage of 1.9%, 3.6%, and 4.8% over one-, two- and three-year horizons, respectively. Correspondingly, the RER falls greatly, by 6.5, 13.1%, and 17.4% per annum in the same three horizons, respectively, leading to a strong appreciation of the VND against the U.S. Dollar.

Hence, Kilian’s (2009) baseline model reveals that, even though supply-side shocks are innocuous, the VN CPI and RER are fairly responsive to both types of oil demand shock.

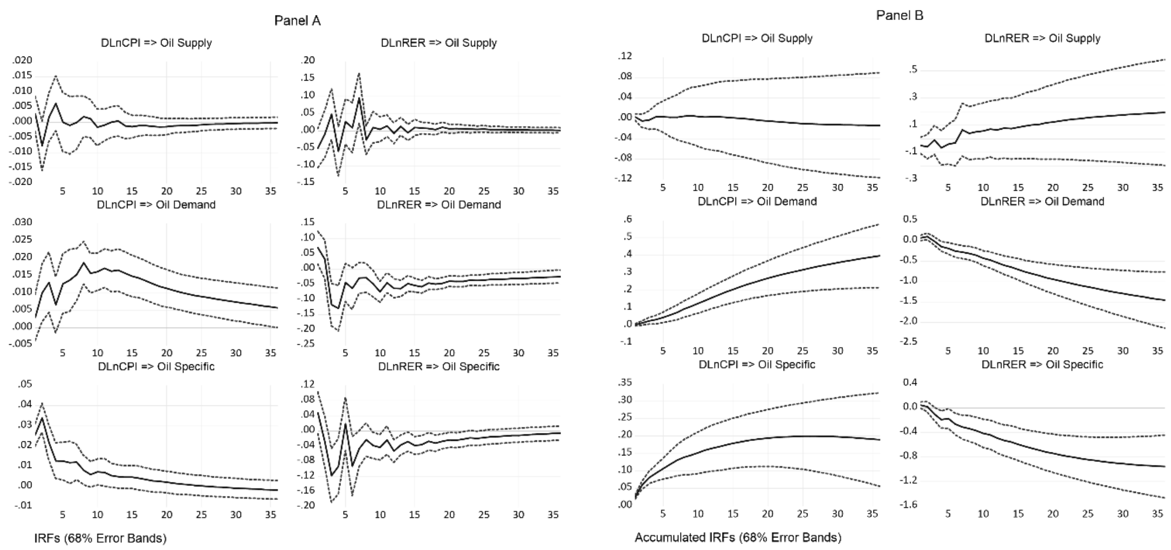


Figure 2: The baseline model (DOilpd, Globix, LnOilpr, DLnCPI, DLnRER). Panel A: Recursive impulse - response functions (IRFs) of VN CPI and RER to one standard deviation structural shocks, defined as oil supply, oil demand, and oil market-specific demand; Panel B: accumulated IRFs of the corresponding ones in panel A. The red dotted lines are the 68% error bands.

3.4.2 Models I and II

The estimated IRFs for models I and II yield particular insights into the dynamics of VN external trade variables. They reveal, first of all, that VN export and trade balances were immune to any surprising change in global oil production. Although, this tends to confirm the innocuous effects of supply-side oil price shocks, Figure 3 clarifies that oil production increases have marginally significant and short-lived impacts on the time-paths of VN (real) effective exchange rates and domestic prices.

The IRFs of RER and REER are quite similar under the effects of oil supply disruptions, but they behave in the opposite way when hit by oil demand shocks. To be precise, a positive global demand oil shock raises the VN REER significantly and persistently for almost a year after the impact. In contrast, an oil-specific demand shock initially decreases REER growth rate by about 0.3 percent per month, but rapidly returns to its equilibrium before climbing to a positive peak of 0.15% in the fourth month. This indicates that the VN economy starts losing its relative competitive advantage in just one quarter after a speculative oil price shock hits the economy. Conversely, the resulting strengthened values of the VND result in cheaper foreign goods for the VN households in a year or more and compromise the trade balance. On this account, note that the responses of TBY depicted in the third row of Figure 3 – Panel B show that oil price increases actually impair the VN trade balance in the short-to-medium run, and of course the current account, even though exports also improve.

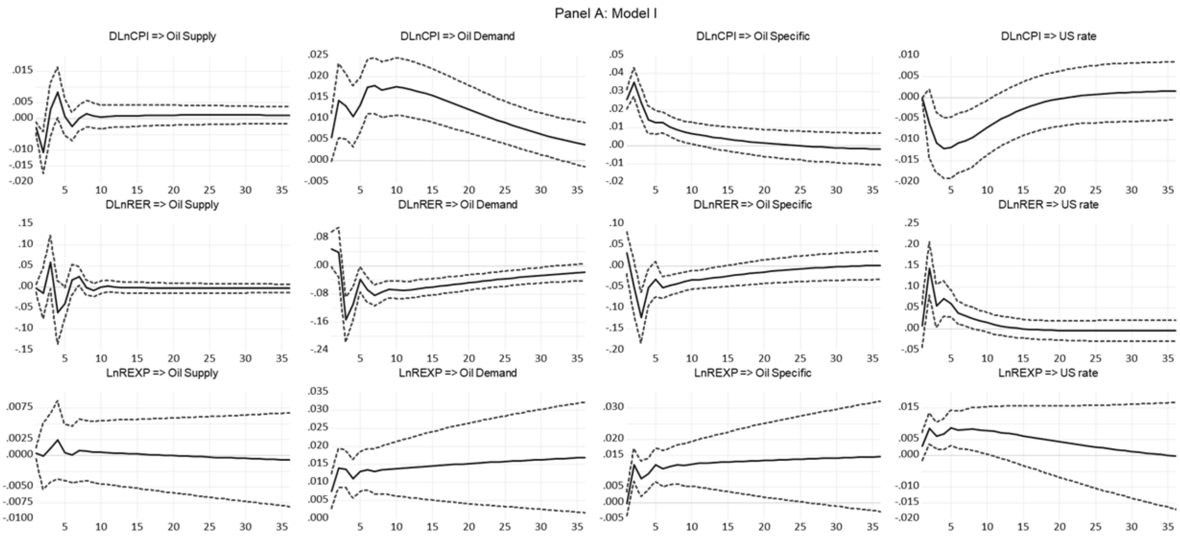


Figure 3: The IRFs of model I (DOilpd, Globix, LnOilpr, Fedfunds, DLnCPI, DLnRER, LnREXP). The dotted lines are the 68% error bands.

Equally important, Kilian's (2009) expanded models uncover that the U.S. Federal Fund rate has remarkable effects on the VN economy because of the strong tie between the two currencies, as pointed out by Anwar and Nguyen (2018). The fourth column of Figure 3 – Panel A shows that the VN inflation rate significantly improves (negative adjustment) in two quarters or more, if there is a hike in the U.S. policy rate. However, it slowly recovers subsequently and then reverts to the zero-line in the mid-term. Note that the response of inflation to the Federal Funds rate in model I is stronger than its counterpart in model II. In the first case, I evaluate the reaction of the RER, which is highly conditioned by VND pegging to the US dollar.[10] In the second case (Figure 3 – Panel B), in contrast, the response is much less persistent in consistence to the multilateral setting captured by the REER. The REER reflects an enlarged system of trade relationships in which adjustments take place quicker than in the bilateral setting depicted by Model I.

[10] Note that $RER = FX\frac{CPI^f}{CPI^d}$. This explains the systematic inverse responses of the RER and domestic inflation.

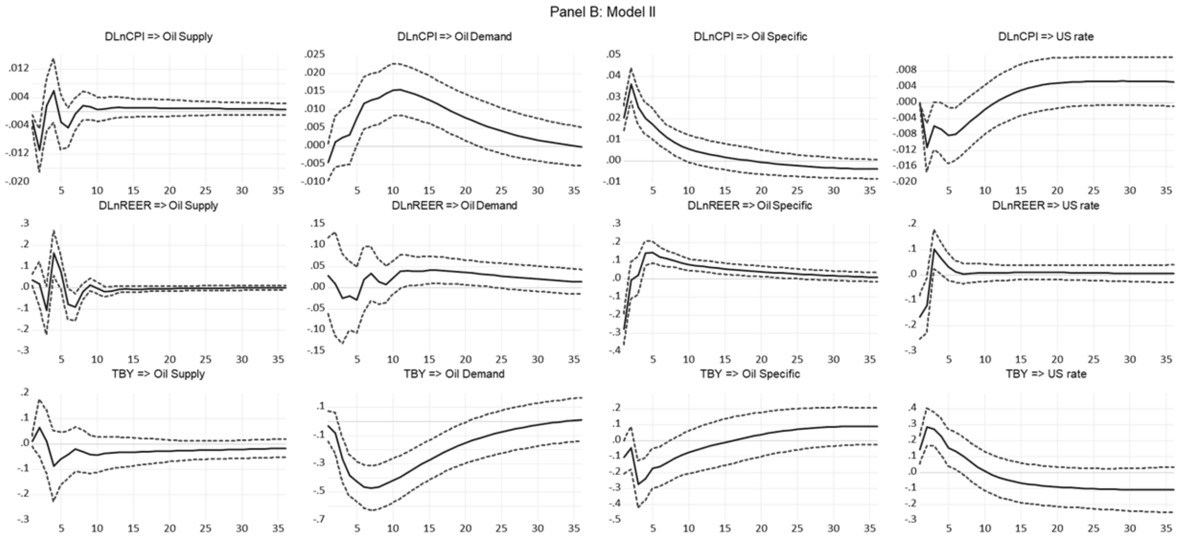


Figure 3 (continue): The IRFs of model II (DOilpd, Globix, LnOilpr, Fedfunds, DLnCPI, DLnREER, TBY). Dotted lines are the 68% error bands.

By the same token, the VN trade balance significantly positively reacts to an increase in the exogenous risk premium in five months after the initial impact, since that type of shock depreciates the VND in the following six months (see the last column of Figure 3 – Panel A), after reaching the peak response of 0.15% at the second month and afterwards diminishing to the negative side but the latter is statistically insignificant.

3.4.3. Model III

Figure 4 presents the adjustments of Vietnam's monetary policy – domestic interest rates and nominal VND/USD exchange rates –, in response to oil prices and international risk premium innovations. Global demand shocks indirectly provoke the rises in VN interest rates, with a lag of two months in response to the consumption price rally. The peak interest rate response occurs in the sixteenth month, nearly ten months after the peak of changes in the inflation rate, implying that the effects of oil demand shocks on the VN economy are prolonged. When the Vietnam's economy is hit by an 8.3% increase in crude oil prices, which is estimated to be one standard deviation of oil market-specific demand shock, the short-term interest rate climbs dramatically to a peak of 0.5% at four months after the response of inflation rate attains its largest magnitude in the second period. Likewise, the nominal VND/USD exchange rate only commences to depreciate significantly after a five-month lag after an oil-specific demand shock. This reflects the strong connection between the nominal exchange rate and the short-term interest rate.

The detachment between the responses of inflation, on one side, and domestic interest rates and nominal VND/USD exchange rates, on the other, during the first two months after a risk-premium shock, can be interpreted as a manifestation of the "price puzzle" (see Castelnuovo and Surico, 2006). My reading of these results is that the VN authorities have implemented the monetary policy in a cautious fashion. In particular, the interest rate policy seems to have been too passive regarding international demand shocks and/or not strong enough to counteract the rapid inflation rate growth resulting from such innovations. This interpretation is consistent with Bhattacharya's (2014) finding of persistently larger rates of inflation in Vietnam than in its neighbouring emerging economies.

Supporting this argument, Figure A.3 adds complementary information showing model's III IRFs of the three variables in the domestic block in response to their own innovations. The autonomous responses of the inflation rate vanish in two quarters, in clear contrast to the interest rate response to this same shock, which steadily rises to reach its peak over the same six-month period and only converges back to zero over a fifteen-month period. Contrariwise, the inflation rate responds weakly to a rise in the interest rate, which is in fact consistent with the findings of Bhattacharya (2014). In the case of an exchange rate shock, the interest rate tends to react strongly in the first half-year (note that its initial response of 0.1% equates 90% of the peak), while the inflation rate only responds significantly in three consecutive months after the initial impact. Finally, the nominal VND/USD exchange rate adjusts upwards in the aftermath of a positive shock to either the inflation or the interest rate, but soon stabilises over the short- to medium-term.

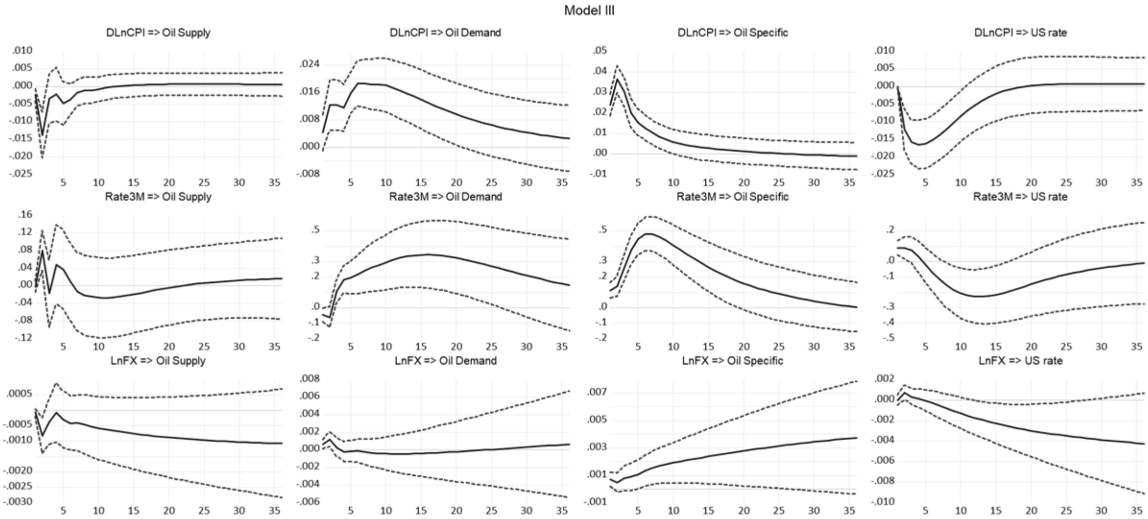


Figure 4: The IRFs of model III (DOilpd, Globix, LnOilpr, Fedfunds, DLnCPI, Rate3M, LnFX). Dotted lines are the 68% error bands.

3.5 Variance analysis

3.5.1 Forecasting variance decomposition

The above analysis tells us how VN macroeconomic indicators behave in response to foreign surprises as well as their own innovations, but it cannot explain how much of their variation is explained by those shocks for either forecasting or historical analysis. Table 3 summarises the variance contribution of each structural shock to each domestic endogenous variable across Models I, II and III (I only show the first twenty-four months, as they are fairly stable afterward). Note that, for the inflation rate, I average the contributions of four foreign shocks because their model-specific values are remarkably similar, and it is not worth presenting them separately. Figure A.4 in the Appendix depicts these variance shares for each endogenous series in detail.

At the first six-month horizon, I find that domestic variables are predominantly affected by their own shocks. It is worth highlighting, however, that among the foreign block shocks, precautionary oil demand shocks have the strongest explanatory power, accounting for around one-fifth of the variation in the VN inflation rate and three-month interest rate.

In the middle run and beyond, from 12 to 24 months, autonomous shocks explain about 50% of the fluctuations in the trade balance and inflation rate, 26% of the fluctuations in the three-month interest rate, and between 66% and 80% of the fluctuations in the other domestic variables. Compared to other structural shocks, global demand shocks seem to be the most important macroeconomic drivers, as they account for about one-third of the fluctuation in TBY, one-fifth in inflation, one quarter in interest rates, and roughly one-tenth in exports and the RER. However, they make only a negligible contribution to nominal exchange rates and the REER. In sharp contrast, oil production disruption has an extremely low explanatory power of these variances.

Additionally, oil market-specific shocks are highly likely to play a crucial role in explaining the long-run variance in interest and inflation rates. To be precise, its contribution to interest rates approximately equals the size of the own shock (close to one-fourth), while its contribution to inflation rates is around 17%. The results also show that an oil-specific shock accounts for 8%, 10%, and 8% of the long-run variance of the nominal exchange rate, exports, and REER, respectively. Regarding the risk premium shock induced by the U.S. monetary policy, I find that, after 24 months, it has only a small impact on all domestic variables, explaining between 2.5% and 7.3% of their long-run volatilities.

Summing up, my findings indicate that apart from the prime power of own shocks in the short run, both types of oil demand shock play an essential role in explaining the long-run variations of several VN macroeconomic indicators. Oil demand shocks most particularly affect the trade balance, whereas the three-month interest rate is strongly influenced by oil-specific demand innovations. In addition, both types of shock are equally important for the inflation rate. Lastly, it is shown that inflation moderately affects interest rates, explaining 18% of its long-run variance, while the reverse influence is insignificant.

Variable	Step	Oil Supply	Oil Demand	Oil Specific	US Rate	DLnCPI	Rate3M	LnFX	LnREXP	TBY	DLnRER	DLnREER
DLnCPI	1	0.09	0.30	7.46	0.00	92.15	0.00	0.00	0.00	0.00	0.00	0.00
	6	1.29	4.39	19.82	4.01	61.80	1.20	1.70	0.68	10.41	3.21	8.90
	12	1.14	12.10	18.69	5.09	53.80	1.96	1.47	0.60	12.79	2.98	7.78
	18	1.09	16.65	17.80	5.01	50.70	1.98	1.41	0.57	12.13	2.80	7.36
	24	1.09	18.47	17.39	5.14	49.38	1.94	1.38	0.56	11.85	2.71	7.19
Rate3M	1	0.00	0.42	2.56	1.56	1.27	94.18	0.00				
	6	0.29	3.73	18.36	0.76	18.87	56.32	1.68				
	12	0.22	9.74	26.78	3.73	22.13	36.09	1.31				
	18	0.20	16.77	26.82	6.26	19.34	29.49	1.11				
	24	0.18	21.81	25.67	6.82	17.57	26.78	1.17				
LnFX	1	0.01	0.73	0.87	0.00	3.03	1.26	94.10				
	6	0.33	0.62	1.54	0.22	2.73	6.80	87.75				
	12	0.43	0.44	3.82	1.43	4.65	7.22	82.02				
	18	0.62	0.38	5.91	4.18	5.25	6.74	76.93				
	24	0.81	0.29	7.96	7.32	5.24	6.21	72.17				
LnREXP	1	0.00	1.21	0.00	0.17	0.14			98.17		0.30	
	6	0.07	8.07	4.81	2.69	1.57			82.28		0.51	
	12	0.05	10.20	7.15	3.40	1.56			77.26		0.39	
	18	0.04	11.68	8.58	3.16	1.58			74.63		0.35	
	24	0.03	12.92	9.63	2.69	1.59			72.83		0.32	
TBY	1	0.01	0.06	0.52	1.05	2.07				93.64		2.66
	6	0.34	12.86	4.11	5.59	2.70				73.04		1.35
	12	0.38	27.92	4.10	4.70	2.73				58.93		1.24
	18	0.44	32.46	3.85	4.69	2.57				54.82		1.17
	24	0.48	33.26	3.96	5.20	2.51				53.45		1.14
DLnRER	1	0.00	0.30	0.13	0.01	4.17			0.00		95.39	
	6	0.88	4.01	2.18	3.25	12.81			3.75		73.12	
	12	0.93	7.11	3.09	3.45	12.62			3.58		69.22	
	18	0.91	8.84	3.31	3.38	12.36			3.52		67.69	
	24	0.90	9.81	3.35	3.34	12.21			3.49		66.89	
DLnREER	1	0.07	0.04	3.39	1.62	8.84				0.00		86.03
	6	2.66	0.18	5.01	2.61	8.69				1.90		78.94
	12	3.03	0.48	7.00	2.56	8.50				2.07		76.35
	18	3.00	0.95	7.71	2.57	8.39				2.06		75.31
	24	2.99	1.27	8.00	2.58	8.33				2.05		74.79

3.5.2 Historical variance decomposition

The preceding subsection answered the question of the variance contribution in a forecasting context. I now turn to explore and contrast how these shocks explain the dynamics of Vietnam's inflation, interest rates and trade balance in the following selected periods of interest: the two inflationary episodes of 2007:06 – 2009:12 and 2010:01 – 2012:06, and the years of deteriorated trade-balance-over-output in 2007:01 – 2011:12. Figures 5, 6 and 7 present the contributions of the structural oil price and risk-premium shocks to their trajectories (oil supply shocks are omitted due to their extremely limited impacts on the domestic variables, as also found in Table 3). The bars represent actual data, while the black solid lines depict the benchmark projection implied by the model in the absence of shocks. In turn, the different dashed lines correspond to the sum of two components: the benchmark projection and the stochastic accumulation accruing from each respective structural shock.

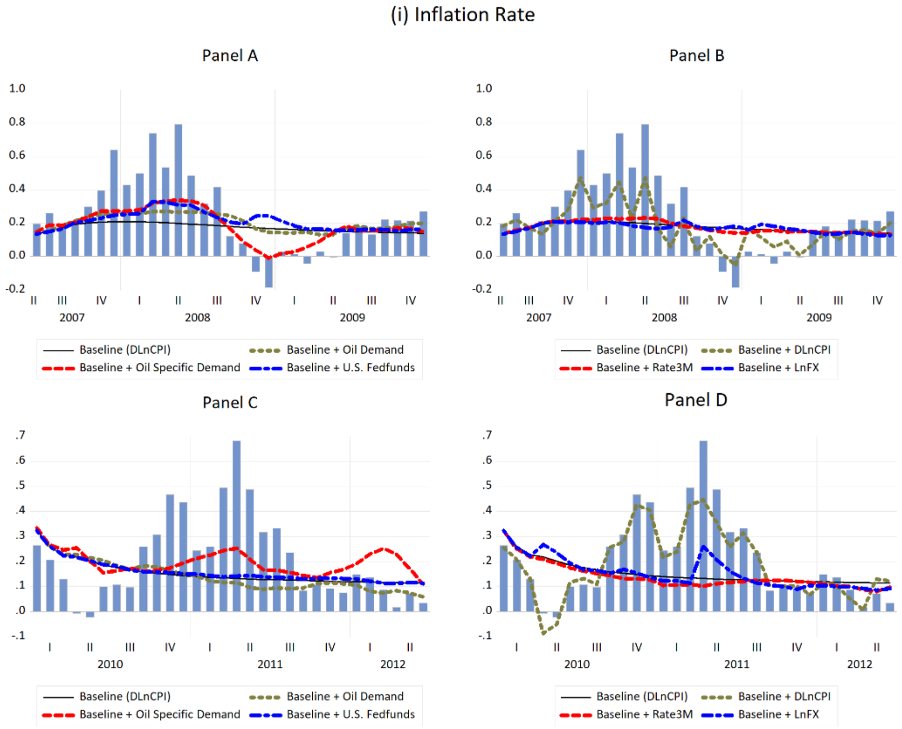


Figure 5: Historical variance decomposition of VN inflation rate.

Figure 5 shows that all three types of foreign shocks produced strong impacts on the VN inflation rate during the first inflationary episode, 2007:06 – 2008:06, but only the global demand oil and oil-specific demand shocks kept their strong influence in 2010:06 – 2011:09. The fall in inflation between 2008:09 and 2009:05 is mostly explained by oil-specific demand and autonomous inflation shocks. Given, in addition, the low impact of nominal interest rates (Panels B and D), one is bound to conclude that Vietnam's monetary policy was largely inefficient in the first inflationary episode, and it was thanks to domestic aggregate demand and oil price declines due to non-fundamental innovations that inflation pulled down subsequently. In sharp contrast, shocks to oil-specific demand were the only foreign factor that accelerated inflation in the second inflationary period. This may be explained by the socio-economic and political tensions around the world that caused the rise in precautionary oil demand between 2011 and 2014 (Lorusso and Pieroni, 2018).

Panel D of Figure 5 also shows that VN aggregate demand was a major force driving inflation in 2010 – 2011 for the reasons mentioned in Bhattacharya (2014), namely, movements in nominal effective exchange rates, real output growth, and credit expansion. Actually, in mid-2011 the nominal VND/USD exchange rate had a considerable impact on VN consumer prices, but it is clear that the impact was short-lived.

Turning to interest rate dynamics, Figure 6 shows that the benchmark projections (black lines) exhibit an opposite influence on the movements of interest rates in the past two periods. A slight upward trend between 2007 and 2009, while a downward line is observed in the period 2010 – 2012. Similar to Figure 5, I find that oil-specific demand and inflation shocks strongly affected interest rates, even though the latter were capped at 13.5% for sixteen months between 2010:12 and 2012:03. Panel D suggests that autonomous interest rate shocks considerably reduced the combined effects of oil-specific, inflation, and nominal exchange shocks that would have raised the interest rate to above 14% had it not been capped.

Figure 7 shows the base projection from 2007:01 to 2011:12, which is remarkably close to the long-run trade-balance deficit of -2.5% of GDP, as highlighted by Pham *et al.* (2019). Panel A shows that the U.S. monetary policy implemented in the aftermath of the GFC helped improve the VN trade balance, but oil price shock due to global demand worsened it notably in most of the assessed months. Finally, I observe that the trade balance was driven essentially by shocks *per se,* while the rest of domestic shocks had limited impacts on it.

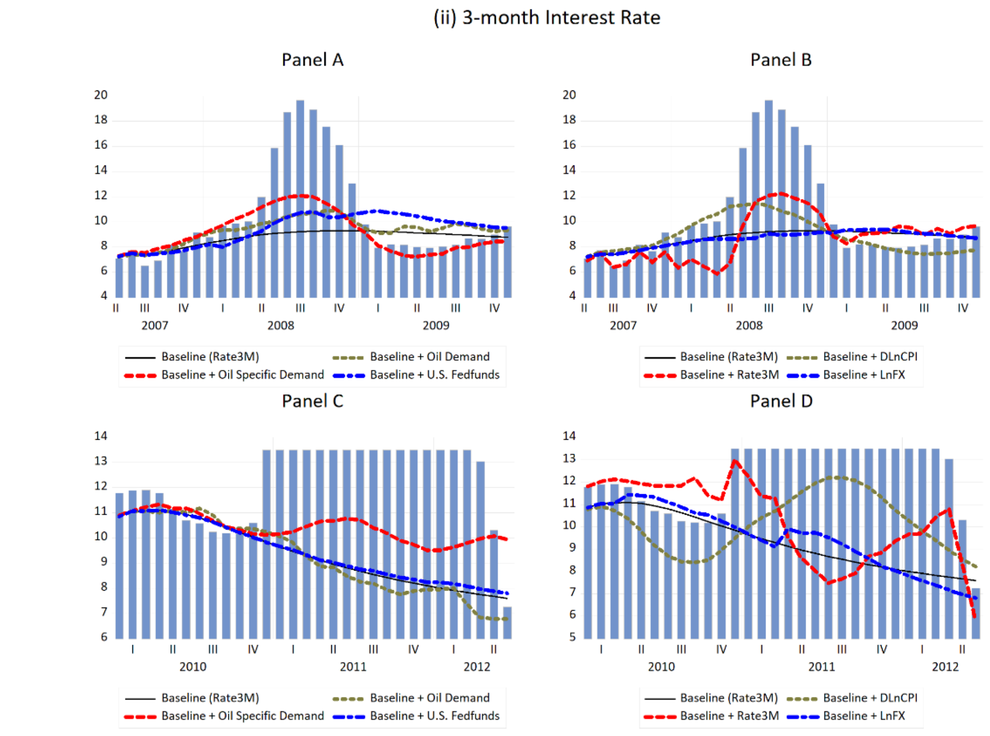


Figure 6: Historical variance decomposition of the VN 3-month interbank interest rate.

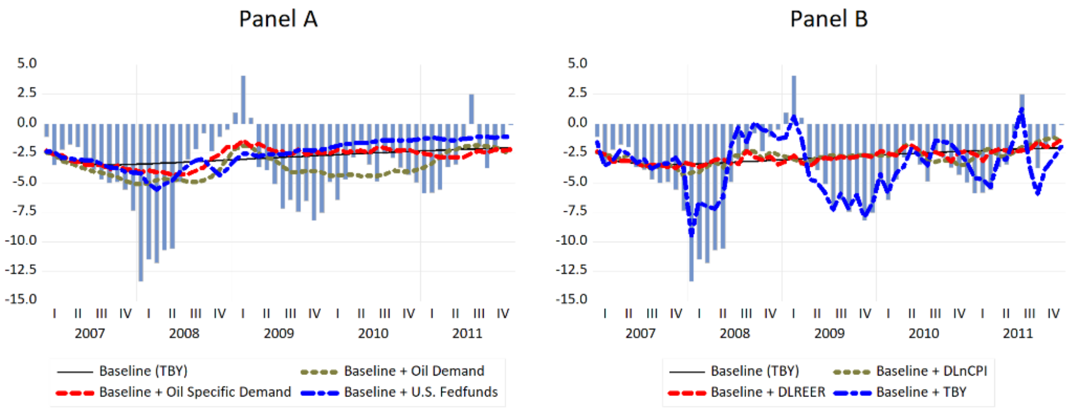


Figure 7: Historical variance decomposition of trade-balance-over-output.

3.6 Conclusions

Vietnam is known to be a small open country in the process of completing its transition to a free-market economy. It is less known, however, that it was a net exporter of crude oil until 2009, at the same time that it had to import up to 70% of the gasoline consumed domestically and most cracked petroleum products. In spite of the long-time planned and expected oil refineries, only Dung Quat started to be in operation in 2009 and could not counterbalance the growing need of oil imports. Given this twofold exposition (to trade in general, and to oil prices in particular), Vietnam's economic prospects crucially depend on the ability to manage the consequences of potential global shocks. This paper provided a step forward in the understanding of the mechanisms through which such shocks may condition Vietnam's economic performance.

To conduct the analysis, I followed Kilian's (2009) framework and examined the macroeconomic consequences of different oil price shocks: (1) oil supply shocks; (ii) oil demand shocks reflecting changes in the level of global economic activity (also called global demand oil shocks); and (3) oil-specific demand shocks, which are also referred to as precautionary, speculative or non-fundamental demand shocks.

Under Kilian's (2009) baseline model, my analysis yielded a first important insight for Vietnam's economy, namely that its CPI and RER are fairly responsive to both types of oil demand shocks (and not to the supply-side shock). In particular, the recovery of inflation to oil-specific demand shocks, whose impact clearly dyes down within one-year, is much faster than to global demand oil shocks. The persistence of the latter implies that one standard deviation of such shock raises VN CPI by 3.6% per annum and reduces the RER by 13.1% per annum in the same 24-month horizon. This leads to a strong appreciation of the VND against the U.S. Dollar. A first important result is, therefore, the harm of global demand oil price shocks in terms of inflation and international price-competitiveness.

Models I and II allow a deeper evaluation on the way oil price and risk-premium shocks affect competitiveness in Vietnam. First, the little influence from oil supply disruptions is confirmed both from the IRFs of the RER and the REER, which display similar responses. Second, the harm in terms of international price-competitiveness is also confirmed when oil price shocks arise either from global demand or speculative activities. Impulses in both cases strengthen the VND during at least a year, thereby leading to cheaper foreign goods for VN households.

Amid this general appraisal, Kilian's (2009) expanded setting allows the identification of idiosyncratic responses of RER and REER when the economy is hit by global demand oil price shocks. On the one hand, the RER and the REER react in opposite ways to such noises. On the other hand, the impact on the REER varies depending on the nature of the demand-side perturbation. A global demand oil price shock significantly raises VN's REER after eleven months, causing a loss in trade competitiveness (exports become more expensive, imports become cheaper, or both since Vietnam is at the same time an oil exporter and an oil importing economy). In contrast, an oil-specific demand shock decreases REER growth rate by about 0.3 percent per month initially and quickly overshoots to reach a peak of 0.15% in the fourth month. This indicates that in case of an oil-specific shock the VN economy may start losing its relative competitive advantage in one quarter. Of course, REER responses were examined together with TBY's ones and I saw that oil price increases actually impair Vietnam's trade balance, and of course, the current account.

Turning to risk-premium shocks, I uncovered remarkable effects of the U.S. Federal Fund rate because of the strong tie between the two respective currencies. In model I with the RER, I disclosed the strong impact of risk-premium shocks on Vietnam's inflation rate due to VND pegging to the US dollar. In contrast, in the multilateral setting brought by the REER, I found short-lived price responses. This is the outcome of international competition beyond the pegging of VND and the US dollar. Inflation responds more quickly to risk-premium shocks because in the multilateral trade context there is more penetration of imported goods and services in the worse currency scenario brought by such shocks. For the same reason, Vietnam's trade balance reacts positively to an increase in the exogenous risk premium, since that type of shock depreciates the VND in the following ten months so that exports are enhanced, and imports restrained.

Model 3 allowed us to assess how reactive the monetary policy is to oil price and risk-premium shocks. It is in this case that I found significant impacts of oil supply shocks on the 3-month interest rate and the nominal VND/USD exchange rates during the first two months after the shock. The possibility of quickly increasing oil imports may be the reason why this response, although significant, tends to be short-lived.

I also found that the 3-month interest rate (and/or short-term interest rates) has been greatly sensitive to both types of oil demand shocks and to changes in international financial risk.

Given the lack of inflation sensitivity to supply-side oil price shocks delivered by Kilian's (2009) baseline model, the significant monetary policy response to these shocks in the first two months after the shock was a new result requiring some interpretation. Especially in a context in which inflation on one side, and the monetary policy response on the other, showed a detachment (moving, respectively, upwards and downwards) consistent with the so-called "price puzzle" (Castelnuovo and Surico, 2006). This may be revealing of a conservative implementation of the monetary policy in Vietnam in recent decades, which would be consistent with Bhattacharya's (2014) assessment of Vietnam's persistently higher inflation vis-à-vis other Asian emerging economies. I therefore conclude that Vietnam's authorities were quite conservative in their reaction to international demand shocks and failed to counteract the inflationary pressures brought from that shocks. In this context, Bhattacharya's (2014) call for a forward-looking monetary policy in Vietnam is also endorsed by the present analysis.

As the variance decomposition analysis showed, this is likely to have had adverse effects on competitiveness on account of the essential role played by both types of demand-side oil price shocks in the long-run variations of several VN macroeconomic indicators, most particularly on the trade balance.

In a context in which inflationary periods are mainly driven by the effects of oil price shocks, I have confirmed that it is likely that Vietnam's monetary policy was to some extent inefficient in the first inflationary period (2007-2009). However, in the second period (2010-2012) domestic aggregate demand and oil price declines due to non-fundamental innovations pulled down Vietnam's inflation rate and counterbalanced the impact of global demand oil price shocks for precautionary reasons. As shown by Lorusso and Pieroni (2018), a set of socio-economic and political tensions around the world from 2011 to 2014 resulted in a rising precautionary demand for oil. This was the only foreign driver of inflation in Vietnam in those years, but it is a key example of how global shocks may affect domestic macroeconomic performance thereby asking for an appropriate policy response.

References

Ahmed, H.J.A., Wadud, I.K.M.M., 2011. Role of oil price shocks on macroeconomic activities: An SVAR approach to the Malaysian economy and monetary responses. Energy Policy 39, 8062–8069.

Amisano, G. and Giannini, C., 1997. Topics in structural VAR econometrics. Springer.

An, L., Jin, X., Ren, X., 2014. Are the macroeconomic effects of oil price shock symmetric?: A Factor-Augmented Vector Autoregressive approach. Energy Economics 45, 217–228.

Anwar, S., Nguyen, L.P., 2018. Channels of monetary policy transmission in Vietnam. Journal of Policy Modeling 40, 709–729.

Basnet, H.C., Upadhyaya, K.P., 2015. Impact of oil price shocks on output, inflation and the real exchange rate: evidence from selected ASEAN countries. Applied Economics 47, 3078–3091.

Baumeister, C., Kilian, L., 2014. Do oil price increases cause higher food prices? Economic Policy 29, 691–747.

Bhattacharya, R., 2014. Inflation dynamics and monetary policy transmission in Vietnam and emerging Asia', Journal of Asian Economics, 34, pp. 16–26.

Bohi, D.R., 1991. On the macroeconomic effects of energy price shocks. Resources and Energy 13, 145–162.

Castelnuovo, E. and Surico, P., 2006. The Price Puzzle: Fact or Artefact? Bank of England Working Paper. No. 288.

Cavalcanti, T., Jalles, J.T., 2013. Macroeconomic effects of oil price shocks in Brazil and in the United States. Applied Energy 104, 475–486.

Choi, S., Furceri, D., Loungani, D., Mishra, S., and Poplawski-Ribeiro, M. 2018. "Oil Prices and Inflation Dynamics: Evidence from Advanced and Developing Economies." Journal of International Money and Finance 82: 71–96.

Considine, T.J., 1988. Oil price volatility and U.S. macroeconomic performance. Contemporary Economic Policy 6, 83–96.

Cunado, J., Jo, S., de Gracia, F.P., 2015. Macroeconomic impacts of oil price shocks in Asian economies. Energy Policy 86, 867–879.

Elgammal, M.M., Eissa, M.A., 2016. Key determinants of inflation and monetary policy in the emerging markets: Evidence from Vietnam. Afro-Asian Journal of Finance and Accounting 6, 210–223.

Faria, J.R., Mollick, A.V., Albuquerque, P.H., León-Ledesma, M.A., 2009. The effect of oil price on China's exports. China Economic Review 20, 793–805.

Fukunaga, I., Hirakata, N., Sudo, N., 2011. The Effects of Oil Price Changes on the Industry-Level Production and Prices in the United States and Japan, in: Commodity Prices and Markets, East Asia Seminar on Economics, Volume 20, University of Chicago Press. pp. 195–231.

Hamilton, J. D., 1983. Oil and the macroeconomy since World War II, Journal of Political Economy, 91(2), pp. 228–248.

Hamilton, J. D., 2003. What is an oil shock?, Journal of Econometrics, 113(2), pp. 363–398.

Holtemöller, O., Mallick, S., 2016. Global food prices and monetary policy in an emerging market economy: The case of India. Journal of Asian Economics 46, 56–70.

Inumula, K.M., Solanki, S., 2017. Exploring causal nexus between crude oil price and exchange rate for India. International Journal of Economic Research 14, 319–331.

Iwaisako, T., Nakata, H., 2017. Impact of exchange rate shocks on Japanese exports: Quantitative assessment using a structural VAR model. Journal of the Japanese and International Economies 46, 1–16.

Jiménez-Rodríguez, R., Sánchez, M., 2012. Oil price shocks and Japanese macroeconomic developments. Asian-Pacific Economic Literature 26, 69–83.

Katircioglu, S.T., Sertoglu, K., Candemir, M., Mercan, M., 2015. Oil price movements and macroeconomic performance: Evidence from twenty-six OECD countries. Renewable and Sustainable Energy Reviews 44, 257–270.

Kilian, L., 2009. Not all oil price shocks are alike: Disentangling demand and supply shocks in the crude oil market. American Economic Review 99, 1053–1069.

Kilian, L., 2019. Measuring global real economic activity: Do recent critiques hold up to scrutiny?. Economics Letters, 178, pp. 106–110.

Kilian, L., Rebucci, A., Spatafora, N., 2009. Oil shocks and external balances. Journal of International Economics 77, 181–194.

Kim, W.J., Hammoudeh, S., Hyun, J.S., Gupta, R., 2017. Oil price shocks and China's economy: Reactions of the monetary policy to oil price shocks. Energy Economics 62, 61–69.

Kisswani, K.M., 2015. Does oil price variability affect ASEAN exchange rates? Evidence from panel cointegration test. Applied Economics 48, 1–9.

Korhonen, I., Ledyaeva, S., 2010. Trade linkages and macroeconomic effects of the price of oil. Energy Economics 32, 848–856. Le, T.H., Chang, Y., 2013. Oil price shocks and trade imbalances. Energy Economics 36, 78–96.

Lee, B.R., Lee, K., Ratti, R.A., 2001. Monetary policy, oil price shocks, and the Japanese economy. Japan and the World Economy 13, 321–349.

Lee, C.C., Chiu, Y.B., 2011. Nuclear energy consumption, oil prices, and economic growth: Evidence from highly industrialized countries. Energy Economics 33, 236–248.

Lizardo, R.A., Mollick, A.V., 2010. Oil price fluctuations and U.S. dollar exchange rates. Energy Economics 32, 399–408.

Lorusso, M., Pieroni, L., 2018. Causes and consequences of oil price shocks on the UK economy. Economic Modelling 72, 223–236.

Narayan, S., 2013. Foreign exchange markets and oil prices in Asia. Journal of Asian Economics 28, 41–50.

Osorio, C., Unsal, D.F., 2013. Inflation dynamics in Asia: Causes, changes, and spillovers from China. Journal of Asian Economics 24, 26–40.

Pham, B. T., Sala, H. and Silva, J. I., 2019. Growth and real business cycles in Vietnam and the ASEAN-5. Does the trend shock matter?. Economic Systems (manuscript accepted).

Rafiq, S., Sgro, P., Apergis, N., 2016. Asymmetric oil shocks and external balances of major oil exporting and importing countries. Energy Economics 56, 42–50.

Ratti, R.A., Vespignani, J.L., 2013. Why are crude oil prices high when global activity is weak?. Economics Letters 121, 133–136.

Salvatore, D., Winczewski, G., 1990. World oil prices and O.E.C.D. trade balance. Open Economies Review 1, 89–106.

Sultonov, M., 2017. The impacts of the oil price fall on the exchange rates of ASEAN-5: Evidence from the 2014 oil price shock. Economics Bulletin 37, 468–479.

Tatom, J.A., 1988. Macroeconomics effects of the 1986 oil price decline. Contemporary Economic Policy 6, 69–82.

Trang, N.T.N., Tho, T.N., Hong, D.T.T., 2017. The impact of oil price on the growth, inflation, unemployment and budget deficit of Vietnam. International Journal of Energy Economics and Policy 7, 42–49.

Ur-Rahman, M., Zoundi, Z., 2018. Macroeconomic response of disentangled oil price shocks: Empirical evidence from Japan. Economics Bulletin 38, 2240–2253.

Zhang, D., 2008. Oil shock and economic growth in Japan: A nonlinear approach. Energy Economics 30, 2374–2390.

Zhao, L., Zhang, X., Wang, S., Xu, S., 2016. The effects of oil price shocks on output and inflation in China. Energy Economics 53, 101–110.

Appendix

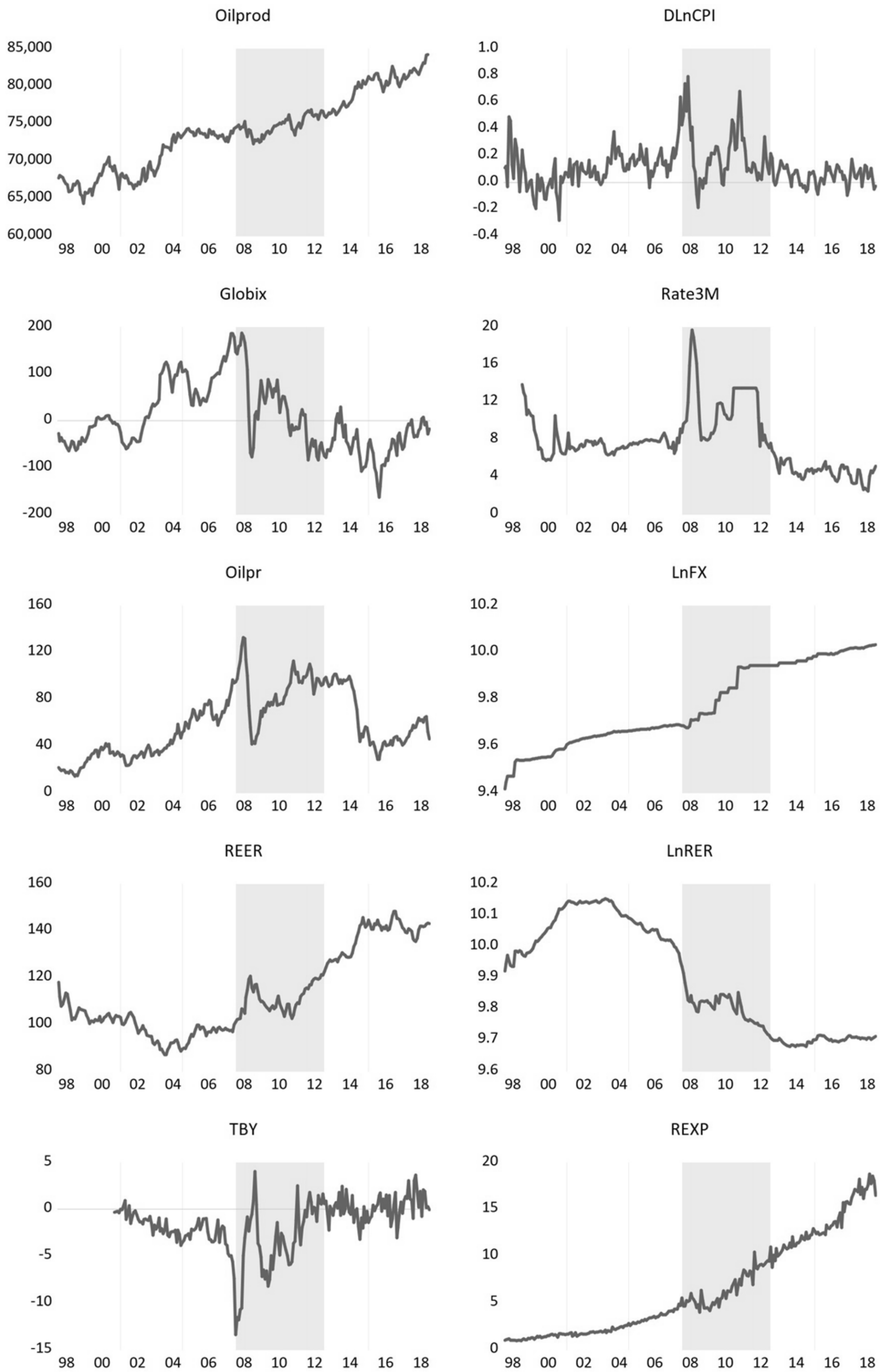


Figure A.1: Global prices and Vietnam’s macroeconomic indicators, 1998 – 2018.

	Stability		Information criteria			
Model	Largest root	*Lag*	FPE	AIC	SC	HQ
Baseline	0.97035	0	10.40853	16.53201	16.60204	16.56019
		1	0.012608	9.815951	**10.23612***	**9.985019***
		2	0.011602	9.732562	10.50287	10.04252
		3	**0.011231***	**9.699457***	10.81991	10.1503
Model I	0.99623	0	7.460511	21.87476	21.9728	21.91421
		1	2.37E-06	6.912553	7.69687	7.228146
		2	9.20E-07	5.96498	**7.435575***	**6.556717***
		3	**8.65e-07***	**5.901157***	8.05803	6.769038
Model II	0.95522	0	72.95096	24.15493	24.26801	24.20066
		1	0.000505	12.27318	**13.17785***	12.63906
		2	**0.000261***	**11.61174***	13.30799	**12.29776***
		3	0.000301	11.75056	14.23839	12.75672
Model III	0.99140	0	0.825113	19.6729	19.7769	19.71484
		1	8.91E-09	1.329123	**2.161094***	1.664648
		2	**4.65e-09***	**0.678190***	2.238136	1.307300*
		3	5.09E-09	0.765657	3.053577	1.688352

Figure A.1: VAR order selection.

	DLnOilpd	LnOilpr	Globix	DLnCPI	Rate3m	DLnRER	DLnREER	DLnREXP	TBY
DLnOilpd	1.000000								

LnOilpr	0.024619	1.000000							
	0.7203	-----							
Globix	0.055407	0.266420	1.000000						
	0.4200	0.0001	-----						
DLnCPI	0.004056	0.428292	0.521264	1.000000					
	0.9530	0.0000	0.0000	-----					
Rate3M	-0.073691	0.434138	0.340182	0.424267	1.000000				
	0.2832	0.0000	0.0000	0.0000	-----				
DLnRER	-0.063261	-0.210418	-0.235422	-0.438477	-0.165433	1.000000			
	0.3571	0.0020	0.0005	0.0000	0.0154	-----			
DLnREER	0.069859	0.210527	0.081512	0.242466	0.139802	-0.353575	1.000000		
	0.3091	0.0020	0.2351	0.0003	0.0410	0.0000	-----		
DLnREXP	0.055455	0.043316	0.064389	0.073961	-0.007027	0.085658	0.003776	1.000000	
	0.4196	0.5285	0.3486	0.2814	0.9186	0.2120	0.9562	-----	
TBY	-0.043473	-0.282740	-0.648303	-0.619260	-0.394005	0.258300	-0.002675	0.034144	1.000000
	0.5270	0.0000	0.0000	0.0000	0.0000	0.0001	0.9690	0.6194	-----

Note: Top-values are pairwise correlations; bottom-values are corresponding p-values.

Table A.2: Correlation matrix (p-values are below the correlation coefficients)

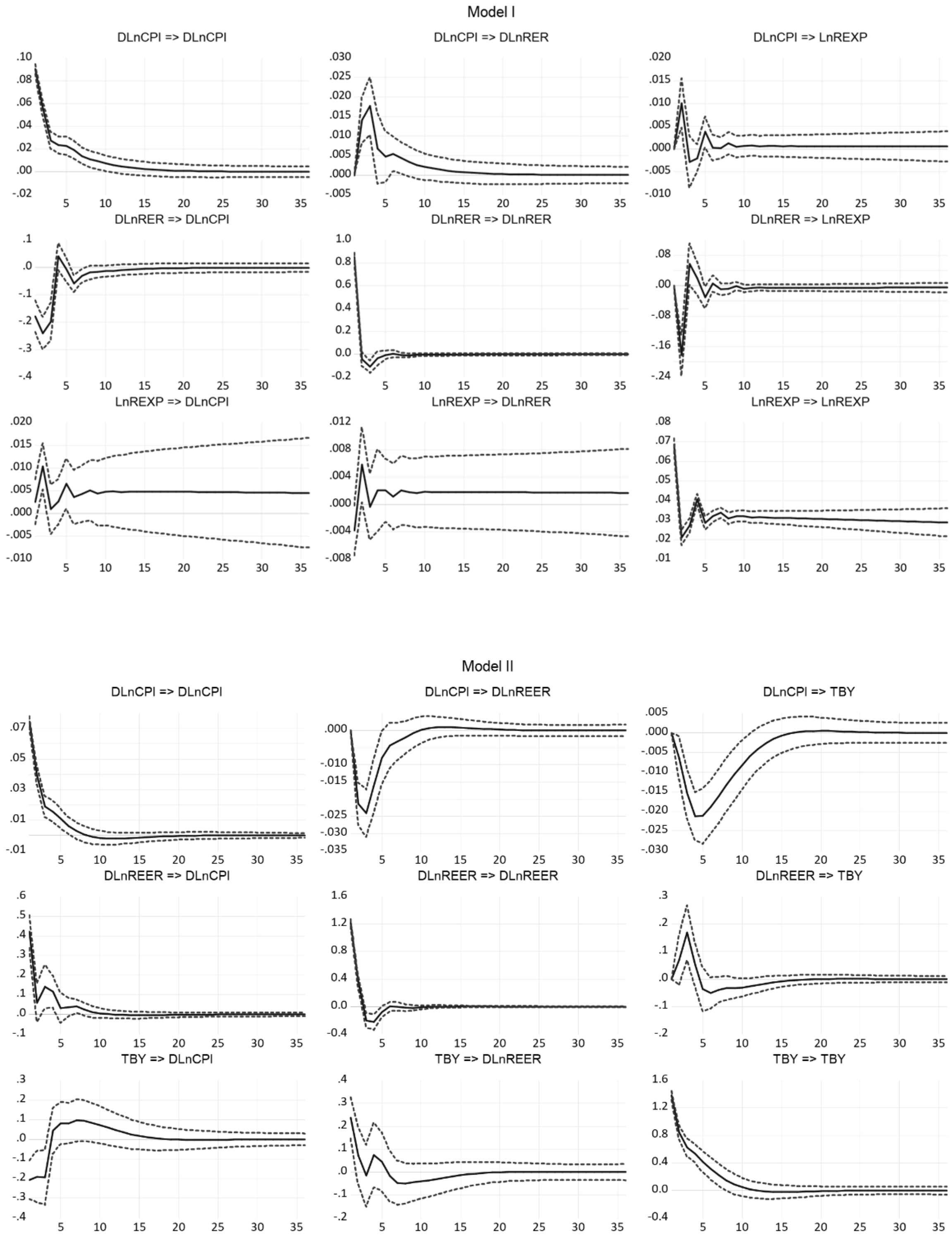


Figure A.3: Impulse - response functions w.r.t domestic structural shocks.

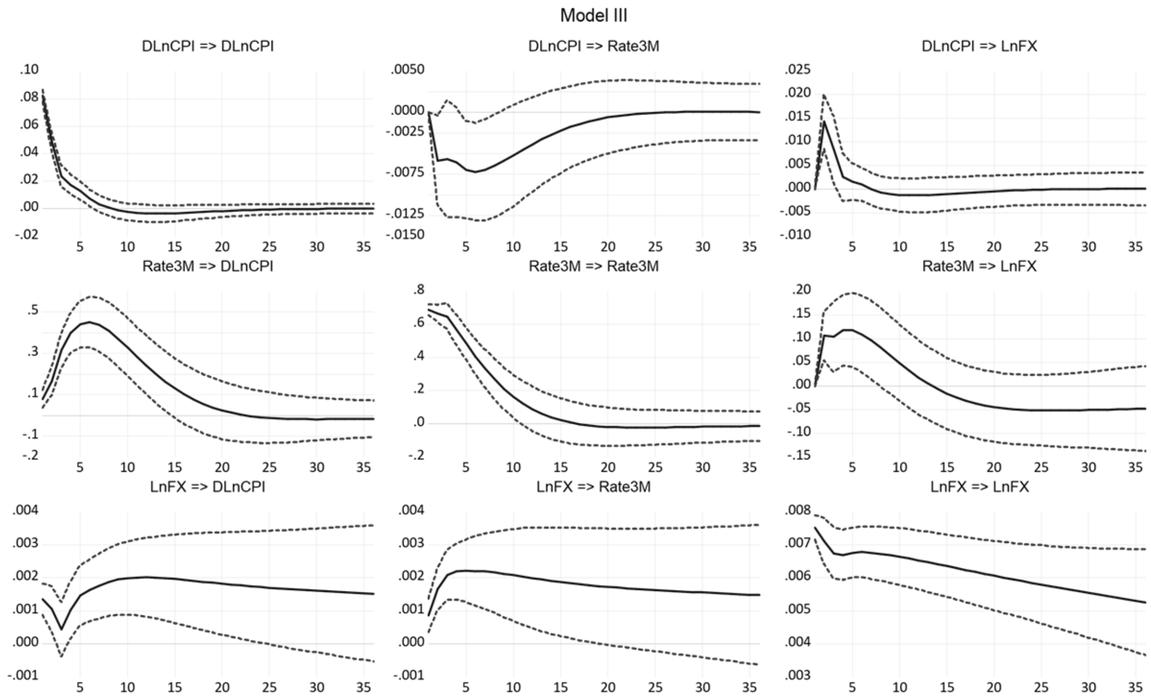


Figure A.3 (continue): Impulse - response functions w.r.t domestic structural shocks.

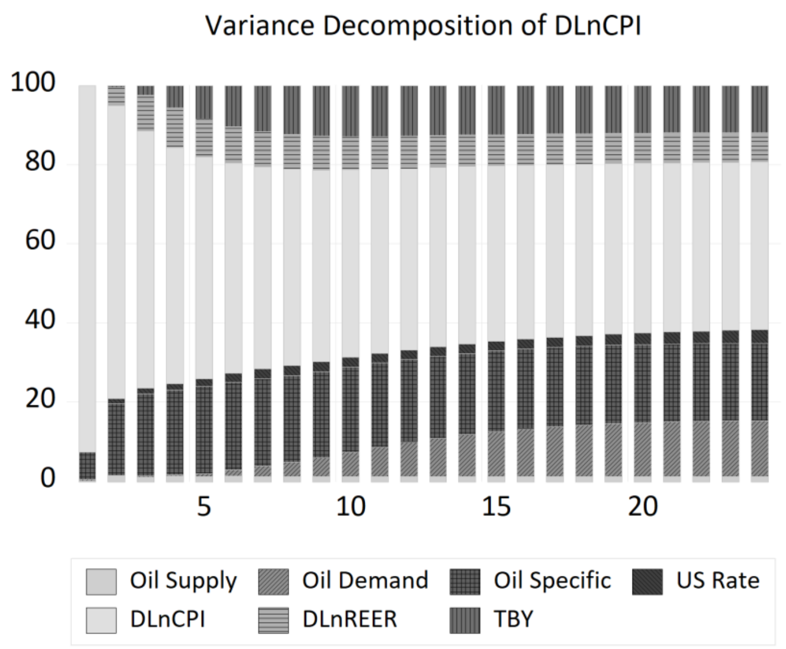
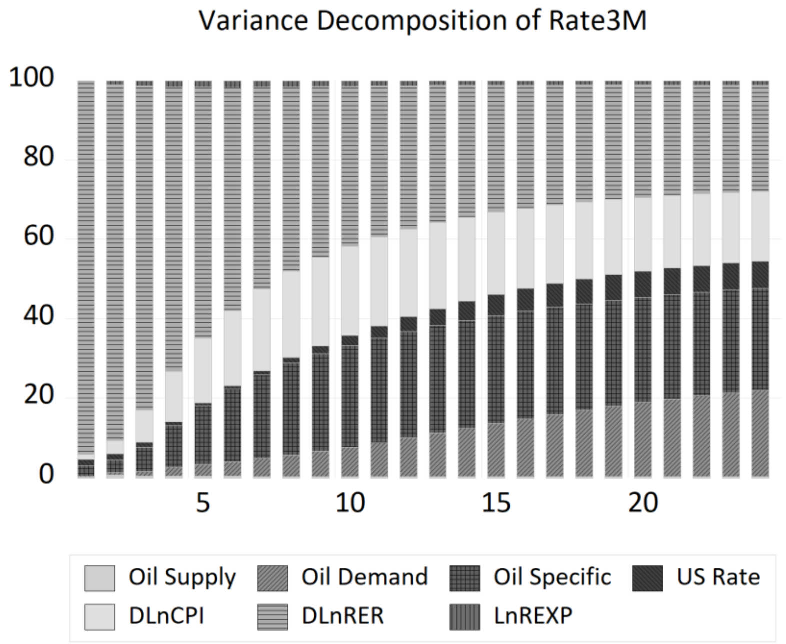


Figure A.4: variance decomposition of selected variables



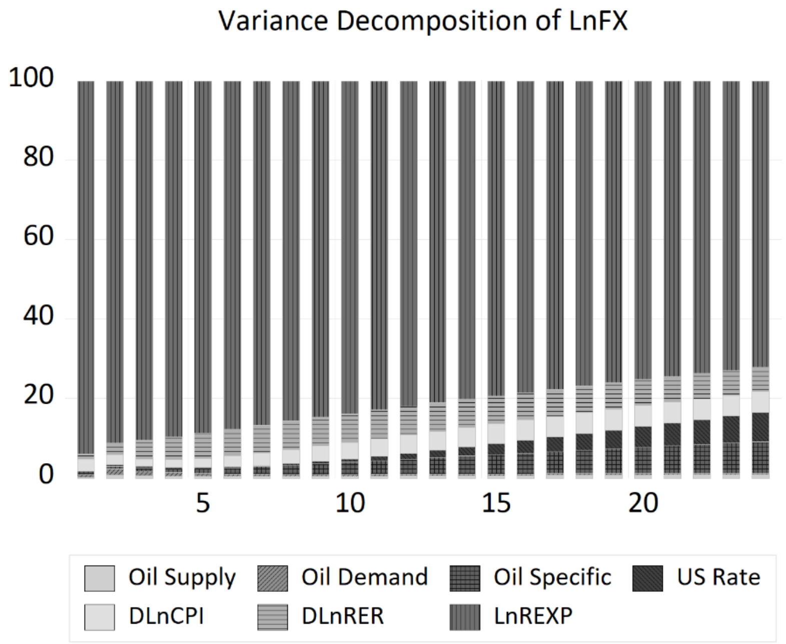
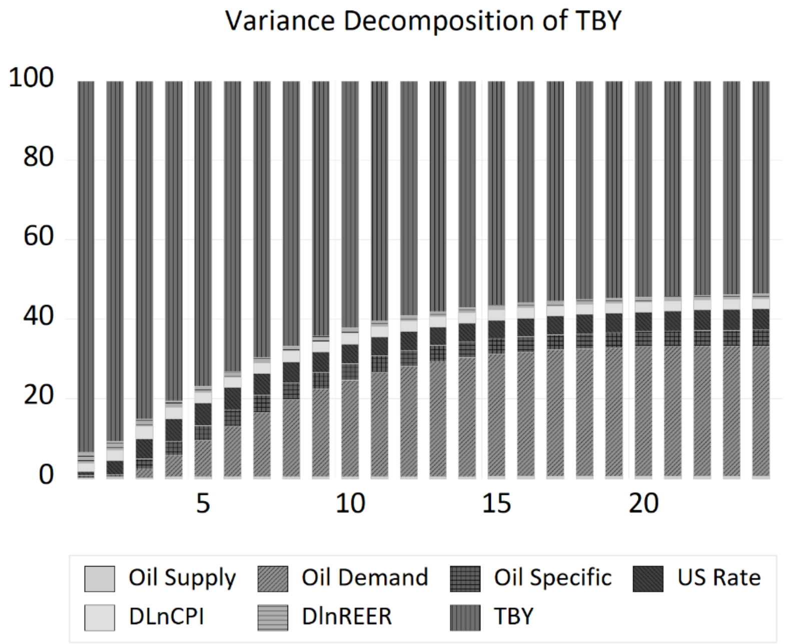


Figure A.4 (continue)



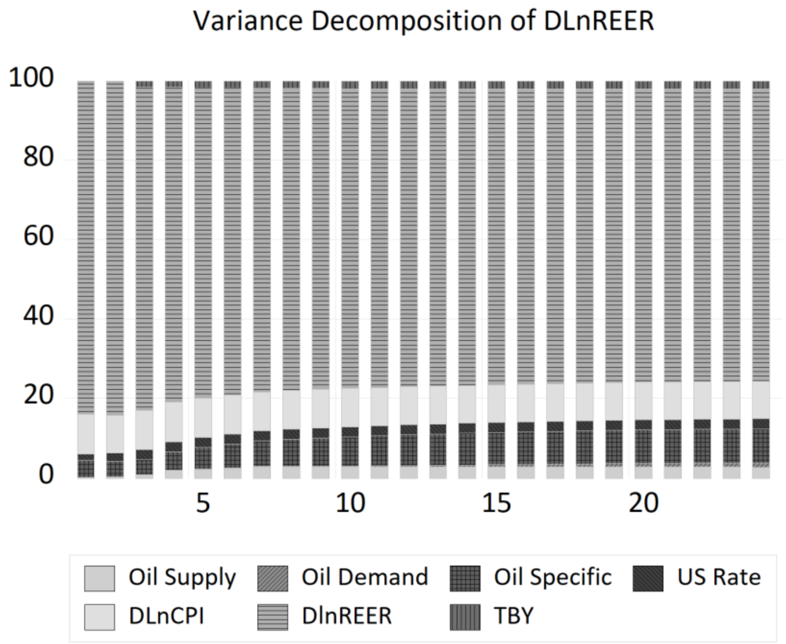


Figure A.4 (continue)

CHAPTER 4: Crowding-in or Crowding-out macroeconomic effect of public investment in Vietnam?♣

Abstract

By considering a regime switch in two structural vector error correction models (SVECMs) over the period 1976 – 2015, I plot the influence of Vietnam's public capital spending and public capital stock on private capital accumulation and output growth. I show that there has been a Granger-causal chain running from public capital to output and private capital, respectively. Furthermore, private capital and output do Granger-cause each other and one-way causal directions are found between them and employment. I also show that public investment (capital) positively impacts on private investment (capital) and output in both short- and long-run, firmly establishing the existence of crowding-in effects in Vietnam. Meanwhile, private investment plays a significant role in explaining variations in the employment. These findings highlight the growth effects of public capital in Vietnam during its transitional phase.

Keywords: public investment, public capital, crowding in, SVECM, Vietnam.

JEL codes: E62, H11, H54, O11.

♣ This is a joint work with Dr. Canh Phuc Nguyen at the University of Economics – Ho Chi Minh City (UEH) during my research visit from 15/07/2019 to 15/10/2019. The manuscript has been submitted to Singapore Economic Review.

4.1 Introduction

Raising expenditure by government, including current and capital spending, is asserted to be in line with economic development under Wagner's Law (Peacock & Scott, 2000). Endogenous growth theories (Ortigueira and Santos, 1997; Romer, 1994) assume that government spending is an endogenous factor in the growth function. As a matter of fact, the nexus between public investment, private investment, and economic growth has become one of the most important topics on the macroeconomic research agenda. Government spending, on the one side, might have positive impacts on growth in the long run through improvements in labour productivity, export capacity, and technological progress, which consequently 'crowds-in' private investment (Aschauer, 1989a; Barro and Sala-i-Martin, 1992; Lucas, 1988). On the other side, it is evident that inefficient public spending could 'crowd-out' private investment (Cavallo and Daude, 2011; Ogibayashi and Takashima, 2017), owing to decreasing returns to private capital (Abel, 2017). As a result, excessive increases in public investment could crowd-out private activities (Abel, 2017; Kandil, 2017) and thus retard economic growth (Buiter, 1977; Ganelli, 2003). Interestingly though, the empirical literature has seen a different mix of conclusions from developed and emerging economies (Cavallo and Daude, 2011; Erden and Holcombe, 2005; Kandil, 2017; Pereira, 2001; among many others).

Another strand in the literature extensively investigates the productive impacts of public capital investment such as 'core infrastructure' on the whole economy. In highly influential works, Aschauer (1989a, b) show that a rise in public capital likely raises the returns to private capital, but that not all types of public capital produce the same crowding-in effect on the accumulation of private capital. Regarding this matter, 'core infrastructure' investment – i.e. streets, highways, public transport systems, power grids, telecommunication networks etc. – appear to have stronger predictive power in explaining productivity progress (Aschauer 1989b; Fernald, 1999). Accordingly, public capital affects output growth directly through the 'total factor productivity' channel in the neoclassical economic framework.

Note, however, that public capital also induces indirect growth effects by “promoting human capital accumulation and innovation capacity”, as Agrénor and Neanidis (2015) assert.

This research empirically fills a void in the current literature by examining the responses of output and private investment (capital) to an injection of public investment from the Vietnamese state. Vietnam is, in fact, a typical example of transition economy from socialist to market economy over past decades. In this process, the public capital is likely to play a major role in economic growth (Su and Bui, 2017), but it important to consider potentially issues related to inefficient investment plans and uncompleted reforms (Anh, 2016; Pincus, Anh, and Le Thuy, 2008; Thanh and Duong, 2009). In this context, the purpose of my endeavour is twofold. First, I explore the long-run relationships and ‘Granger causalities’ between public (private) capital accumulation and output. Second, I investigate the macroeconomic transmission of public investment in Vietnam. To this end, I recruit two common cointegrated vector autoregression (VAR) models found in Kamp (2005) and Bahal et al. (2018). Notably, time breaks are tested and then incorporated in the analysis to deal with structural changes in the Vietnam economy (see e.g., Mallon and Irvin, 2001; Thanh and Duong, 2009) over the forty-year sample period, 1976 – 2015.

My research findings robustly document the positive impacts of public capital spending on private capital and output growth, highlighting the crowding-in effects of public capital investment in the transitional economy of Vietnam. Interestingly, the analysis shows that variations in employment are mostly explained by private capital accumulation. As a result, an appropriate public investment strategy is recommended for Vietnam economy, with the focus on ‘core’ or ‘productive’ infrastructure to improve economic development.

Section 2 and 3 discuss some relevant literature and the estimation strategy. Sections 4 and 5 present the empirical models and the univariate analysis of individual time-series, respectively. Model estimations and their corresponding impulse – response and variance investigations are in turn presented in Sections 6 and 7, respectively. The final section contains some concluding remarks and a policy discussion.

4.2 Background and related literature

4.2.1 Background on the Vietnamese economy

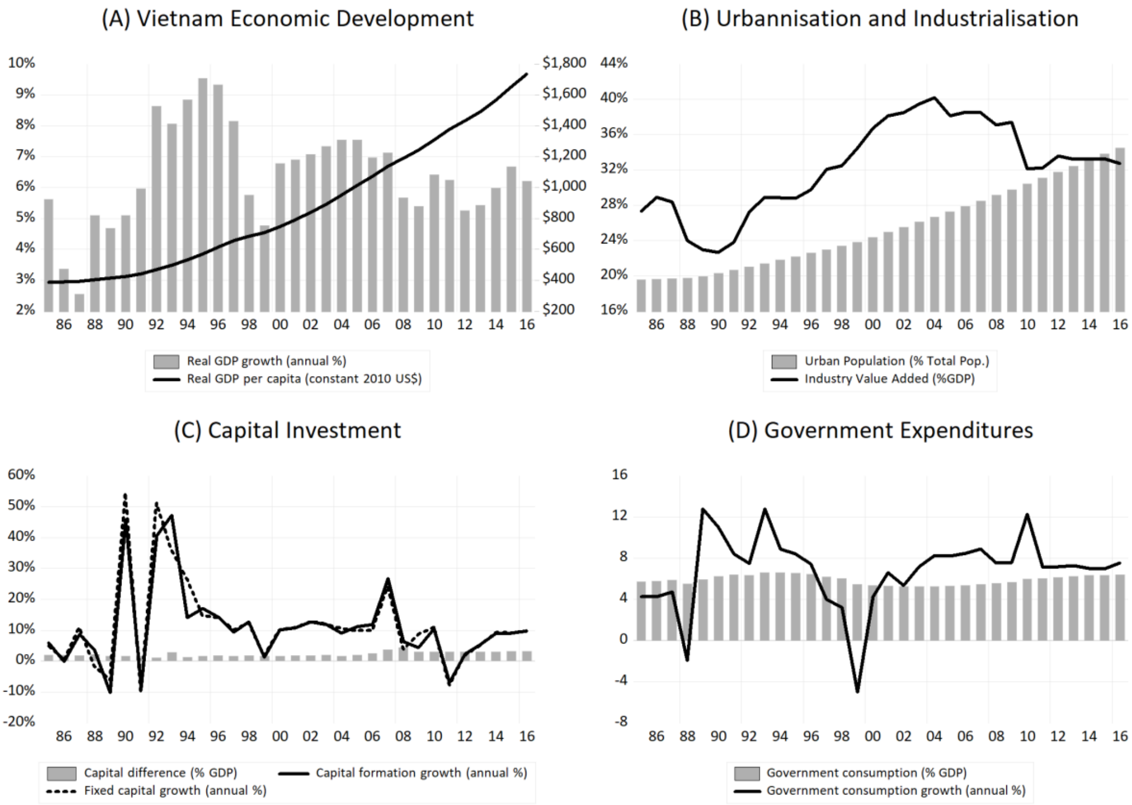


Figure 1: Vietnam's economic dynamics, 1985 – 2016.

Panel A of Figure 1 shows that the rates of real VN output growth were high in the period 1990 – 1996, i.e. five years after the 'Doi Moi', before falling in the three consecutive years 1997 – 1999 due to the Asian economic crisis. The growth rate in the period 2000 – 2007 continued to be stable, but at a slightly slower pace than in the previous decade. Suffering declines during the 2008 economic turmoil, Vietnam GDP growth fell to below 6% in 2008 and 2009, then rebounded thereafter, but was not as strong or stable as in the previous year. It should, however, be noted that the absolute income level of

Vietnam households has still been in line with the diminishing income growth rates in the 2010s of the group of lower-middle-income economies, implying the existence of a middle-income trap (Aiyar et al., 2013). Barker and Üngör (2019) claim that Vietnam economy should thus be less dependent on factor accumulation as its source of growth and that Vietnam should raise its technological capabilities and introduce appropriate economic policies to boost productivity growth rates in both agricultural and non-agricultural sectors. Furthermore, the productivity of the whole economy needs to be improved to avoid the middle-income trap (Ohno, 2009).

The urbanisation rate – the grey bars of Panel B in Figure 1 – has apparently increased over this period, reaching 35% in 2017 from a low of 20% in 1985, whereas the industrialisation rate (industrial value added in terms of percentage of GDP) has fallen since the 2008 global financial crisis after fifteen years of moderate growth, 1992 – 2007. Prior to the global financial crisis, it rose from about 25% in 1992 to its peak of 40% in 2004. The urbanisation and industrialisation rates, however, are far below the averages for high-income economies, indicating the need for infrastructure and public facilities to foster economic development and raise living standards (Arouri et al., 2017; Kang and Imai, 2012; Lanjouw and Marra, 2018).

Panel C presents a panoramic view of capital investment in Vietnam. It is interesting to note that the gross capital formation and gross fixed capital formation lines are not significantly different, suggesting that investment predominantly goes into building up fixed assets. The graph also shows that there was a sharp rise in investment in the short period 1991 – 1993, before a stable period of capital investment from 1995 to 2005. Joining the World Trade Organisation in 2007 sparked new capital investment that year, but the subsequent crises – the 2008 global financial crisis and the Vietnam banking system crisis – created a rapidly downward investment trend until 2013. Panel D shows that Vietnam (general) government spending decreased sharply in 1999 due to the 1997 Asian economic crisis, since Vietnam government had limited policy instruments and capability to expand its fiscal budget at that time. Government spending was then gradually increased to nearly 9% of GDP in 2007 from its low of 5% in 2000. It peaked at 12.3% in 2010 as a consequence of fighting the 2008 global recession. In the period 2012 – 2017, government spending fluctuated around a fairly high rate of 7.1%.

Since the 'Doi Moi', there have been several reforms in fiscal policy – enacted by the 1996 and 2002 State Budget Law, resulting in several markedly positive effects on economic development (Leung, 2010; Nguyen and Anwar, 2011). However, it has also witnessed problematic fiscal issues that are impairing the effectiveness of public spending in promoting and sustaining economic growth (Anh, 2016; Fritzen, 2006; Pincus, 2009; Su and Bui, 2017). As pointed out by Rao (2000), the budget determination process in Vietnam was predominantly 'top-down', and local governments had no power to raise revenue. Anh (2016) recently argues that fiscal decentralisation in Vietnam has not always been accompanied by institutional autonomy and sufficient financial resources for local governments. Nguyen and Anwar (2011) show that even though revenue decentralisation is positively associated with economic growth, expenditure decentralisation has caused reverse impacts on 61 provinces' economies over the period 2002 – 2007. Su and Bui (2017) study 63 VN provinces over the period 2005–2013 and find a non-linear relationship (inverted-U shape) between provincial government size and private investment growth. Nguyen et al. (2017) analyse the interrelationships between output, foreign direct investment, international trade, the inflation rate, and state investment in Vietnam. They find the existence of pair-wise bidirectional and unidirectional causalities among output, inflation, and state investment, but their impulse-response analysis, surprisingly, shows that public investment negatively impacts output.

All in all, Vietnam has remained a lower-middle country and has experienced rapid changes in urbanisation, industrialisation, and globalisation in the past decades (Abbott and Tarp, 2012; Anwar and Nguyen, 2011a, 2011b; Le and Tran-Nam, 2018), which have subsequently induced many socio-economic problems such as urban poverty, inequality, and environmental degradation (Dollar, 2002; Kang & Imai, 2012; Magrini et al., 2018). A clear understanding of the short- and long-run impacts of public capital spending on the dynamics of Vietnam economy is to approach these problems effectively from a macroeconomic perspective.

4.2.2 A brief literature review

Several studies have documented the crowding-in effects of public spending. For instance, Erenburg (1993) and Pereira (2001) find evidence of a crowding-in effect of public investment on private investment in the United States (U.S.). Dreger and Reimers (2016) report crowding-in effects in twelve Eurozone members between 1991 and 2012, but Afonso and St. Aubyn (2009) only confirm the existence of such an expansionary effect in six out of the twelve countries over a much longer period between 1960 and 2005. Romero-Ávila and Strauch's (2008) findings robustly demonstrate that public investment had a positive impact on long-run growth and private capital accumulation in the EU-15 countries.

The finding that there are crowding-out effects or mixed effects arising from investment by the state is, however, not uncommon in empirical research. Voss (2002), for example, shows that public investment crowded-out private investment in the U.S. and Canada over the five-decade period between 1951 and 1997. Facchini and Melki (2013) review 84 empirical studies on the nexus between government spending and economic growth and documented that 66.6% of these studies showed negative growth effects, only 8.3% showed positive effects, and a surprising 25.1% were inconclusive.

Abiad et al. (2015) study 17 advanced countries and claim that "increased public investment raises output, both in the short term and in the long term, crowds-in private investment, and reduces unemployment". Kamps (2005) shows positive effects of public capital on output, but not on employment, in 22 OECD economies. De Jong et al. (2018) confirm most of the results in Kamp (2005). Likewise, Hunt (2012) also finds the crowding-in impacts of public capital in a study of 20 OECD members.

Atukeren (2005) documents both crowding-in and crowding-out effects in 25 developing countries, while Erden and Holcombe (2005) note crowding-in effects in 19 developing countries, but crowding-out effects in 12 developed countries, between 1980 and 1996. Furthermore, Cavallo and Daude (2011) emphasize a dominant crowding-out effect in a panel of 116 developing countries from annual data between 1986 and 2006. They argue that public investment, on one hand, raises the marginal productivity of private capital and leads to potential crowding-in of private investment, but weak institutions and restricted access to financing could, on the other hand, diminish the positive effects of public investment projects and then crowd-out private investment.

4.3 Econometric methodology

This section presents a multivariate econometric framework for non-stationary time-series analysis. First, Sims' (1980) theory of Vector Autoregression (VAR), where variables in the system are possibly integrated at order 1 but not necessarily cointegrated is summarised first. Next, I present a vector error-correction model (VEC) for dealing with cointegrated covariates, as they are of vital interest for modelling Vietnamese national account aggregates.

4.3.1 Vector Autoregression (VAR)

Let us consider the *K*-dimensional VAR(*p*) process having the form

$$y_t = \upsilon + A_1 y_{t-1} + \cdots + A_t y_{t-p} + u_t \tag{1}$$

where $y_t = (y_{1t}, \dots, y_{Kt})'$ is the set of *K*-variables of interest and u_t is a corresponding *K*-dimensional zero-mean white noise process, which has the variance-covariance matrix $E(u_t u_t') = \Sigma_u$, that is $u_t \sim N(0, \Sigma_u)$. By construction, A_i is a *K-by-K* matrix of i^{th}-lag of y_t with $i = 1...p$. The deterministic term υ may contain deterministic variables such as a constant, time and or seasonal dummies, so it should be a $(K \times m)$ matrix. If the VAR only has a constant, then $m = 1$. The above VAR(*p*) system can be rewritten in the companion form VAR(1) as

$$Y_t = \Upsilon + AY_{t-1} + U_t \tag{2}$$

where $Y_t \equiv (y_t, y_{t-1}, \dots, y_{t-p+1})'$ and $\Upsilon \equiv (\upsilon, 0, \dots, 0)'$ are $Kp \times 1$ vectors. For any stable VAR(1) process, there always exists a Moving Average Representation (MAR) defined as the sum of past and current innovations, so that the system (2) is also read

$$Y_t = \sum_{i=0}^{\infty} A^i \Upsilon + \sum_{i=0}^{\infty} A^i U_{t-i} \qquad (3)$$

4.3.2 Vector Error Correction Model (VECM)

The concept of cointegration was pioneered in the seminal works of Granger (1981) and Engle and Granger (1987). Technically speaking, a set of I(1) time-series variables is said to be cointegrated if there exists a linear combination of these variables that is I(0).

The VECM($q = p - 1$) corresponding to the reduced VAR(p) in (1) can be written as

$$\Delta y_t = \mathbf{\Pi} y_{t-1} + \Gamma_1 \Delta y_{t-1} + \cdots + \Gamma_{p-1} \Delta y_{t-p+1} + \Phi D_t + u_t \qquad (4)$$

where $\mathbf{\Pi} = -(I_K - A_1 - \cdots - A_p)$; $\Gamma_i = -(A_{i+1} + \cdots + A_p)$ with $i = 1...q$. D_t denotes a vector of deterministic terms such as constant, trend, or intervention dummies *etc*. The vector of error terms, u_t, has the usual meaning.

It is worth noting that Δy_{t-k} for $k = 0...q$ is stationary by construction, hence, $\mathbf{\Pi} y_{t-1}$ must be stationary. Matrix $\mathbf{\Pi}$ can be decomposed as $\mathbf{\Pi} = \alpha\beta'$, so that $\beta' y_{t-1}$ describes cointegrating relationships, whereas matrix α presents how these vectors of linear combinations are loaded into the system. Thus, estimating a VEC model requires knowing the cointegrating vectors β and the loading matrix α, implying that the rank of $\mathbf{\Pi}$ must be known. Johansen (1995) provides the suitable analysis to determine this rank. The likelihood ratio rank tests of $\mathbf{\Pi}$ suggested by Johansen (1995), however, depend upon the involvement of constant and linear trends in both the VECM(q) and implied VAR(p) specifications. Accordingly, let us rewrite (4) as

$$\Delta y_t = \alpha\beta' y_{t-1} + \sum_{i=1}^{q} \Gamma_i \Delta y_{t-i} + \mu_0 + \mu_{1,t} + u_t \qquad (5)$$

Johansen (1995) consider five cases depending upon how the constant $\mu_0 = \alpha\beta_0 + \gamma_0$ and the linear trend $\mu_1 = \alpha\beta_1 + \gamma_1$ are restricted into the cointegrating space. Of these, we opt for the situation μ_0 unrestricted but $\beta_1 \neq 0$ and $\gamma_1 = 0$ are restricted into $\mathbf{\Pi}$ such that $\mathbf{\Pi} y_{t-1} = \alpha[\beta' : \beta_1][y_{t-1} : 1]'$. Given the chosen setting, the testing hypotheses are thus written

H(K): rank of $\mathbf{\Pi} = K$, *i.e.* y_t has no unit-roots;

H(r): rank of $\mathbf{\Pi} = r$, *i.e.* y_t has r cointegration relationships or $(K - r)$ unit-roots.

The test should run from $r = 0$ to $r = K$ (top-to-bottom approach), and if the null hypothesis H($r = r^*$) has been correctly rejected, then the conclusion is that the endogenous vector y_t has at most $(r^* + 1)$ cointegrating vectors.

Similarly to the MAR representation (3) of system (1), Johansen (1995) showed that any VECM(q) also has a so-called Granger's representation as

$$y_t \;=\; \boldsymbol{C}\sum\nolimits_{i=1}^{t} u_i \;+\; \boldsymbol{C}^*(L)u_t \;+\; y_0^* \tag{6}$$

where

$$\boldsymbol{C} = \beta_\perp\left[\alpha'_\perp\left(I_K - \sum\nolimits_{i=1}^{q=p-1}\Gamma_i\right)\beta_\perp\right]^{-1}\alpha'_\perp$$

where $\alpha_\perp$ and $\beta_\perp$ are orthogonal complement matrices so that $\beta'\beta_\perp = 0$ and $\alpha'\alpha_\perp = 0$. $\mathbf{C}$ is the long-run impact matrix because $\sum u_i$ is a K-dimensional unit-root process. $\mathbf{C}^*(L)u_t$ and y_0^* are K-vector of I(0) and initial value, respectively. Moreover, the rank of matrix $\mathbf{C}$ is of $(K - r)$, since $\mathbf{\Pi}$ has a rank of r.

4.3.3 Long- and Short-run identifications

If u_t in (6) is replaced by $u_t = \mathrm{B}\omega_t$ such that $\omega_t \sim N(0, I_K)$, $\Sigma_u = \mathrm{BB}'$, with B being a $(K \times K)$ matrix accounting for contemporaneous relationships among variables in u_t (hence, y_t), then VEC model (6) has a structural form such that matrix $\mathbf{F} = \mathbf{C}\mathrm{B}$ represents the long-run effects of unit structural shocks ω_t on y_t in level. Since $\mathbf{C}$ has rank $(K - r)$, then $\mathbf{F}$ contains at most r columns of zero elements, i.e. there are r transitory shocks in the system (6).

Responses of y_t to unit shocks ω_t can be estimated with imposed constraints on $\mathbf{F}$ or B or both matrices, which correspond to long-, short-, and short-and-long-run identification schemes, respectively. Specifically, vector ω_t can be decomposed as $\omega_t = (\omega'_{P,t} : \omega'_{T,t})'$, where $\omega_{P,t}$ contains $(K -$

r) × 1 innovations having permanent effects on y_t as $\lim_{h\to\infty} \partial E_t(y_{t+h})/\partial\omega_{P,t} \neq 0$; the ($r$ × 1) matrix $\omega_{T,t}$ produces only transitory effects as $\lim_{h\to\infty} \partial E_t(y_{t+h})/\partial\omega_{T,t} = 0$. As stated by Gonzalo and Ng (2001) and Breitung *et al.* (2004), if one places restrictions on both F and B matrices, then a total of ½ $K(K-1)$ conditions are required, of which each zero column in F is equivalent to $(K - r)$ restrictions.

4.3.4 Unit-root tests

From the above discussions, it can be seen that it is necessary to test whether or not a time series is difference-stationary, i.e. following a unit-root process, since cointegration itself only appears to be valid among I(1) variables. Dicker and Fuller's (1979) unit-root test has the form

$$\Delta z_t = (1-\theta)z_{t-1} + \sigma\epsilon_t \tag{8}$$

where z_t is a univariate process, c and δ are constant and linear trends, respectively, and ε_t is normally distributed with constant variance of σ^2. If the equality of $(1-\theta) = 0$ cannot be significantly statistically rejected, then z_t has a unit-root. The alternative is $(1-\theta) < 0$, implying that z_t is stable. Note, however, that equation (9) does not account for either the deterministic trend of variables in level or the serial correlation in the innovations. Therefore, the test in the augmented form (ADF test) below is frequently used.

$$\Delta z_t = \underbrace{\gamma_0 + \gamma_1 t + \gamma_2 t^2}_{\boldsymbol{D_t\gamma}} + (1-\theta)z_{t-1} + \sum_{j=1}^{p} \delta_j \, \Delta z_{t-1} + \epsilon_t \tag{9}$$

Besides this, there are some popular alternatives, such as the Phillips and Perron (1988) (or PP) and Kwiatkowski, Phillips, Schmidt, and Shin (1992) (or KPSS) tests. While the former validates the null in (10) by controlling serial correlation via non-parametric method, the latter adopts a different approach as its null assumes the process z_t being (trend-) stationary. Consequently, one might expect the alternative rather than the null in the KPSS procedure if $z_t \sim I(1)$.

4.3.5 Structural break tests

In all foregoing unit-root tests, the model does not allow for change in the level of series or change in growth rate or both. Perron (1989) showed that the probability of accepting the false null increases when one ignores an existing breakpoint. Presuming a single known break date, T_b, Perron (1989) considers three distinguishing situations: (i) case A validates the unit-root null of one-time 'crash' in the intercept μ for date $t > T_b$; (ii) meanwhile, case B allows for a drift term shifting from μ_1 to μ_2 at time T_b; (iii) consequently, the two effects are simultaneously modelled in case C. Favouring the alternative hypothesis in Perron (1989) implies the process z_t is trend-stationary with: (A) a one-time shift in the intercept; (B) a '*changing trend slope*' without any '*crash*' in the level; and (C) a jump in the level followed by a different growth trend.

There are two major weaknesses in the Perron (1989) test, which I will discuss in turn. The first is quite obvious, i.e. that the break date T_b is often unknown, rather than an endogenously testable parameter. Zivot and Andrews (1992), Banerjee *et al*. (1992), and Volgelsang and Perron (1998), among others, extended the ADF test by allowing one data-dependent breakpoint, meaning that one break-date is endogenously identified. For a long macroeconomic series, there are possibly three or more regimes in the unit-root process, but it seems practical to not consider more than three regime switches with respect to low-frequency data, so I will not pursue further two-break unit-root tests,[1] given the commonly short annual capital stock data.

4.4 Data and model settings

To quantify the macroeconomic impacts of public investment and public capital accumulation on the Vietnamese economy over the post-war decades 1976 – 2015, I adopt two common VEC systems from the literature (see e.g., Pereira and Roca-Sagales, 2001; Everaert, 2003; Kamps, 2005; Bahal *et al.,* 2018; and Afonso and Aubyn, 2018; among others), as follows:

- Model I: (*Output*, *Public Investment*, *Private Investment*) = (y, i_g, i_p).
- Model II: (*Public Capital*, *Private Capital*, *Output*, *Employment*) = (k_g, k_p, y, h).

[1] I refer interested readers to Lumsdaine and Papell (1997) and Lee and Strazicich (2003) for multiple break tests.

With Model I, I study two aspects: (i) whether investment by the state crowds-in its private counterpart; and (ii) to what extent output responds to an impulse of public investment. In the meantime, Model II aims to (iii) test the crowding-in hypothesis on public capital, i.e., whether public capital positively Granger-causes economic growth and private capital accumulation; and (iv) to assess the impacts of public capital shocks quantitatively.

I make use of national account data downloaded from the United Nations Statistical Division (UNStad)[2], and the IMF investment and capital stock dataset.[3] All variables are expressed in logarithms of levels (represented by the small letters of their corresponding capital letters) then taken first-difference to obtain their respective rates of growth. It is also worth noting that UNStad provides aggregate data in terms of constant 2010 USD, whereas IMF public capital and investment data are adjusted to constant 2011 international dollars. Since currency adjustment factors are multiplicative, all log-level time series have essentially preserved their own characteristics, *i.e.* their growth rates and trends.

4.5 Univariate analysis

Figure 2 presents the time-paths of the growth rate (prefixed by Δ) and log-level variables of output (y), public investment (i_g), private investment (i_p), public capital (k_g), private capital (k_p), and employment (h). From the subplots of investment and capital components, one can visually observe a regime switch in the early 1990s, *i.e.* within five years after the 'Doi Moi'. Specifically, both public and private investment strongly fluctuated and shifted to higher levels in 1991. As a consequence, the slopes of capital stock trend-lines became much steeper after the new constitutional law of 1992. Within this context private capital grew at a diminishing rate after the Asian economic crisis, so that Δk_p is deemed to be a potential I(1) process, implying that K_p is potentially an I(2) variable. This is consistent with Everaert (2003), who documented that the capital growth rate could be non-stationary in a finitely small

[2] http://unstats.un.org
[3] http://www.imf.org/external/np/fad/publicinvestment/

capital stock sample. However, Musolesi (2011) showed that if a structural break (regime switch) is evident in the time series, one should take this break into account for the unit-root testing process.

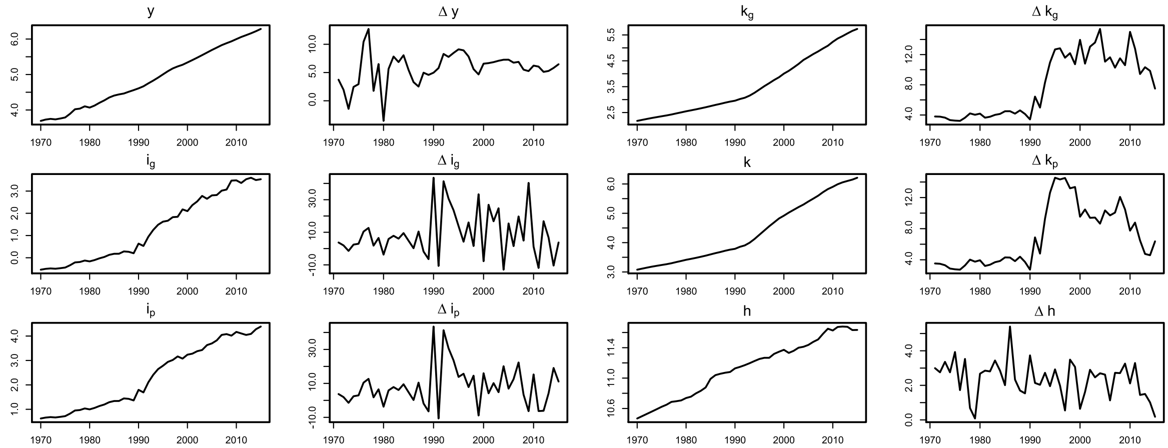


Figure 2: Vietnam aggregate time series. Data sample: 1970 – 2015.

	ADF test (H_0: unit root)					KPSS test (H_0: stationary)				
Variables	Specs	Stats	5%	1%	p-value*	Specs	Stats	5%	1%	p-value**
y	(CT, 5)	-1.44	-4.21	-3.53	> 0.10	(CT, 4)	0.16	0.22	0.15	< 0.05
k_g	(CT, 2)	-2.47	-4.21	-3.53	> 0.10	(CT, 5)	0.20	0.22	0.15	< 0.05
k_p	(CT, 6)	-2.26	-4.21	-3.53	> 0.10	(CT, 5)	0.17	0.22	0.15	< 0.05
i_g	(CT, 0)	-2.06	-4.21	-3.53	> 0.10	(C, 5)	0.76	0.74	0.46	< 0.01
i_p	(CT, 5)	-1.98	-4.21	-3.53	> 0.10	(C, 5)	0.76	0.74	0.46	< 0.01
h	(CT, 0)	-0.81	-4.21	-3.53	> 0.10	(CT, 4)	0.17	0.22	0.15	< 0.05
Δy	(C, 4)	-3.45	-3.61	-2.94	< 0.02	(C, 2)	0.08	0.74	0.46	> 0.10
Δk_g	(C, 1)	-1.41	-3.61	-2.94	> 0.10	(C, 5)	0.58	0.74	0.46	< 0.05
$\Delta \mathrm{k}_p$	(C, 5)	-1.48	-3.61	-2.94	> 0.10	(C, 5)	0.36	0.74	0.46	< 0.10
Δi_g	(C, 0)	-8.09	-3.61	-2.94	< 0.01	(C, 1)	0.16	0.74	0.46	> 0.10
$\Delta \mathrm{i}_p$	(C, 4)	-3.34	-3.61	-2.94	< 0.02	(C, 4)	0.10	0.74	0.46	> 0.10
Δh	(C, 0)	-5.87	-3.61	-2.94	< 0.01	(C, 2)	0.24	0.74	0.46	> 0.10

Note: C = constant; T = trend; AIC lag-length for ADF, Newey bandwidth for KPSS.
Small letters express variables in logarithms.
* MacKinnon (1996) one-sided p-values; ** Probability based on KPSS (1992, Table 1).

Table 1: Unit root tests (1976 – 2015)

Unsurprisingly, *Table 1* shows that all series in logs are non-stationary variables as we strongly statistically (i) fail to reject the null of unit root in all augmented Dickey and Fuller (1979), or ADF tests; and (ii) reject the null of stationary in Kwiatkowski, Phillips, Schmidt, and Shin (1992), or KPSS, tests. Nevertheless, the null of unit-root cannot be rejected at the significant level of 10% for the two rates of capital growth–Δk_g and Δk_p–in the ADF test. KPSS tests provide a similar conclusion, which leads us to conclude that Δk_g and Δk_p are potentially I(1) variables.

Since Vietnam economy started to reform in the mid-1980s, unit-root tests with respect to growth variables should handle potential regime switches properly. This is the reason why I select single-break unit root tests as displayed in *Table 2*. Note that the reported break-year (T_b) has been endogenously determined from the data rather than a priori specified as in Perron (1989). It is shown that the null of single-break ADF unit root test (see *e.g.*, Zivot and Andrews, 1992; Volgelsang and Perron, 1998) has not been statistically accepted in all cases, suggesting that the Vietnam capital stock data could be appropriately treated as normal I(1) variables in a VEC model if the structural break is carefully handled.

ADF test with intercept (H_0: unit root)					
Variables	Break	T_b	Stats	1%	p-value*
Δy	(C, IO, 3)	1991	-7.93	-4.94	< 0.01
Δk_g	(C, IO, 0)	1992	-4.80	-4.94	< 0.02
Δk_p	(C, IO, 4)	1990	-4.99	-4.94	< 0.01
Δi_g	(C, IO, 0)	1989	-8.93	-4.94	< 0.01
Δi_p	(C, IO, 0)	1989	-7.66	-4.94	< 0.01

*Vogelsang (1993) asymptotic one-sided p-values.
(C, IO, *k*) = (crash model, innovational outlier, SIC lag-length).

Table 2: Single-break unit root tests for growth variables (1976 – 2015)

4.6 I(1) cointegration analysis

Given that all variables in model I and II are first order integrated, I now validate whether they are cointegrated at some orders before estimating their respective multivariate systems. In this context, Johansen's (1995) rank test has been employed to verify the long-run relationships among endogenous covariates in (y, i_g, i_p) and (k_g, k_p, y, h).

To estimate the rank of matrix $\mathbf{\Pi}$, two settings must be specified: firstly, the treatment of deterministic terms in the VECM; secondly, dummy variables accounting for structural breaks. Because all components of model I and II straightforwardly exhibit linear trends, the VECM should be modelled with unrestricted intercept and restricted trend terms. In addition, two unrestricted dummy variables, *D*79 and *S*91, are used to handle breaks and mean shifts in all the time series, respectively. D79 is a blip dummy accounting for transitory effects of the northern border war in 1979, taking zeroes everywhere except for 1 and –1 in the years 1979 and 1980, respectively. S91 is a shift variable taking value 1 for 1991 onward, so as to account for the endogenous breakpoints obtained from the reported single-break unit-root tests.[4] Note, however, that no dummies have entered the cointegrating space.

Model	$p-r$	r	Eig.Val	Trace	Trace*	Frac95	p-value	p-value*
I	3	0	0.538	55.10	49.08	46.68	0.007	0.029
	2	1	0.335	25.02	23.10	28.57	0.130	0.196
	1	2	0.208	9.12	7.21	14.90	0.272	0.431
II	4	0	0.719	88.34	66.66	63.66	0.000	0.027
	3	1	0.504	38.86	30.86	42.77	0.120	0.458
	2	2	0.236	11.48	10.19	25.73	0.843	0.910
	1	3	0.024	0.96	0.91	12.45	0.995	0.996

Star "*" indicates Barlett correction value for the small sample.
VECM specification: unrestricted constant, restricted trend.
Model I and II have two common dummies D79 and S91.

Table 3: Johansen (1995) cointegration tests. Data sample: 1976 – 2015.

Table 3 show that the null (zero rank of $\mathbf{\Pi}$) is strongly statistically rejected in both models, implying that at least one cointegrating vector exists in each case. The subsequent null, however, can not be rejected, implying that only one long-run equilibrium relationship among (y, i_g, i_p) and (k_g, k_p, y, h) can be drawn. To validate this conclusion, I compute and report roots of the companion matrices in *Table 4*. Models I and II have two- and three-unit-roots, respectively, if *rank*($\mathbf{\Pi}$) is set to one, and their second largest eigenvalues are markedly below unity. Setting *rank*($\mathbf{\Pi}$) to two implies the appearance of some close-to-unity roots that signal one more non-stationary process in the model. Besides this, I conduct

[4] In fact, our results are robust to the choice of break-year (1990, 1991 or 1992) but we pick up $T_b = 1991$ because of several important political events happened in Vietnam and the world.

several typical model mis-specification tests and report their statistics and cointegrating relations in the Appendix. It is shown that models with blip and shift dummies generally exhibit better statistical properties, and that their respective cointegrating relations have a lower variance.

Choosing the cointegrating rank of one produces empirical vector β_s, as displayed in *Table 5* below. It is shown that private output (y) only cointegrates with public investment (i_g) in model I since the coefficient of i_p in the cointegration space (β_{ip}) is statistically insignificant and fails to reject a zero-restriction test on β_{ip}. Likewise, Model II shows that the coefficient of employment variable (β_h) can be restricted to zero without affecting its **Π** matrix. Besides, the sign of all coefficients on i_g, k_g and k_p is opposite to that on y, showing that a rise of public (private) capital stock or state investment would have positive impacts on private output in the long run.

Model	I		II		
Modulus	$r = 1$	$r = 2$	$r = 1$	$r = 2$	$r = 3$
Root 1	1.000	1.000	1.000	1.000	1.000
Root 2	1.000	0.834	1.000	1.000	0.828
Root 3	0.608	0.587	1.000	0.812	0.828
Root 4	0.608	0.587	0.301	0.812	0.320
Root 5	0.374	0.425	0.301	0.283	0.320
Root 6	0.254	0.425	0.575	0.283	0.313
Root 7	-	-	0.575	0.020	0.331
Root 8	-	-	0.009	0.020	0.002

Table 4: Roots of companion matrices

Choosing the cointegrating rank of one produces empirical vector β_s, as displayed in *Table 5* below. It is shown that private output (y) only cointegrates with public investment (i_g) in model I since the coefficient of i_p in the cointegration space (β_{ip}) is statistically insignificant and fails to reject a zero-restriction test on β_{ip}. Likewise, Model II shows that the coefficient of employment variable (β_h) can be restricted to zero without affecting its **Π** matrix. Besides, the sign of all coefficients on i_g, k_g and k_p is opposite to that on y, showing that a rise of public (private) capital stock or state investment would have positive impacts on private output in the long run.

Model | Variable	k_g	k_p	h	y	i_g	i_p	*Trend*
I				1.000	-0.176	-0.011	-0.042
					(-5.878)	(-0.345)	(-22.832)
I*				1.000	-0.185	0.000	-0.042
[0.797]					(-12.101)		(-22.479)
II	-0.097	-0.140	0.163	1.000			-0.045
	(-2.307)	(-2.981)	(0.820)				(-8.008)
II*	-0.093	-0.155	0.000	1.000			-0.041
[0.508]	(-2.224)	(-3.385)					(-34.955)

t-statistics are in parentheses. "*" indicates zero-restriction on β's elements.
p-values of β-restriction tests are in squared-brackets.

Table 5: Cointegrating vectors and β-restrictions

The endogenous variables under assessment have had long-run equilibrium relationships but not all of them can have 'levels feedback' properties. As stated by Juselius and Hendry (2000), if a VECM system has r cointegrating vectors (or $rank(\mathbf{\Pi}) = r$), then it has at most $(p - r)$ common trends, indicating that matrix α of size $(p \times r)$ only has r non-zero rows. Therefore, Models I and II could have two and three weakly exogenous variables, respectively. The weakly exogenous hypothesis is

$$H_{0,\alpha}(r) : \alpha = A\tilde{\alpha}$$

where A is a $(p \times s)$ matrix with $s \geq r$ and $\tilde{\alpha}$ is a $(s \times r)$ matrix of α-coefficients differing from zero. The alternative can thus be written

$$H_{1,\alpha}(r) : B\alpha = 0$$

with $B = A_{\perp}$ such that $A_{\perp}A = 0$.

The test results are summarised in *Table 6*. It can be easily seen that there is only one variable in both systems, output, which adjusts to long-run relations so that the mixture of uni- and bidirectional Granger causality between output and other variables could be as reported in *Table 7*. It is shown that neither public capital nor private capital directly 'Granger-cause' employment. Interestingly, K_p appears to Granger-cause K_g, but not *vice versa*, although one might expect a bidirectional causality. Private capital and output do Granger-cause each other, while one-way causal directions are found between them and employment. Consequently, a causal chain exists among the three variables K_g, K_p and Y, which suggests that public capital would indirectly Granger-cause private capital and employment. Since I_g and I_p are components of the national account identity, they obviously Granger-cause output as shown in the last two columns of *Table 7*. But it does not exist a causality relationship between I_g and I_p. Nor do such causalities run from Y to I_g, and from Y to I_p.

In sum, I find that public investment and public capital unambiguously have positive effects on Vietnamese output growth in general and on private capital accumulation in particular over the sample period. This finding supports the Aschauer's (1989a, b) hypothesis, but only partially because the cointegration analysis refers to a long-run relationship. The question regarding the extent to which public capital (investment) crowds-in private capital (investment) and output has not yet fully been addressed.

Model	k_g	k_p	h	y	i_g	i_p
I				13.951	0.031	1.791
				[0.000	[0.860]	[0.181]
II	0.492	1.059	0.941	21.300		
	[0.483]	[0.303]	[0.332]	[0.000		

p-values of $\chi^2(1)$ statistics are in squared-brackets.

Table 6: Weak exogeneity tests

Model	$H_0, j \mid H_0, i$	k_g	k_p	h	y	i_g	i_p
I	y				-	0.000	0.000
	i_g				0.794	-	0.952
	i_p				0.128	0.180	-
II	k_g	-	0.001	0.704	0.576		
	k_p	0.295	-	0.420	0.032		
	h	0.584	0.436	-	0.582		
	y	0.000	0.000	0.000	-		

H_0: variable i is Granger non-causal for variable j. Numbers are p-values.

Table 7: Granger causality tests

4.7 Impulse–response and variance analysis

By appropriately establishing some identification schemes, I now assess the short- and long-run impacts of various structural shocks on private investment and capital in the two models considered. It is worth recalling that Cholesky decomposition (recursive identification), adopting multivariate time series techniques, has often been used in recent public capital (investment) literature. For instance, Kamps (2005) estimates a VECM, but implicitly sets up a recursive VAR model to explore the short-run effects. Kamp argues that a permanent change in government expenditure would permanently affect private output, so he intentionally focuses on a short-run analysis. In contrast, in order to undertake a deep study of the effects of public investment and public capital, I impose zero restrictions on elements of both matrices **F** (long-run impact matrix) and *B* (short-run impact matrix).

Let $u^I{}_t = (u^y{}_t, u^{ig}{}_t, u^{ip}{}_t)$ and $u^{II}{}_t = (u^{kg}{}_t, u_t^{kp}, u^h{}_t, u^y{}_t)$ be vectors of the reduced VEC models I and II, respectively. Then, $\omega_t^I = (\omega_t^y, \omega_t^{ig}, \omega_t^{ip})$, and $\omega_t^{II} = (\omega_t^{kg}, \omega_t^{kp}, \omega_t^h, \omega_t^y)$ are the corresponding vectors of structural shocks, such that $u_t^I = B\omega_t^I$ and $u_t^{II} = B\omega_t^{II}$. In these structural shocks, ω_t^y represents innovation in productivity, which, along with ω_t^i, possibly generates long-term effects, as suggested by Bahal *et al*. (2018). Conversely, private investment and labour supply shocks – ω_t^{ip} and ω_t^h, respectively – are assumed to be transitory, as it is well established after Blanchard and Quah (1989) and King *et al.* (1991). Following Aschauer (1989b), 'core' public capital is supposed to be a productive asset, and therefore, ω_t^{kg} and ω_t^{kp} could have both long- and short-run effects on the variables in the system.

As discussed in the methodology section, the number of cointegrating vectors in the VEC system is the number of transitory shocks defined in **F**, so that either Model I or II has one variable that only produces a short-run impact. I thus assume that both structural innovations, ω_t^{ip} and ω_t^h, have no long-run impacts on themselves and other variables and that, as a consequence, their corresponding columns in F are full of zeros, of which one only counts two and three zero restrictions, respectively. To fulfil the restriction requirements of Model I, I impose the third zero on B to enforce that a public investment shock has no contemporaneous effects on its private counterpart since it has generally been implemented with lags.

For Model II, three additional zero restrictions are needed. First, it is not unreasonable to assume that a one-time shock to technological progress would not produce long-lasting impacts on public capital and employment. Second, I deliberately limit the impact of private capital on public capital in a short period. Accordingly, *Table 8* summarises zero constraints placed on **F** and B.

Model	Specification	**F** = **C**B	B
I	(y, i_g, i_p)	$\begin{bmatrix} * & * & 0 \\ * & * & 0 \\ * & * & 0 \end{bmatrix}$	$\begin{bmatrix} * & * & * \\ * & * & * \\ * & 0 & * \end{bmatrix}$
II	(k_g, k_g , h, y)	$\begin{bmatrix} * & 0 & 0 & 0 \\ * & * & 0 & * \\ * & * & 0 & 0 \\ * & * & 0 & * \end{bmatrix}$	$\begin{bmatrix} * & * & * & * \\ * & * & * & * \\ * & * & * & * \\ * & * & * & * \end{bmatrix}$

"*" denotes parameters to be estimated.

Table 8: Long- and short-run identification schemes

Panels A and B of *Figure 3* show the responses of output and private investment to one standard deviation impulse of public (private) investment and capital. The shaded bands – from lightest to darkest – represent 90%, 70%, and 50% bootstrapped error bands, respectively. The time axis represents forecasting horizons, while the vertical axis expresses response values in percentages.

Because of the constraints on the long-run impacts in **F**, output and public investment statistically significantly revert to their equilibrium states within five years after their initially positive responses, by around 0.85% and 2.1%, respectively, with respect to one standard deviation, or 3%, private investment shock (see the plots (c) and (d) of Panel A). In sharp contrast, a rise of private capital does not promote output, but instead considerably crowds out government investment, as shown in the plot (i) in the second row of Panel B. This suggests that private capital in Vietnam appears to be a substitute for, rather than a complement to, public capital.

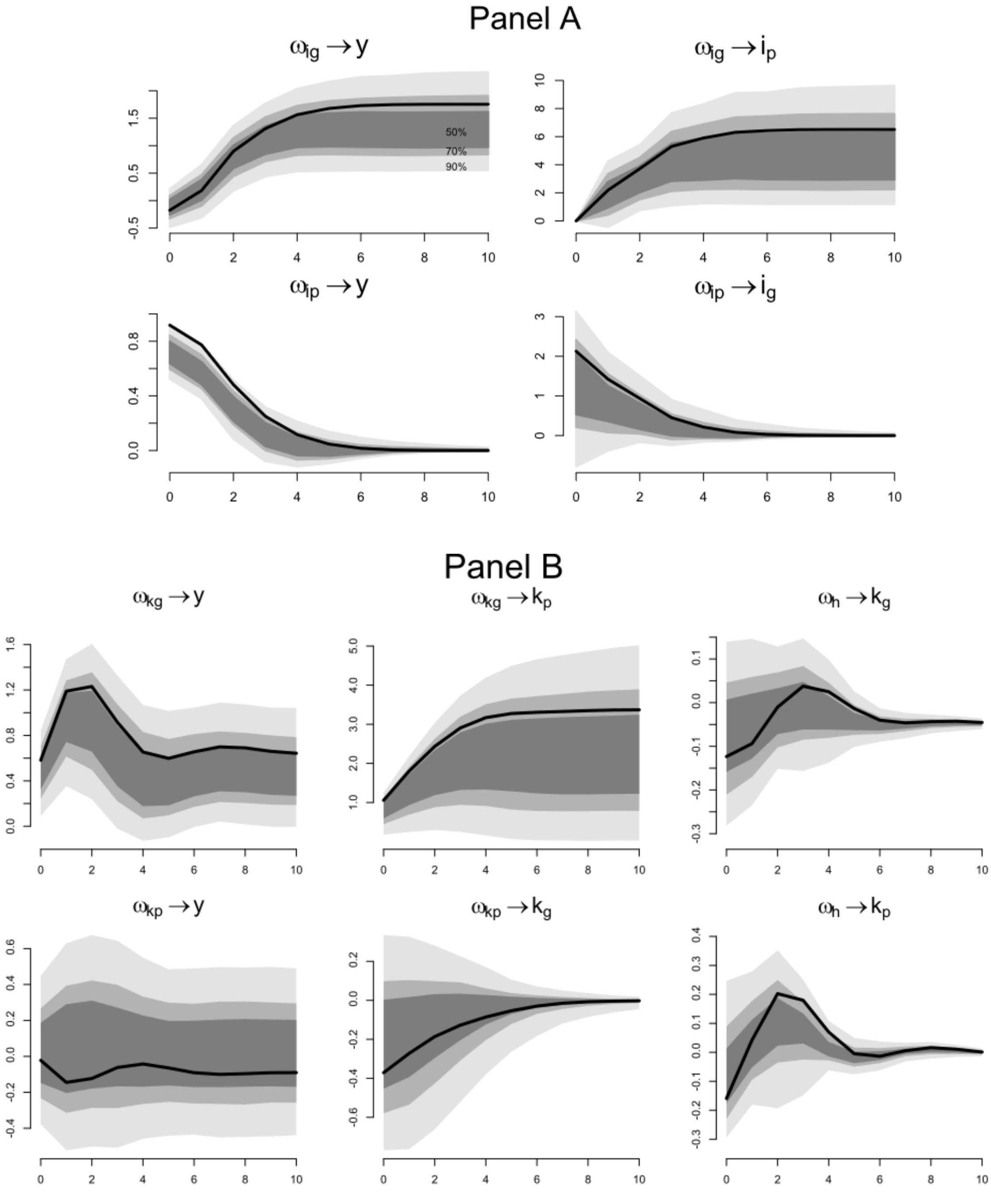


Figure 2: Responses of endogenous variables to structural shocks (*shock* → *var*). Shaded bands are 90%, 70%, and 50% (from lightest to darkest, respectively) bootstrapped error bands with 2000 replications. Y-axis is in terms of percentage, while X-axis represents the forecasting horizon. Panel A refers to selected IRFs of model I, whilst Panel B refers to that of model II.

Crucially, public (investment) capital seems significantly to crowd-in private (investment) capital, as depicted in the plots (b) and (f) of Panels A and B in *Figure 3*. It has been shown that an increase in public (investment) capital equivalent to a 0.8% (11.8%) shock induces 3% (6%) of growth in the private sector. Moreover, public investment spurs a steady growth of 1.5% in the output level from the fourth year onward, whereas the response of output to public capital shock reaches its peak of about 1.25% in the second year and then statistically vanishes to 0.5% in year five and beyond. Regarding employment, I find that shocks to both private and public capital stock positively affect labour supply, but the estimates are somewhat imprecise and insignificant.[5] The third column of Panel B shows that an increase in labour supply initially decreases public (private) capital stock, however, these impacts are not long-lasting as equilibria tend to recover within a four-year horizon. Note that, private capital seems to react more quickly and strongly than its public counterpart to labour force shocks, probably because of the longer implementation time lags of public investment projects.

Table 9 shows the forecast error variance decomposition of the variables under evaluation. It is easily seen that public investment predominantly explains the variations of private investment in both the short and long run, accounting for about 90% and 74% of the first and tenth forecasting horizons, respectively. Contrariwise, private capital seems to have an equally contributing variance of public capital and output in the middle and long run, although the former shock gives a higher contribution to the private capital's variability in the first three years. In the short run, output variance is mostly described by private investment and employment in Models I and II, at 96.4% and 72.2%, respectively, but public investment and public capital make a major contribution, at about 60%, to output variance in the long run. Interestingly, VN employment rate variations have been totally caused by private capital innovations.

In a nutshell, I empirically answer two questions proposed at the outset. First, I find that public investment and public capital stock have had positive impacts on VN output and private capital stock accumulation, reflecting the 'so-called' 'crowding-in effect'. Secondly, there is a Granger causality circle among the three variables, K_g, Y, and K_p, but public capital seems to *Granger*-cause private capital indirectly via income growth channel. Finally, it is shown that the positive effect of public capital on

[5] The IRFs are not shown but available upon request.

employment is statistically insignificant, but private capital shocks appear to exhibit the reverse effects on employment.

Variable	Shock Horizon	Model I ω^{y}	Model I ω^{ig}	Model I ω^{ip}	Model II ω^{kg}	Model II ω^{kp}	Model II ω^{h}	Model II ω^{y*}
y	1	0.05	3.52	96.43				
	3	11.11	30.52	58.37				
	5	18.59	60.45	20.96				
	10	20.88	72.77	6.35				
i_g	1	2.08	94.82	3.10				
	3	3.79	94.13	2.08				
	5	4.96	93.87	1.16				
	10	5.93	93.55	0.52				
i_p	1	90.01	0.00	9.99				
	3	85.59	7.66	6.75				
	5	79.64	16.84	3.52				
	10	73.94	24.61	1.45				
k_g	1				31.51	7.39	0.29	60.81
	3				59.34	2.37	0.09	38.21
	5				77.30	1.15	0.09	21.46
	10				91.45	0.42	0.04	8.09
k_p	1				59.21	1.41	1.34	38.04
	3				54.29	1.84	0.36	43.51
	5				51.93	1.62	0.19	46.26
	10				47.86	1.57	0.06	50.51
h	1				10.19	84.74	1.14	3.93
	3				16.08	81.32	0.40	2.20
	5				17.70	80.65	0.24	1.41
	10				20.34	78.86	0.12	0.69
$y*$	1				26.18	0.04	72.23	1.56
	3				60.40	0.68	23.89	15.03
	5				57.75	0.54	18.03	23.69
	10				58.45	0.71	12.41	28.43

"*" denotes output in Model II. All numbers are in terms of percentage.

Table 9: Variance decomposition of Model I and II.

4.8 Conclusions and implications

There is a mixed of findings on the influence of public investment and public capital on private investment and economic growth are mixed, especially in the empirical literature on developing countries (see *e.g.*, Bahal *et al.*, 2018; Dinh Thanh and Canh, 2019; Shanmugam, 2017; Xu and Yan, 2014). This study investigates the dynamic relationships between public investment, public capital, private investment and economic development in Vietnam over the forty years of economic transformation towards a market-friendly economy, 1976 – 2015.

I employ two standard error correction models with dummy variables to deal with structural changes in the Vietnam economy. My estimations show that public (investment) capital appears to have positive impacts on private (investment) capital and output growth in Vietnam. This implies the existence of crowding-in effects of public (investment) capital in Vietnam over the past four decades. Notably, most variations in the employment rate are explained by private capital accumulation, suggesting that private investment is important for creating jobs and tackling unemployment.

These results have two main policy implications. First, they suggest that public investment is an appropriate strategy for Vietnam, given that its economy has only just escaped the 2008 economic turmoil and that the middle-income trap threatens it (Aiyar et al., 2013). The government could thus promote the development of the private sector by channelling funds to essential infrastructure projects such as new bridges, highways, mass transport systems *etc*. The positive influence of public investment on private investment would be very significant for the labour market, since private investment has been shown to be a main driver of variations in the employment rate. However, public investment should be undertaken under the best possible public governance to intensify its efficacy.

Second, as suggested by Barker and Üngör (2019), Vietnam economy should become less dependent on factor accumulation as its source of growth and should raise its technological capability to boost productivity growth in both agricultural and non-agricultural sectors. That is, the threat of productivity stagnation should be considered in the Vietnam government's investment agenda. The literature on endogenous growth has emphasised the strong link between technological progress, innovation, and growth (Grossman and Helpman, 1991; King and Levine, 1993). This implies that public investment has also a key role to play in support of innovative industries which should shape Vietnam's economy towards an entrepreneurial economy.

Reference

Abbott, P. and Tarp, F. (2012) ‘Globalization Crises, Trade And Development In Vietnam’, Journal of International Commerce, Economics and Policy, 03(01), p. 1240006.

Abel, A. B. (2017) ‘Crowding out in Ricardian economies’, Journal of Monetary Economics, 87, pp. 52–66.

Abiad, M. A., Furceri, D. and Topalova, P. (2015) The macroeconomic effects of public investment: evidence from advanced economies. International Monetary Fund.

Afonso, A. and St. Aubyn, M. (2009) ‘Macroeconomic rates of return of Public and Private investment: Crowding-in and Crowding-out effects’, The Manchester School, 77, pp. 21–39.

Afonso, A. and St. Aubyn, M. (2019) ‘Economic growth, public, and private investment returns in 17 OECD economies’, Portuguese Economic Journal, 18(1), pp. 47–65.

Agénor, P.-R. and Neanidis, K. C. (2015) ‘Innovation, public capital, and growth’, Journal of Macroeconomics, 44, pp. 252–275.

Aiyar, M. S. et al. (2013) Growth slowdowns and the middle-income trap. International Monetary Fund.

Anh, V. T. T. (2016) ‘Vietnam: decentralization amidst fragmentation.’, Journal of Southeast Asian Economies, pp. 188–208.

Anwar, S. and Nguyen, L. P. (2011a) ‘Foreign direct investment and export spillovers: Evidence from Vietnam’, International Business Review, 20(2), pp. 177–193.

Anwar, S. and Nguyen, L. P. (2011b) ‘Foreign direct investment and trade: The case of Vietnam’, Research in International Business and Finance, 25(1), pp. 39–52.

Arouri, M., Youssef, A. Ben and Nguyen, C. (2017) ‘Does urbanization reduce rural poverty? Evidence from Vietnam’, Economic Modelling, 60, pp. 253–270.

Aschauer, D. A. (1989a) ‘Does public capital crowd out private capital?’, Journal of Monetary Economics, 24(2), pp. 171–188. doi: 10.1016/0304-3932(89)90002-0.

Aschauer, D. A. (1989b) ‘Is public expenditure productive?’, Journal of Monetary Economics, 23(2), pp. 177–200.

Atukeren, E. (2005) ‘Interactions Between Public and Private Investment: Evidence from Developing Countries’, Kyklos, 58(3), pp. 307–330.

Bahal, G., Raissi, M. and Tulin, V. (2018) ‘Crowding-out or crowding-in? Public and private investment in India’, World Development, 109, pp. 323–333.

Barker, T. and Üngör, M. (2019) ‘Vietnam: The next asian Tiger?’, The North American Journal of Economics and Finance, 47, pp. 96–118.

Barro, R. J. and Sala-i-Martin, X. (1992) ‘Public Finance in Models of Economic Growth’, Review of Economic Studies, 59(4), pp. 645–661.

Blanchard, O. J. and Quah, D. (1989) ‘The Dynamic Effects of Aggregate Demand and Supply Disturbances’, The American Economic Review, 79(4), pp. 655–673.

Breitung, J., Brüggemann, R. and Lütkepohl, H. (2004) ‘Structural vector autoregressive modeling and impulse responses’, Applied time series econometrics. Cambridge University Press.

Buiter, W. H. (1977) ‘‘Crowding out’and the effectiveness of fiscal policy’, Journal of public economics, 7(3), pp. 309–328.

Caldara, D. and Kamps, C. (2017) ‘The analytics of SVARs: A unified framework to measure fiscal multipliers’, Review of Economic Studies, 84(3), pp. 1015–1040.

De Jong, J F M, Ferdinandusse, M., and Funda, J. (2018). ‘Public Capital in the 21st Century: As Productive as Ever?’, Applied Economics, 50 (51): 5543–60.

Dickey, D. A. and Fuller, W. A. (1979) ‘Distribution of the estimators for autoregressive time series with a unit root’, Journal of the American statistical association, 74(366a), pp. 427–431.

Dinh Thanh, S., and Canh, N. P. (2019). ‘Dynamics between government spending and economic growth in China: an analysis of productivity growth’, Journal of Chinese Economic and Business Studies, 1-24.

Dollar, D. (2002) Reform, growth, and poverty in Vietnam. The World Bank.

Dreger, C. and Reimers, H.-E. (2016) ‘Does public investment stimulate private investment? Evidence for the euro area’, Economic Modelling, 58, pp. 154–158.

Engle, R. F. and Granger, C. W. J. (1987) ‘Co-integration and error correction: representation, estimation, and testing’, Econometrica: journal of the Econometric Society, pp. 251–276.

Erden, L. and Holcombe, R. G. (2005) ‘The effects of public investment on private investment in developing economies’, Public Finance Review, 33(5), pp. 575–602.

Erenburg, S. J. (1993) ‘The real effects of public investment on private investment’, Applied Economics, 25(6), pp. 831–837.

Everaert, G. (2003) ‘Balanced growth and public capital: An empirical analysis with I(2) trends in capital stock data’, Economic Modelling, 20(4), pp. 741–763.

Facchini, F. and Melki, M. (2013) ‘Efficient government size: France in the 20th century’, European Journal of Political Economy, 31, pp. 1–14.

Fernald, J. G. (1999) ‘Roads to prosperity? Assessing the link between public capital and productivity’, American Economic Review, 89(3), pp. 619–638.

Fisher, A. C. et al. (2016) ‘Progressive taxation, public capital, and endogenous growth’, International Review of Applied Economics, 24(1), pp. 35–47.

Fritzen, S. (2002) ‘Growth, inequality and the future of poverty reduction in Vietnam’, Journal of Asian Economics, 13(5), pp. 635–657.

Ganelli, G. (2003) ‘Useful government spending, direct crowding-out and fiscal policy interdependence’, Journal of international money and finance, 22(1), pp. 87–103.

Gonzalo, J. and Ng, S. (2001) ‘A systematic framework for analyzing the dynamic effects of permanent and transitory shocks’, Journal of Economic Dynamics and Control, 25(10), pp. 1527–1546.

Granger, C. W. J. (1981) ‘Some properties of time series data and their use in econometric model specification’, Journal of econometrics, 16(1), pp. 121–130.

Grossman, G. M. and Helpman, E. (1991) Innovation and growth in the global economy. MIT press.

Hendry, D. F. and Juselius, K. (2000) ‘Explaining Cointegration Analysis: Part 1’, The Energy Journal. International Association for Energy Economics, 21(1), pp. 1–42.

Hendry, D. F. and Juselius, K. (2001) ‘Explaining Cointegration Analysis: Part II’, The Energy Journal. International Association for Energy Economics, 22(1), pp. 75–120.

Hunt, C. (2012) ‘The interaction of public and private capital: a study of 20 OECD members’, Applied Economics. Routledge, 44(6), pp. 739–764.

Johansen, S. (1995) Likelihood-based inference in cointegrated vector autoregressive models. Oxford University Press.

Kamps, C. (2005) ‘The Dynamic Effects of Public Capital: VAR Evidence for 22 OECD Countries’, International Tax and Public Finance, 12(4), pp. 533–558.

Kandil, M. (2017) ‘Crowding out or crowding in? Correlations of spending components within and across countries’, Research in International Business and Finance, 42, pp. 1254–1273.

Kang, W. and Imai, K. S. (2012) ‘Pro-poor growth, poverty and inequality in rural Vietnam’, Journal of Asian Economics, 23(5), pp. 527–539.

Kilian, L. and Lütkepohl, H. (2017) Structural vector autoregressive analysis. Cambridge University Press.

King, R. G. and Levine, R. (1993) ‘Finance and Growth: Schumpeter Might Be Right’, The Quarterly Journal of Economics, 108, pp. 717–737.

King, R. G., Plosser, C. I. and Stock, J. H. (1991) ‘Stochastic Trends and Economic Fluctuations’, The American Economic Review, 81(4), pp. 819–840.

Kwiatkowski, D. et al. (1992) ‘Testing the null hypothesis of stationarity against the alternative of a unit root: How sure are we that economic time series have a unit root?’, Journal of econometrics, 54(1–3), pp. 159–178.

Lanjouw, P. and Marra, M. R. (2018) ‘Urban poverty across the spectrum of Vietnam’s towns and cities’, World Development, 110, pp. 295–306.

Le, T.-H. and Tran-Nam, B. (2018) ‘Relative costs and FDI: Why did Vietnam forge so far ahead?’, Economic Analysis and Policy, 59, pp. 1–13.

Leung, S. E. (2010) ‘Vietnam: an economic survey’, Asian-Pacific Economic Literature, 24(2), pp. 83–103.

Lucas Jr., R. E. (1988) ‘On the mechanics of economic development’, Journal of Monetary Economics, 22(1), pp. 3–42.

Lütkepohl, H. (2005) New introduction to multiple time series analysis. Springer Science & Business Media.

Magrini, E., Montalbano, P. and Winters, L. A. (2018) ‘Households’ vulnerability from trade in Vietnam’, World Development, 112, pp. 46–58.

Mallon, R. and Irvin, G. (2001) ‘Systemic change and economic reform in Vietnam’, in Globalization and Third-World Socialism, pp. 153–167.

Musolesi, A. (2011) ‘On public capital hypothesis with breaks’, Economics Letters, 110(1), pp. 20–24.

Nguyen, D. T. H., Sun, S. and Anwar, S. (2017) ‘A long-run and short-run analysis of the macroeconomic interrelationships in Vietnam’, Economic Analysis and Policy, 54, pp. 15–25.

Nguyen, L. P. and Anwar, S. (2011) ‘Fiscal decentralisation and economic growth in Vietnam’, Journal of the Asia Pacific Economy, 16(1), pp. 3–14.

Ogibayashi, S. and Takashima, K. (2017) ‘Influence of Inefficiency in Government Expenditure on the Multiplier of Public Investment’, Computational Economics, 50(4), pp. 549–577.
Ohno, K. (2009) ‘Avoiding the middle-income trap: renovating industrial policy formulation in Vietnam’, ASEAN Economic Bulletin. ISEAS--Yusof Ishak Institute, 26(1), pp. 25–43.
Ortigueira, S. and Santos, M. S. (1997) ‘On the Speed of Convergence in Endogenous Growth Models’, American Economic Review, 87(3), pp. 383–399.
Peacock, A. and Scott, A. (2000) ‘The curious attraction of Wagner’s law’, Public Choice. Springer, 102(1–2), pp. 1–17.
Pereira, A. M. (2001) ‘On The Effects of Public Investment on Private Investment: What Crowds in What?’, Public Finance Review, 29(1), pp. 3–25.
Pereira, A. M. and Roca-Sagales, O. (2001) ‘Infrastructures and private sector performance in Spain’, Journal of Policy Modeling, 23(4), pp. 371–384.
Perron, P. (1989) ‘The great crash, the oil price shock, and the unit root hypothesis’, Econometrica: Journal of the Econometric Society, pp. 1361–1401.
Pincus, J., Anh, V. T. T. and Le Thuy, T. (2008) ‘Vietnam: A tiger in turmoil’, Far Eastern Economic Review, 171(4), p. 28.
Romer, P. M. (1994) ‘The origins of endogenous growth’, Journal of Economic perspectives, 8(1), pp. 3–22.
Romero-Ávila, D. and Strauch, R. (2008) ‘Public finances and long-term growth in Europe: Evidence from a panel data analysis’, European Journal of Political Economy, 24(1), pp. 172–191.
Shanmugam, M. (2017). ‘Does public investment crowd-out private investment in India’, Journal of Financial Economic Policy, 9(1), 50-69.
Sims, C. A. (1980) ‘Macroeconomics and Reality’, Econometrica. [Wiley, Econometric Society], 48(1), p. 1.
Sturm, J. E. and de Haan, J. (1995) ‘Is public expenditure really productive?. New evidence for the USA and The Netherlands’, Economic Modelling, 12(1), pp. 60–72.
Su, T. D. and Bui, T. M. H. (2017) ‘Government size, public governance and private investment: The case of Vietnamese provinces’, Economic Systems, 41(4), pp. 651–666.
Thanh, V. T. and Duong, N. A. (2009) ‘Vietnam after two years of WTO accession: What lessons can be learnt?’, ASEAN Economic Bulletin. ISEAS--Yusof Ishak Institute, 26(1), pp. 115–135.
Vogelsang, T. J. and Perron, P. (1998) ‘Additional tests for a unit root allowing for a break in the trend function at an unknown time’, International Economic Review, pp. 1073–1100.
Voss, G. M. (2002) ‘Public and private investment in the United States and Canada’, Economic Modelling, 19(4), pp. 641–664.
Xu, X. and Yan, Y. (2014), ‘Does government investment crowd out private investment in China?’, Journal of Economic Policy Reform, 17(1), 1-12.
Zivot, E. and Andrews, D. W. K. (2002) ‘Further evidence on the great crash, the oil-price shock, and the unit-root hypothesis’, Journal of Business and Economic Statistics, 20(1), pp. 25–44.

CHAPTER 5: Implications of public expenditures on Vietnam economy♠

Abstract

This paper explores the influence of government expenditures on the business cycle dynamics of a transition country such as Vietnam. I develop a stochastic-growth DSGE/RBC model based on Leeper et al (2010a; 2010b), with notable features such as government consumption in utility, internal habit persistence in consumption, private capital utilisation, and consideration of two fiscal shocks in addition to shocks to household preferences and technological progress. The estimated DSGE model adequately captures the dynamics of Vietnam's economy and is helpful to uncover a diversity of findings. First, public investment has been an important driver of Vietnam's high output growth in the early-to-mid 1990s, although its contribution decreased during the 2000s. Second, government expenditures explain up to 30% and 20% of Vietnamese output variations in the long and short run, respectively. Third, impulse–response functions reveal a 3% GDP cumulative five-year gain from a public investment shock of 1% GDP. Fourth, simulation analysis uncovers a significant and positive impact of productive public investment by crowding in private investment.

Keywords: Vietnam, public investment, crowding-in, DSGE, business cycle.

JEL codes: EE62, H11, H54, O11.

♠ This chapter has awarded the best paper of the Asia Conference on Business and Economic Studies (ACBES 2019) organised by the Journal of Asian Business and Economic Studies, and sponsored by University of Economics Ho Chi Minh City, held on 13^{th} – 14^{th} Sep 2019 at Ho Chi Minh City, Vietnam.

5.1 Introduction

Being one of the fastest growing economies in Asia, Vietnam is subject to congestion in transport infrastructures and basic supplies (e.g., education and public health service, electricity, and tap water). For example, in the two biggest cities, Ho Chi Minh and Ha Noi, traffic jams put the burden of billion dollars per year on its citizens. On a countrywide level, Vietnam suffers from delays in the north-south national highway, which is a most expected project but will not be completed before 2022. Current congestion and delays are certainly costly in general, as the literature has shown, and also in Vietnam, given that the numerous infrastructure projects developed since the 'Doi Moi' (an economic renovation program commenced in 1986) have sustained Vietnam's rapid economic growth for three consecutive decades by crowding in private investment and output (Pham et al., 2019b).

In spite of the undeniable role played by public expenditures in a transition economy such as Vietnam, there is hardly any empirical evidence exploring the interactions between its economic fluctuations and public expenditures. Pham *et al.* (2019a) was perhaps the first real business cycle (RBC) analysis of the Vietnamese (hereafter VN) economy over the post-war period, 1980 – 2015. Their small open economy setting, however, leaves aside the role of public capital spending and government final consumption in explaining VN economic fluctuations. Thus, this paper fills a void in the literature and complements the main findings in Pham *et al.* (2019a).

Although New Keynesian (NK) literature was not much concerned on the relationship between government fiscal policy and economic growth until the quake of the 2008 global financial crisis (GFC), this crisis has lured substantial quantitative research on this nexus. One strand of the fiscal literature preferably employs vector autoregression models showing mixed evidence. For instance, Abiad *et al.* (2016) showed that public investment could raise output in both the short- and long-term and reduce unemployment, which in turn crowds in private investment (Pereira and Andraz, 2013; Dreger and Reimers, 2016) and raises the marginal productivity of private capital (Cavallo and Daude, 2011). Other studies show, however, that excessive and inefficient public investment may crowd out private investment and thus affect income growth negatively (Abel, 2017; Cavallo and Daude, 2011; Kandil, 2017; Ogibayashi and Takashima, 2017; among others). On a different perspective, Ilzetzki *et al.* (2013) documented a larger output multiplier associated to government spending in developed (closed) economies than in developing (open) economies. This multiplier was found to be zero in countries having flexible exchange rates, large in those having predetermined exchange rates, and negative in highly indebted countries.

Another strand of literature focuses on dynamic stochastic general equilibrium (DSGE) models. Early contributions, such as Baxter and King (1993), claimed that permanent changes in government expenditures may result in larger short- and long-run output multipliers, and that permanent changes seem to be more important than the temporary changes, thereby they stressed on the remarkable impacts of public investment on private output and investment. Christiano and Eichenbaum (1992) analysed the influences of government consumption on labour-market dynamics and showed that modelling government consumption significantly improves the moment matching ability of the neo-classical general equilibrium model. More recently, Leeper et al. (2010a, b) extended the standard real business cycle (RBC) model by incorporating several NK features such as household preference persistence, investment adjustment costs, and a variety of exogenous shocks other than the conventional total-factor productivity shock. Specifically, their evidence suggests that the scheduling implementation of public investment is of utmost importance in the short run, while less productive public capital and distorting financing could harm government investment in the long run. Moreover, debt-financed fiscal shocks could have prolonged impacts, giving rise to notable differences in the short- and long-run fiscal multipliers.

Aiming to investigate the effects of government expenditures on the VN business cycles in the light of neo-classical economic theory, in this chapter I build and estimate a DSGE model, which is based on the ones in Leeper et al. (2010a, b). To this benchmark setting, I incorporate the presence of adjustment costs in private capital stock, following in CEE (2005); internal habit persistence in consumption, following Boldrin et al. (2001) and Pham et al. (2019a); government consumption in utility as in Baxter and King (1993); and private capital utilisation, which is a refinement with respect to the seminal model in King and Rebelo (1999). The specific target of my analysis is to comprehend two specific economic matters. First, the extent to which output volatility can be explained by variations in public expenditures. Second, the impact of new productive public capital investment on private income, private investment, and hours-work in a real business cycle sense. To conduct this analysis, I move beyond the consideration of standard shocks in TFP and consumption preferences and bring in the role of two fiscal shocks in the literature such as the ones on government final consumption and public investment. To the best of my knowledge, this is the first attempt to provide such a DSGE investigation for Vietnam's economy.

The DGSE RBC model fits to a large extent the cyclical aggregate moments in Vietnam. In the Bayesian analysis, the IRFs are quite revealing of the economic relevance of public expenditures. More precisely, I find that a one percentage point (pp) increase in public investment (as % of GDP) causes a 0.91 pp increase in income, and a 3.0 pp accumulated income gain in five years. It is important to remark that a 1 pp increase in public investment as % of GDP is roughly equivalent to 10% of total public investment in absolute values.

With respect to the variance decomposition analysis, three periods are worth examining. First, 1985-1996, triggered by the Doi Moi, the Collapse of Communist Bloc, and the 1992 Constitutional law. Second, 1997-2007, triggered by the East Asian Crisis. Third, 2008-2015, triggered by the GFC. In the first of these three periods, TFP shocks dominate output fluctuations, followed by shocks in public investment. In the second one, TFP shocks still dominate, followed by the shock in consumption preferences. In the third period, TFP remain as the most significant driver of economic fluctuations, but there is again a substantial impact of shocks in public investment (which is in turn similar to the one on consumption preferences). From this analysis, I posit that productive public expenditures are relevant and should be taken into account in any economic plan aiming to foster growth.

This is confirmed by the conditional and unconditional variance decomposition analyses. The conditional variance decomposition shows that TFP shocks account for 24.6% of output variation, preference shocks for 56.8%, and public investment shocks for 18.3%. In turn, the unconditional variance decomposition reveals that TFP shocks account for 38.3% of output variation, closely followed by preference and public investment shocks, which account, respectively, for 32.4% and 29.2% of output variation.

Given the relevance of public expenditures, but also their different role depending on the form taken (consumption or investment). Knowing, in addition, the existence of studies that point out the necessity of ensuring the efficiency of any public investment project, I conduct a deterministic simulation to compare the impact of two fiscal programmes with a different policy mix in terms of funding mechanisms and expenditure composition. In this setting, I show that in the event of an expansionary policy design matters and the efficiency of public projects is of paramount importance.

The remainder of this chapter is organized into five sections. In the next section, I economically summarise the stylised facts of VN business cycles. The DSGE model and its Bayesian estimation results are discussed in Section 3 and 4, respectively. I then provide fiscal policy implications in Section 5. The concluding remarks close the chapter.

5.2 Stylised facts

The Vietnamese (VN) real GDP per capita (hereafter Y) has been steadily growing from US$ 200 per capita in the 1980s to several times higher over the past three decades. Vietnam is now a lower-middle income country as the nominal income of VN people reached above US$ 2,300 per person in 2017, marking a great success of the economic renovation, 'Doi Moi', program. Figure 1 shows the upward trend of gross capital investment (black line) and public investment (grey bars) over the period 1986 – 2016. The latter investment apparently experienced two different phases. The first phase indicates that public investment constantly grew up from the low of 4% of GDP in 1986 to the peak of 21.4% in 2002, whereas it has diminished to the stable level of 12.3% in the second phase.

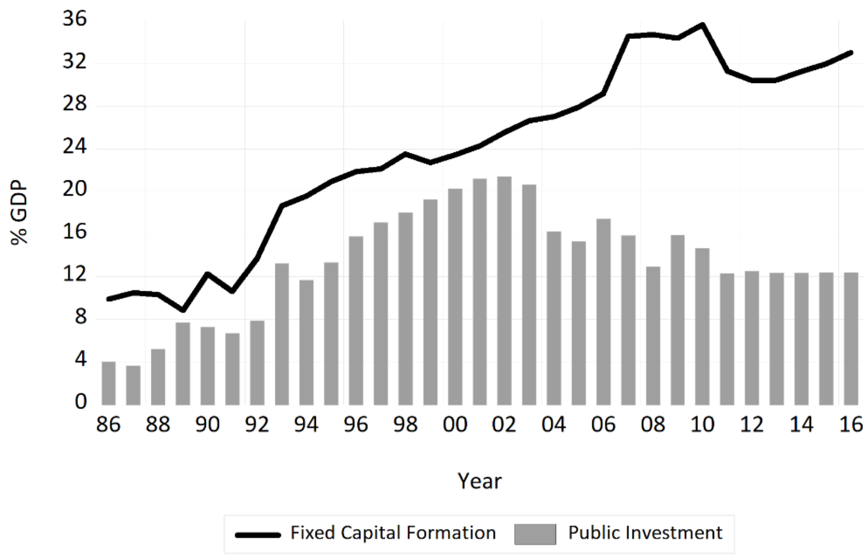


Figure 1: Vietnam's investments, 1986 – 2016.

The dashed-dotted lines in Figure 2 present the dynamic of Hodrick – Prescott (HP) detrended main aggregates in the national income identity equation. It has been shown that their movements are fairly close to the paths of respective growth variables (the black solid lines). Technically speaking, the business cycle or phenomenon (Prescott, 1986) is measured by examining dynamic characteristics of cyclical (detrended) components extracted from their original non-stationary levels. Pham *et al.* (2019a) extensively investigated Vietnam's economic regularities alongside with other ASEAN-5 partners, but they omitted two sub-investment variables, i.e. public investment ($\mathrm{I_g}$) and private investment ($\mathrm{I_p}$), as well as public capital and private capital stock series, $\mathrm{K_g}$ and $\mathrm{K_p}$, respectively. As the matter of facts, Table 1 complements the findings in Pham *et al.* (2019a), but the analysis is limited to the usage of standard HP filter with $\lambda = 100$.[1]

[1] I refer the interested readers to Pham, Sala, and Silva (2019a) for further details.

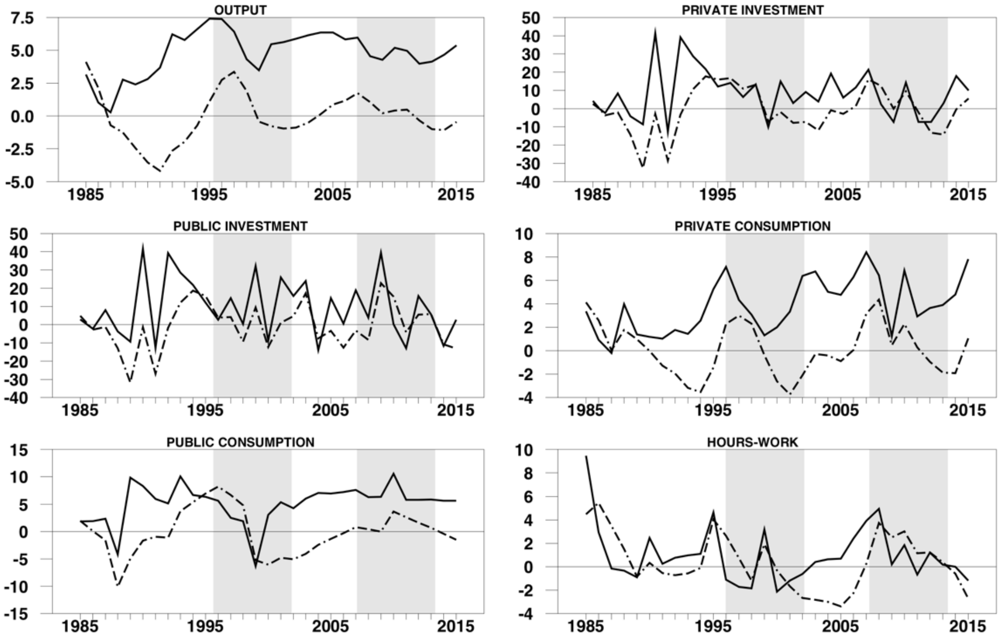


Figure 2: Dynamic of Vietnam main aggregate variables. Continuous lines are growth rates. Dashed-dotted lines are HP filtered ($\lambda = 100$) data.

As can be seen from the top-block of Table 1, the VN output growth rate sustained the high pace of 4.8% and been less volatile than other main aggregates such as consumption and investment over the past thirty years. On the sample average, capital stock and investment components grew faster than output as their gaps varying in the range of 2.3% – 4.7% but there are some notable differences if one has a closer look at the two recent economic crises, namely the Asian financial crisis (1997 – 2002) and the GFC (2008 – 2013). For instance, in the former crisis, the rates of public and private investment growth were about 13.3% and 6.1%, respectively, while their growth rates were only 8.62% and - 0.4%, respectively, in the second crisis. This fact signifies the crucial role of public investment spending in sustaining and stabilising the VN economy over the last two decades.

The cyclical output line appears to be the smoother curve among those of other plots in Figure 2 since its standard error is as low as 1.73% in comparison with around 2.4% of household consumption and hours-work variables, 3.8% of government consumption, 2.4% (3.5%) of public (private) capital, and above 11% of investment variables. This implies that public investment (I_g) and private investment (I_p) are the most fluctuating aggregate components.

Measure | Variables	Y	C_p	C_g	Inv	H	K_g	K_p	I_g	I_p
Growth, %	4.79	3.84	5.08	8.81	0.98	8.32	7.12	9.53	8.54
	(0.31)	(0.42)	(0.63)	(2.12)	(0.43)	(0.70)	(0.68)	(2.88)	(2.37)
HP filtered $\sigma_{(.)}$ / σ_y	1.00	1.18	2.23	5.87	1.29	1.35	2.05	6.71	6.60
		(0.11)	(0.35)	(1.18)	(0.18)	(0.26)	(0.36)	(0.85)	(0.98)
	Cross-correlation with Output								
Lag	Y	C_p	C_g	Inv	H	K_g	K_p	I_g	I_p
-3	0.001	0.264	-0.597	-0.420	0.114	0.501	0.692	-0.329	-0.388
-2	0.343	0.430	-0.227	-0.012	0.317	0.605	0.812	-0.047	-0.004
-1	0.737	0.597	0.218	0.290	0.427	0.620	0.754	0.099	0.309
0	1.000	0.629	0.497	0.574	0.380	0.512	0.514	0.272	0.599
1	0.737	0.231	0.553	0.646	0.147	0.223	0.162	0.410	0.639
2	0.343	-0.185	0.479	0.625	-0.098	-0.104	-0.170	0.492	0.581
3	0.001	-0.419	0.311	0.414	-0.352	-0.336	-0.390	0.372	0.368
	Autocorrelation								
Lag	Y	C_p	C_g	Inv	H	K_g	K_p	I_g	I_p
1	0.737	0.596	0.727	0.508	0.658	0.721	0.838	0.237	0.492
2	0.343	0.136	0.398	0.386	0.302	0.421	0.532	0.157	0.343
3	0.001	-0.16	0.082	0.065	0.006	0.106	0.164	-0.133	0.061

Standard error in parentheses estimated by GMM method.

Table 1: Vietnam's business cycle facts, 1985 – 2015.

The middle block of Table 1 provides cross-correlations across lags and leads of output with the rest. It is suggested that Vietnam economy business cycles subjected to high volatility because the correlations among output and other aggregates are mild and all ratios of aggregate standard errors over output above unity. The contemporaneous correlation between output and private consumption occupies the highest position despite being below two-third. The output adequately correlates with government consumption and investment leads, but it weakly ties up with their lags. Contrariwise, the leads of capital stock variables (K_g and K_p) seem to be unrelated to output but not their respective lags, showing the time-to-build effects. This means that past capital stock accumulation would significantly affect the current and future output. It is also evident that C_g, C_p, and I_g, I_p exhibit a weak pro-cyclical property.

The auto-correlation measure describes how persistent a time-series would be. In general, the persistence of public and private capital stocks, K_g and K_p, respectively, are higher than output and consumption. The public investment appears to be a strongly unstable process in the past, and it is almost unrelated to output movements since their contemporaneous and first-lag correlations are low, as of 0.272 and 0.099, respectively. It, however, has a moderate positive correlation with Y from the first lead to the third one, giving rise to the fact that public investment would affect the output in three years from the outset. This impact is stronger with respect to private investment, indicating the vital role of the private sector in the VN economy as analysed in Pham *et al.* (2019b).

5.3 Economic environment

5.3.1 Model

I build a zero-growth closed economy model comprising of three economic agents, namely household, firm and government. The representative household consumes one good, invests in private capital, and contributes taxes to the government. But he has received a lump-sum transfer from the government. Technically speaking, the economy is assumed to be symmetric and endowed with constant-returns-to-scale Cobb-Douglas production function so that firms are identical in terms of technological aspect.[2]

There are two roles of government in the economy: as the tax levying entity by imposing three kinds of taxes to households i.e. taxes on labour and capital income and consumption tax, and as the public input supplier. The latter role accumulates public capital by building up infrastructure. Since electricity, water, transport and telecommunication supplies are inputs in most industries, government practically fosters economic growth by making large investments in these sectors. And firms enjoy infrastructure

[2] I shall present different constant returns to scale definition later in the model development.

improvements so that they could earn extra profits by making use of public inputs. It is the Figure 3 which best demonstrates the interactions and flows in the modelled economy.

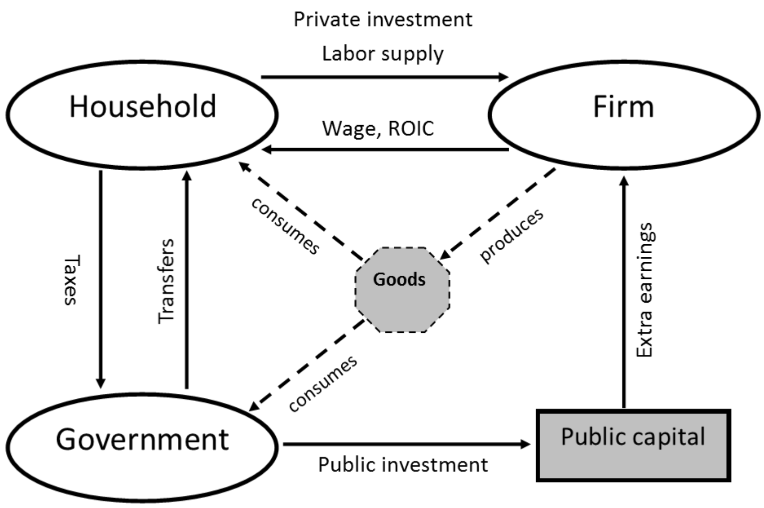


Figure 3: A bird-eyes views of the modelled economy

Besides, the model contains components reflecting the real frictions of the economy. It is worth considering capital-related rigidities such as capital utilization and quadratic-type of investment adjustment costs Christiano, Eichenbaum, and Evans (2005) (CEE for short). The persistence of household consumption has also been considered as it is recently a typical feature of NK-DSGE modelling due to empirically supporting facts (see e.g., Boldrin et al. 2001; Pham et al. 2019a).

There is a shock attached to household intertemporal consumption preference so that the time discount factor β will vary over time instead of a constant as in the earlier RBC models (e.g., King and Rebelo, 1999; Hansen, 1985). In addition, two shocks linking with government final consumption and public investment are incorporated since I want to perturb the government fiscal schedule.

Households,

With additively logarithmic utility function, the representative household seeks to maximize the expected value function

$$\max E_0 \sum_{t=0}^{\infty} \beta^t z_{p,t} \left[\log(C_{p,t} - hC_{p,t-1} + \pi C_{g,t}) - \chi \frac{H_t^{1+\sigma_L}}{1+\sigma_L}\right] \quad (1)$$

where β and $z_{p,t}$ are time discount factor and intertemporal preference shock, respectively. H_t represents labor supply in terms of hours-work; $\eta = \frac{1}{\sigma_L}$ is Frisch labor elasticity; h is internal habit formation coefficient of the household. $C_{p,t}$ denotes private consumption, whilst $C_{g,t}$ is government purchases exogenously as given. Time endowment is normalized so that $H_t + L_t = 1$, with L_t denoting leisure time. χ is dis-utility labor parameter to be calibrated such that H_t satisfies empirical hours-work. The preference shifter $z_{p,t}$ is assumed following an AR(1) process such that $\log z_{p,t} = \rho_p \log z_{p,t-1} + \sigma_p \epsilon_{p,t}$ with $\epsilon_p \sim N(0,1)$.

The involvement of government consumption and public capital in RBC model have been well-developed in Aschauer (1985, 1989), Aiyagari, et al. (1992), and Baxter and King (1993) among others. Baxter and King (1993) did specify a function $\Gamma(C_{g,t}, K_{g,t-1})$ with K_g denoting public capital stock, although they argued that K_g has no direct influence on household's decisions. The variable C_g is consequently assumed as an uncontrollable stochastic process (Christiano and Eichenbaum, 1992). Parameter π governs the elasticity of substitution between private and public consumption. Basically, there is no condition upon the sign of π, yet the setting of $\pi \geq 0$ implies government purchases crowding out private consumption until $\pi = 1$.

The household only consumes on what he earns so that his budget should be even in every period. By assuming physical capital owned by household, the intertemporal budget constraint is read

$$(1 + \tau_c)C_{p,t} + I_{p,t} + B_t =$$

$$= R_{b-1}B_{t-1} + (1 - \tau_w)W_t H_t + (1 - \tau_k)r_{k,t}u_t K_{p,t-1} + G_t + (1 - \tau_k)\text{Profits}_t \quad (2)$$

where W_t and $K_{p,t}$ denote real wage and private capital, respectively. Household pays consumption tax τ_c, capital gain tax (or returns on capital rental, τ_k) and labor tax bracket τ_w, yet he would receive a lump-sum transfers G_t from the government along with net extraordinary profits due to free access to public infrastructures. B_t is a financial asset in terms of unit of consumption issued by the government, which yields the gross return R_b.

Household chooses the optimal rate of capital service u_t in each period and level of investment, which are conditionally upon costs of over capital utilisation and costs of investment congestion.

$$K_{p,t} = (1 - \delta_p - \delta(u_t))K_{p,t-1} + I_{p,t}\left[1 - S\left(\frac{I_{p,t}}{I_{p,t-1}}\right)\right] \tag{3}$$

where $S(\frac{I_{p,t}}{I_{p,t-1}})$ is a function representing investment adjustment costs, which commonly has a quadratic form as such $S(\frac{I_{p,t}}{I_{p,t-1}}) = \frac{\kappa}{2}(\frac{I_{p,t}}{I_{p,t-1}} - 1)^2$, with $\kappa = \frac{\psi}{1-\psi}$ defined in Smets and Wouters (2003) (henceforth SW). In the steady-state, adjustment costs do not exist, one therefore claim that $S(1) = 0$, $S'(1) = 0$ and $S''(1) > 0$.

Parameter δ_p denotes depreciation rate of private capital. The function $\delta(u)$ represents the rate of accelerated capital depreciation, hence it should be an increasing function with respect to u_t such that $\delta'(u_t) > 0$, $\delta''(u_t) > 0$. Adopting the quadratic form: $\delta(u) = \delta_1(u_t - 1) + \frac{\delta_2}{2}(u_t - 1)^2$ similar to Schmitt-Grohé and Uribe (2012), it thus gives $\delta(\overline{u}) = \delta(1) = 0$.

Household maximizes (1) by choosing the sequence $\{C_{p,t}, H_t, K_{p,t}, u_t\}_{t=0}^{\infty}$ subject to (2) and (3) taking as given other stochastic processes and suitable initial conditions. The Lagrangian function can be written:

$$\begin{aligned}\mathcal{L} \quad &\equiv \max E_0 \sum_{t=0}^{\infty} \beta^t \{ z_{p,t}[\log(C_{p,t} - hC_{p,t-1} + \pi C_{g,t}) - \chi \frac{H_t^{1+\sigma_L}}{1+\sigma_L}] \\ &-\lambda_t[(1+\tau_c)C_{p,t} + I_{p,t} - ((1-\tau_w)W_t H_t + (1-\tau_k)r_{k,t}u_t K_{p,t-1} + G_t + (1-\tau_k)\text{Profits}_t)] \\ &-\mu_t[K_{p,t} - (1-\delta_p - \delta(u_t))K_{p,t-1} - I_{p,t}(1 - S(\frac{I_{p,t}}{I_{p,t-1}}))] \}\end{aligned}$$

Defining the ratio $q_t = \frac{\mu_t}{\lambda_t}$ as the Tobin´s Q, it could be understood as "the benefit from investment per unit of benefit from capital" Wickens (2012, p37). Solving the Lagrangian first-order conditions defines the household's optimal choices in the equilibrium at time index $t = 0$ as

$$q_t = \frac{(1-\tau_k)r_{k,t}}{\delta'(u_t)} = \frac{(1-\tau_k)r_{k,t}}{\delta_1 + \delta_2(u_t - 1)}$$

$$\frac{(1+\tau_c)z_{p,t}\chi H_t^{\sigma_L}}{(1-\tau_w)W_t} = z_{p,t}\left(C_{p,t} - hC_{p,t-1} + \pi C_{p,t}\right)^{-1} - \beta h E_t[z_{p,t+1}\left(C_{p,t+1} - hC_{p,t} + \pi C_{g,t+1}\right)^{-1}]$$

$$1 \quad = q_t[1 - \frac{\kappa}{2}(\frac{I_{p,t}}{I_{p,t-1}} - 1)^2 - \kappa(\frac{I_{p,t}}{I_{p,t-1}} - 1)\frac{I_{p,t}}{I_{p,t-1}}] + \beta E_t q_{t+1}\kappa(\frac{I_{p,t+1}}{I_{p,t}} - 1)(\frac{I_{p,t+1}}{I_{p,t}})^2[\frac{z_{p,t+1}W_t}{z_{p,t}W_{t+1}}(\frac{H_{t+1}}{H_t})^{\sigma_L}]$$

$$q_t \quad = \beta E_t[\frac{z_{p,t+1}W_t}{z_{p,t}W_{t+1}}(\frac{H_{t+1}}{H_t})^{\sigma_L}][q_{t+1}(1 - \delta_p - \delta(u_{t+1})) + (1-\tau_k)r_{k,t+1}u_{t+1}]$$

$$R_b = \frac{\lambda_t}{\beta\lambda_{t+1}}$$

It is clear that if $S(\cdot) = 0$, then $q_t = 1$ implying $\lambda_t = \mu_t$ and $\delta'(u_t) = (1-\tau_k)r_{k,t}$ so that the relationship between δ_1 and capital rental rate r_k is known in the stationary state. The Tobin's Q definition shows that the household only earns positive benefit $(q > 1)$ from utilisation of his own capital if the rate of returns (net of capital tax) is greater than accelerating wearing-out-rate of that capital. Hence, accelerating depreciation rate δ_1 should be greater than or equal to the normal depreciation rate of private capital δ_p in the modelled economy. This exposition will be validated in the next sub-section.

Firms,

I assume identical firms exhibit the Cobb-Douglas production technology as

$$Y_t = z_{a,t}(u_t K_{p,t-1})^{\alpha_1} K_{g,t-1}^{\alpha_2} H_t^{\alpha_3}$$

where K_p denotes public capital such as infrastructures, public utilities, etc. The productivity process z_a is assumed following an AR(1) such that $\log \mathrm{z_a} = \rho_a \log \mathrm{z_{a,t-1}} + \sigma_a \epsilon_{a,t-1}$ and $\epsilon_a \sim \mathrm{N}(0,1)$.

If one assumes the production technology exhibits constant-returns-to-scale (CRS) with respect only to private factors as in Baxter and King (1993) and Leeper *et al.* (2010a, b), then $\alpha_1 + \alpha_2 + \alpha_3 \geq 1$ implying increasing returns to scale with respect to all input factors. But it is not the only case found in the literature. Instead, I am modelling and estimating two settings. The second is to constrain $\alpha_1 + \alpha_2 + \alpha_3 = 1$ implying decreasing returns to scale of private factors because labour and private capital are paid more than their marginal products, i.e. $\alpha_1 + \alpha_3 < 1$. Notice that, firms always enjoy extraordinary profits with factor α_2 in both settings.

Firm management optimizes period-profits by setting level of physical capital and labour input rented from households. Profit function is thus written

$$\Pi_t = z_{a,t}(u_t K_{p,t-1})^{\alpha_1} K_{g,t-1}^{\alpha_2} H_t^{\alpha_3} - r_{k,t} K_{p,t-1} - W_t H_t \tag{5}$$

In perfect competitive environment, firm makes no profit with respect to private factors. That is marginal returns on private capital and labour input equal zeroes. Because of capital utilisation already chosen by the firm's owners, firm's optimal decisions are the first-order conditions of equation (4). They are

$$\begin{aligned} \frac{\partial \Pi_t}{\partial H_t} = 0 \Rightarrow W_t &= \alpha_3 \frac{Y_t}{H_t} \\ \frac{\partial \Pi_t}{\partial K_{p,t-1}} = 0 \Rightarrow r_{k,t} &= \alpha_1 \frac{Y_t}{K_{p,t-1}} \end{aligned}$$

Since there is no market price for public input, firm makes extra profits by an amount of $\mathrm{Profits}_t = \frac{\partial \Pi_t}{\partial K_{g,t-1}} = \alpha_2 \frac{Y_t}{K_{g,t-1}}$, which then returns to household net of capital gain tax, as that of $(1 - \tau_k)\mathrm{Profits}$.

Government,

Government raises fund by collecting taxes from households, then makes lump-sum transfers, investments and spends on operational activities. Its budget constraint is given by

$$
\begin{aligned}
&\mathrm{B_t + \tau_w W_t H_t + \tau_k \left(r_{k,t} u_t - \delta_p - \delta(u_t)\right) K_{p,t-1} + \tau_k Profits_t + \tau_c C_{p,t}} \\
&= \mathrm{R_{b,t-1} B_{t-1} + G_t + C_{g,t} + I_{g,t}}
\end{aligned}
$$

where the simple fiscal rule is applied as $B_t = \zeta_b Y_t$ with ζ_b a calibrated constant expressing the long-run ratio of government domestic bond over output.

Adopting a basic law of motion for public capital, one reads

$$\mathrm{K_{g,t} = (1 - \delta_g) K_{g,t-1} + I_{g,t}}$$

The model is closed with the national income identity equation as

$$\mathrm{Y_t = C_{p,t} + C_{g,t} + I_{p,t} + I_{g,t}}$$

where

$$
\begin{aligned}
\mathrm{C_{g,t}} &= \mathrm{z_{cg,t} \zeta_{cg} Y_t} \\
\mathrm{I_{g,t}} &= \mathrm{z_{ig,t} \zeta_{ig} Y_t}
\end{aligned}
$$

with ζ_{cg} and ζ_{ig} are the ratios of public consumption and public investment over GDP, respectively. I let C_g and I_g expose to AR(1) and ARMA(1,1) process, respectively, for capturing variations in government expenditure schedule.

$$
\begin{aligned}
\log z_{cg,t} &= \rho_{cg} \log z_{cg,t-1} + \sigma_{cg} \epsilon_{cg,t-1} \\
\log z_{ig,t} &= \rho_{ig} \log z_{ig,t-1} + \sigma_{ig} \epsilon_{ig,t} + \mu_{ig} \sigma_{ig} \epsilon_{ig,t-1}
\end{aligned}
$$

where $\rho_{cg,\,ig}$ and $\sigma_{cg,\,ig}$ are persistent coefficient and standard deviations of z_{cg} and z_{ig} shocks, respectively; and $\epsilon_{cg,\,ig} \sim N(0,1)$. The MA(1) coefficient μ_{ig} is followed SW (2007) to void serial correlation in cyclical public investment due to implementation lags.

5.3.2 Linearisation

From above equilibrium conditions, the modelled economy has been determined by the set of eighteen equations which comprise of four exogenous stochastic process ($z_{a,t}$, $z_{p,t}$, $z_{cg,t}$, $z_{ig,t}$) and fourteen equations defining corresponding endogenous variables

($Y_t, C_{p,t}, C_{g,t}, I_{p,t}, I_{g,t}, K_{p,t}, K_{g,t}, W_t, H_t, u_t, r_{k,t}, R_b, G_t, q_t$).

It is convenient to write all variables in terms of deviation from its steady-state value. The Uhlig's (1997) rule is to define $\tilde{x}_t = \log X_t - \log\overline{X_t}$, so that $e^{\tilde{x}_t} = e^{\log X_t - \log\overline{X_t}}$. Providing that $\tilde{x}_t$ is small, therefore $X_t \approx \overline{X_t}(1+\tilde{x}_t)$. One can interpret $\tilde{x}_t$ as the percentage deviation from its steady-state value, where the latter is denoted by the over-bar capital letter.

Consequently, I yield $\bar{I}_g = \zeta_{ig}\bar{Y}$, $\bar{C}_g = \zeta_{cg}\bar{Y}$ and $\bar{Y}(1-\zeta_{ig}-\zeta_{cg}) = \bar{C}_p + \bar{I}_p$. The law of motion for public capital gives $\frac{\bar{K}_g}{\bar{Y}} = \frac{\zeta_{ig}}{\delta_g}$, whereas wage $\overline{W} = \alpha_3\frac{\bar{Y}}{\bar{H}}$ and capital rental rate $\bar{r}_k = \alpha_1\frac{\bar{Y}}{\bar{K}_p}$. By construction, $\delta(\bar{u}) = \delta(1) = 0$, $\delta'(\bar{u}) = \delta_1$ and $S(\frac{I_p}{I_p} \equiv 1) = 0$ so that I collect the relation of $\delta_1 = (1-\tau_k)\bar{r}_k$, which then implies that the after-tax capital rental rate in the equilibrium equates the accelerating depreciation rate. Since Tobin's Q equation gives $\frac{1}{\beta} = (1-\delta_p) + (1-\tau_k)\bar{r}_k$, I obtain $\delta_1 = \frac{1}{\beta} + \delta_p - 1 \geq \delta_p$ since β is at most of one by definition.

Cobb-Douglas production function provides additional relations

$$\bar{Y} = \bar{K}_p^{\alpha_1}\bar{K}_g^{\alpha_2}\bar{H}^{\alpha_3} = [\bar{H}^{\alpha_3}(\frac{\bar{K}_p}{\bar{Y}})^{\alpha_1}(\frac{\bar{K}_g}{\bar{Y}})^{\alpha_2}]^{\frac{1}{1-(\alpha_1+\alpha_2)}} = [\bar{H}^{\alpha_3}(\frac{\alpha_1}{\bar{r}_k})^{\alpha_1}(\frac{\zeta_{ig}}{\delta_g})^{\alpha_2}]^{\frac{1}{1-(\alpha_1+\alpha_2)}}$$

From the equation $\frac{\chi\bar{H}^{\sigma_L}}{(1-\tau_w)\overline{W}} = \frac{1}{1+\tau_c}[(\bar{C}_p - h\bar{C}_p + \pi\bar{C}_g)^{-1} - \beta h(\bar{C}_p - h\bar{C}_p + \pi\bar{C}_g)^{-1}]$

with $\bar{I}_p = \delta_p\bar{K}_p$ so that $\frac{\overline{C_p}}{\bar{Y}} = (1-\zeta_{ig}-\zeta_{cg}) - \delta_p\frac{\bar{K}_p}{\bar{Y}} = (1-\zeta_{ig}-\zeta_{cg}) - \delta_p\frac{\alpha_1}{\bar{r}_k}$.

Substituting out, one may write the relationship between $\bar{H}$ and other deep parameters as follows

$$\overline{H} = \left[\frac{\alpha_3(1-\tau_w)(1-\beta h)}{\chi(1+\tau_c)[(1-\zeta_{ig}-\zeta_{cg}-\delta_p\frac{\alpha_1}{\overline{r}_k})(1-h)+\pi\zeta_{cg})]}\right]^{\frac{1}{1+\sigma_L}}$$

In the RBC literature, it is more convenient if hours-work $\overline{H}$ is calibrated to a constant $\overline{H}^*$ such that $0 < \overline{H}^* < 1$, then χ is correspondingly calibrated or estimated. As the result of this, $\overline{Y}$ and others steady-state values are derived easily.

I close this section with the set of eighteen log-linearized equations that describes the modelled economy (noting that I use the equal sign "=" instead of "≈" for convenience). The four stochastic processes have the basic form $\log z_t = \rho_z \log z_{t-1} + \epsilon_t$, rewriting $\log \overline{z}e^{\tilde{z}_t} = \rho \log \overline{z}e^{\tilde{z}_{t-1}} + \epsilon_t$. In equilibrium $\overline{z}$ equals to 1, it leads to

$$\tilde{z}_t = \rho\,\tilde{z}_{t-1} + \epsilon_t$$

The national income identity is

$$\overline{Y}\,\tilde{y}_t = \overline{C}_P\,\tilde{c}_{p,t} + \overline{C}_G\,\tilde{c}_{g,t} + \overline{I}_P\,\tilde{i}_{P,t} + \overline{I}_G\,\tilde{i}_{g,t}$$

Public consumption and investments are

$$\begin{aligned}\overline{C}_g\,\tilde{c}_{g,t} &= \zeta_{cg}\overline{Y}(\tilde{z}_{cg,t} + \tilde{y}_t)\\ \overline{I}_g\,\tilde{i}_{g,t} &= \zeta_{ig}\overline{Y}(\tilde{z}_{cg,t} + \tilde{y}_t)\end{aligned}$$

The government budget constraint is

$$\begin{aligned}\overline{G}\,\tilde{c}_t + \overline{C}_g\,\tilde{c}_{g,t} + \overline{I}_g\,\tilde{i}_{g,t} = \quad & \tau_w\overline{WH}(\widetilde{w_t} + \tilde{h}_t) + \tau_k\overline{K}_{P,t-1}[\overline{r}_k(\tilde{u}_t + \tilde{r}_{k,t} + \tilde{k}_{p,t-1}) - \delta_P\,\tilde{k}_{p,t-1} - \delta_1\,\tilde{u}_t]\\ & + \tau_k\overline{Profits}\,\widetilde{profits}_t + \tau_c\overline{C}_p\,\tilde{c}_{p,t}\end{aligned}$$

The law of motion of public capital is

$$\overline{K}_g\,\tilde{k}_{g,t} = \overline{K}_g(1-\delta_g)\,\tilde{k}_{g,t-1} + \overline{I}_g\,\tilde{i}_{g,t}$$

Regarding firms

$$
\begin{aligned}
\tilde{w}_t &= \alpha_3(\tilde{y}_t - \tilde{h}_t) \\
\tilde{r}_{k,t} &= \alpha_1(\tilde{y}_t - \tilde{k}_{p,t-1}) \\
\widetilde{profits} &= \alpha_2(\tilde{y} - \tilde{k}_{g,t-1}) \\
\tilde{y}_t &= \tilde{z}_{a,t} + \alpha_1\,\tilde{u}_t + \alpha_1\,\tilde{k}_{p,t} + \alpha_2\,\tilde{k}_{g,t-1} + \alpha_3\,\tilde{h}_t
\end{aligned}
$$

Regarding households

$$\tilde{r}_{b,t} = \tilde{\lambda}_t - \tilde{\lambda}_{t+1}$$

$$(1-\tau_k)\,\tilde{r}_{k,t} = \tilde{q}_t + \frac{\delta_2}{\delta_1}\tilde{u}_t \Rightarrow \tilde{r}_{k,t} = \frac{\tilde{q}_t}{1-\tau_k} + \frac{\delta_2}{\delta_1(1-\tau_k)}\tilde{u}_t$$

$$\frac{(1+\tau_c)z_{p,t}\chi H_t^{\sigma_L}}{(1-\tau_w)W_t} = z_{p,t}aux_t - \beta h E_t[z_{p,t+1}aux_{t+1}]$$

where $aux_t = \left(C_{p,t} - hC_{p,t-1} + \pi C_g\right)^{-1}$ and $\widetilde{aux_t} = \frac{(h\tilde{c}_{p,t-1} - \tilde{c}_{p,t})\overline{C}_P - \pi\overline{C}_G\tilde{c}_g}{\overline{aux}}$, then applying the first-order Taylor expansion one gets the linear relationship:

$$\frac{\chi(1+\tau_c)\overline{H}^{\sigma_L}}{(1-\tau_w)\overline{W}_t}(\tilde{z}_{p,t} + \sigma_L\,\tilde{h}_t - \tilde{w}_t) = \overline{aux}(\tilde{z}_{p,t} + \widetilde{aux_t}) + \beta h\overline{aux}\,E_t[\tilde{z}_{p,t+1} + \widetilde{aux}_{t+1}]$$

Similarly, the remaining linear equations related to private investment and Tobin's Q, respectively, are

$$\tilde{i}_{p,t} - \tilde{i}_{p,t-1} \equiv \Delta i_{p,t} = \frac{1}{\kappa}\left(\frac{1+\tilde{q}_t}{1+\beta}\right)$$

$$\tilde{q}_t = \tilde{r}_{k,t} - \tilde{u}_t$$

From the last two equations, it can be seen explicitly that the lower investment costs κ and intertemporal discount factor β the higher the changes in private investment given the same level of Tobin's Q; and, the Tobin's Q only positively changes when $\tilde{r}_{k,t} > \tilde{u}_t$, implying that to preserve "the benefit from investment per unit of capital" the marginal rental rate from over-utilisation capital should be at least equating the changes in rate of capital utilisation.

5.4 Bayesian estimation

I adopt the Bayesian Markov Chain Monte Carlo (MCMC) estimation strategy – which is now implemented in Dynare (Adjemian *et al.*, 2011) and similar packages – used in SW (2003, 2007) and the most recent DSGE literature. Technical analysis of MCMC algorithms could be further found in An and Schorfheide (2007) or Herbst and Schorfheide (2015). At the core of Bayesian estimator, the Bayes rule for the conditional distribution of parameter set $\theta \in \Theta$ given observational data Υ is written as

$$\pi(\theta|\Upsilon) = \frac{f(\Upsilon|\theta)\pi(\theta)}{f(\Upsilon)} \tag{6}$$

where $\pi(\theta|\Upsilon)$ is the 'so-called' posterior probability distribution function of the parameter θ (or posterior distribution shortly) conditional on observational data Υ. The prior distribution $\pi(\theta)$ is the unconditional probability distribution of θ whilst the likelihood function $f(\Upsilon|\theta)$ is defined as in classical econometric methods. The last component $f(\Upsilon)$, i.e. marginal likelihood, is a constant as such $f(\Upsilon) = \int f(\Upsilon|\theta)\pi(\theta)d\theta$.

By construction, marginal likelihood is independent of estimating parameter set θ, it is convenient to rewrite equation (6) in the following form

$$\pi(\theta|Y) \propto f(\Upsilon|\theta)\pi(\theta) \tag{7}$$

The posterior $\pi(\theta|Y)$ in equation (7) retains all distributional characteristics as in the original form (6) and embraces information before knowing the data, i.e. $\pi(\theta)$, and information contained in the observables, $f(\Upsilon|\theta)$. In Bayesian view, the prior will be updated by the likelihood function when receiving new data. Thus, equation (7) explicitly requires us specifying prior distributional knowledge about modelled parameters and evaluating the likelihood function for every data point. In practice, one works with log-likelihood function so that the posterior is the sum of two right-hand-side components.

5.4.1 Model settings

Theoretically, the prior distribution, $\pi(\cdot)$, could be any known distribution basing on one's informative beliefs in advance about parameters to be estimated. Practically, I choose prior distribution in such the way that its support could afford the parameter boundary. Taking technological parameter of public capital, α_2, as an example, I assume the positive elasticity α_2 has the lower bound of zero, hence, the gamma distribution having the support of $[0, +\infty)$ could be a suitable choice.

I am going to estimate the first group of seven structural parameters $\boldsymbol{\Theta_1} = \{h, \eta, \chi, \pi, \psi, \alpha_1, \alpha_3\}$. η is the Frisch elasticity as $\sigma_L = \frac{1}{\eta}$. Literature suggests η has a support of $[0, +\infty)$, so I specify $\eta \sim \mathrm{Gamma}(0.5, 0.2)$, where the first and second parameter in the parentheses represent mean and standard deviation of the Gamma distribution, respectively. Habit formation and investment adjustment costs have been set similar to SW (2003) as $h \sim \mathrm{Beta}(0.7, 0.1)$ and $\psi \sim \mathrm{Beta}(0.8, 0.05)$, respectively. Stressing that the prior of ψ is fairly tight stemming from my belief that investment costs in Vietnam could be comparatively higher than one in a typical advanced economy, say, the United States.

Given earlier discussions, π is assumed to be non-negative but not greater than one, giving rise to the prior of $\mathrm{Beta}(0.5,\ 0.2)$. This implies government spending would interfere household utility whenever $0 \leq \pi < 1$. For technological parameters of private inputs, α_1 and α_2, I consider two cases: (i) calibrating $\alpha_2 = 0.1$ and constant returns to scale with respect to private inputs as in Baxter and King (1993) and Christiano and Eichenbaum (1992), which then implies $\alpha_1 + \alpha_3 = 1$; (ii) assuming $\alpha_1 + \alpha_2 + \alpha_3 = 1$, therefore, α_2 is implicitly derived by estimating α_1 and α_2. The parameter denoting labour dis-utility, χ, comes with a prior of $\mathrm{Gamma}(8, 3)$, whose has the mean such that $\overline{H} \approx 0.3$.

The second set consists of parameters that characterises the propagation mechanisms of model endogenous, namely $\boldsymbol{\Theta_2} = \{\rho_a, \sigma_a, \rho_p, \sigma_p, \rho_{ig}, \sigma_{ig}, \mu_{ig}, \rho_{cg}, \sigma_{cg}\}$. According to the DSGE literature, their priors are pinned down straightforwardly such that AR(1) coefficients are of $\mathrm{Beta}(0.6, 0.15)$

except ρ_{cg} is calibrated to 0.7 because of its failing to Iskrev's (2010) test; the rest of standard deviations $\sigma_{(\cdot)}$ follow the same $\mathrm{Inverse} - \mathrm{Gamma}(0.01, 0.1)$ distribution.

Finally, I calibrate other 'deep' parameters that are ill-identified due to limited supporting information by aggregate data. Depreciation rates, government spending-related ratios, and tax codes are justified from publicly exogenous data and the RBC literature. Notably, I set annual discount rate by 6.5% (Pham *et al.*, 2019a) implying $\beta = 0.939$.

It is obvious to us that public capital is normally expected to be depreciated at the lower pace than the private counterpart, and government spending schedule should fluctuate around a fixed ratio to GDP. In general, the choice of depreciation rate in any RBC analysis varies from 3% to 12% depending upon the specific sector and modelled economy. Cooley and Prescott (1995) argued that steady-growth economy should be depreciated by 5% per annum but the higher rate, about 10%, has been recommended for a zero-growth economy. Consequently, I assume private capital wearing out at 8% per annum, meanwhile, the depreciation rate of public capital takes 4%.

Vietnam's government consumption was around 6% in the 2000s, yet public investment has been much higher. The VN statistical yearbooks have reported public investment accounting for 35% – 40% of gross investment over past decades (see also Figure 1). As gross investment over GDP in the thirty-years period had dramatically increased from the low of 14% - 17% in years before 1992 to the high of 33% - 39% in the mid of 2010s, I calibrate public consumption and public investment ratios as of 6% and 10%, respectively. Noting that my choice of ζ_{ig} ratio is slightly higher than that of IMF dataset because the latter only refers to capital investment. Also, the level of government debt, ζ_b, is set to the long-run value of 0.35.

θ	α_2	δ_P	δ_G	τ_c	τ_k	τ_w	ζ_{CG}	ζ_{IG}	ζ_b
0.065	0.10	0.08	0.04	0.10	0.25	0.25	0.06	0.10	0.35

Table 2: Calibrated parameters

Tax codes are another difficult thing for the young economy like Vietnam. Before the Asian financial crisis, Vietnam government had not imposed VAT tax on goods and services, since then an identical consumption tax (τ_c) of 10% is applied for most goods and services. Labour income tax (τ_w) bracket has mainly been social contributions for a long time. It is worth mentioning that there are two personal income tax bases, namely basic wage and actual wage, in Vietnam. The basic wage is only used as tax base for social and pension contributions while the progressive income tax is levied on actual income net of deductibles. Since family deductibles is relatively high compared to Vietnam average labour wage, the "de-facto" tax revenue from high income person is actually as low as 5% of the total fiscal revenue in most past years. Corporate income tax decreased considerably from 28% in 2000s to 22% by 2014 but tax on capital gain such as leasing is straight as of 20%. As a matter of fact, I calibrate the flat corporate and labour income taxes as of 25%. Table 3 and Figure 4 thus report just-discussed priors and their corresponding posteriors, and Table 2 presents all calibrated parameters.

The last step is to point out measurement equations in order to evaluate the likelihood function described in equation (6). I employ six annual data series, namely output, private consumption, investment, government consumption, public investment, and hours-work. They all are normalized to per capita values and detrended with the one-sided Hodrick-Prescott filter (Stock and Watson,1999) because the model belongs to the class of stochastic growth.

In DSGE estimation one must have as many exogenous shocks in the model as the number of observation series, otherwise the estimation is in trouble with the singularity matter. My model has four shocks, I cope with singularity problem by incorporating two measurement errors into two investment linking equations. To this end, six measurement equations are defined as follows

$$\tilde{\mathbf{y}}_{\mathbf{obs}} = \begin{bmatrix} y_t - \bar{y} \\ c_{p,t} - \bar{c}_p \\ c_{g,t} - \bar{c}_g \\ i_t - \bar{i} \\ i_{g,t} - \bar{i}_g \\ h_t - \bar{h} \end{bmatrix} + \begin{bmatrix} 0 \\ 0 \\ 0 \\ \epsilon_{ime} \\ \epsilon_{igme} \\ 0 \end{bmatrix} \tag{8}$$

where $\tilde{\mathbf{y}}_{\mathbf{obs}} \equiv \{\tilde{y}_t, \tilde{c}_{p,t}, \tilde{c}_{g,t}, \tilde{i}_t, \tilde{i}_{g,t}, \tilde{h}\}$ is a vector of cyclical components of the respective aggregates; the small letter denotes variable in logarithm, and the over bar denotes steady-state value. Thus, the first block of the above equations expresses variable deviations from their equilibrium states in the model.

		Prior			Case I: $\alpha_1 + \alpha_3 = 1$			Case II: $\alpha_1 + \alpha_2 + \alpha_3 = 1$		
Param	**Description**	**Mean**	**SD**	**Dist.**	**Mean**	**5%**	**95%**	**Mean**	**5%**	**95%**
h	Habit formation	0.70	0.10	Beta	0.760	0.693	0.826	0.753	0.684	0.821
η	Frisch elasticity	0.50	0.20	Gamma	0.931	0.556	1.299	0.909	0.535	1.268
χ	Labour dis-utility	8.00	3.00	Gamma	7.884	3.189	12.38	7.962	3.416	12.682
π	Public spending elasticity	1.00	0.50	Beta	0.694	0.179	1.219	0.745	0.180	1.276
ψ	Investment costs	0.80	0.05	Beta	0.769	0.708	0.832	0.768	0.705	0.832
α_1	Private capital elasticity	0.35	0.05	Beta	0.265	0.215	0.315	0.260	0.204	0.314
α_3	Labour elasticity	0.50	0.20	Beta	-	-	-	0.606	0.370	0.848
ρ_a	TFP persistence	0.60	0.15	Beta	0.594	0.459	0.737	0.607	0.462	0.751
ρ_p	Preference persistence	0.60	0.15	Beta	0.373	0.214	0.528	0.371	0.209	0.518
ρ_{ig}	Public investment	0.60	0.15	Beta	0.724	0.607	0.844	0.719	0.601	0.839
μ_{ig}	MA(1) persistence	0.60	0.15	Beta	0.397	0.242	0.542	0.399	0.253	0.554
σ_a	Productivity shock	0.01	0.10	InvGam	0.022	0.017	0.027	0.020	0.014	0.026
σ_p	Preference shock	0.01	0.10	InvGam	0.096	0.069	0.121	0.095	0.070	0.119
σ_{ig}	Public investment shock	0.01	0.10	InvGam	0.065	0.051	0.077	0.065	0.051	0.078
σ_{cg}	Government spending shock	0.01	0.10	InvGam	0.021	0.017	0.025	0.021	0.017	0.025
σ_{ime}	Measurement error of i_t	0.05	0.20	InvGam	0.080	0.064	0.095	0.080	0.065	0.096
σ_{igme}	Measurement error of i_g	0.05	0.20	InvGam	0.103	0.081	0.123	0.103	0.083	0.125

Table 2: Priors and Posteriors. Sample for estimation: 1981 – 2015.

5.4.2 Results

The structural parameters are estimated over the sample period 1981 – 2015, which is chosen to be consistent with Pham *et al.* (2019a). The last three columns of Table 2 report the posteriors of case (ii) constant returns to scale with respect to all inputs, whereas the next three columns on the left are those of the case (i) fixing $\alpha_2 = 0.1$ and letting $\alpha_1 + \alpha_3 = 1$ (Baxter and King, 1993). In both cases, I simulate the posterior distributions through 800,000 draws from the MHMC[3] algorithm.

It is my very first observation that the model has been informatively supported by data as most of the posterior distributions are visually distinguishable from the priors (see Figure 4), indicating that the

[3] Metropolis – Hastings Monte Carlo algorithm.

estimates are generally robust except for ρ_{cg} being failed the identification test. Basically, I obtain the highly compatible values between two returns-to-scale settings even though there are some minor differences among household-related structural parameters.

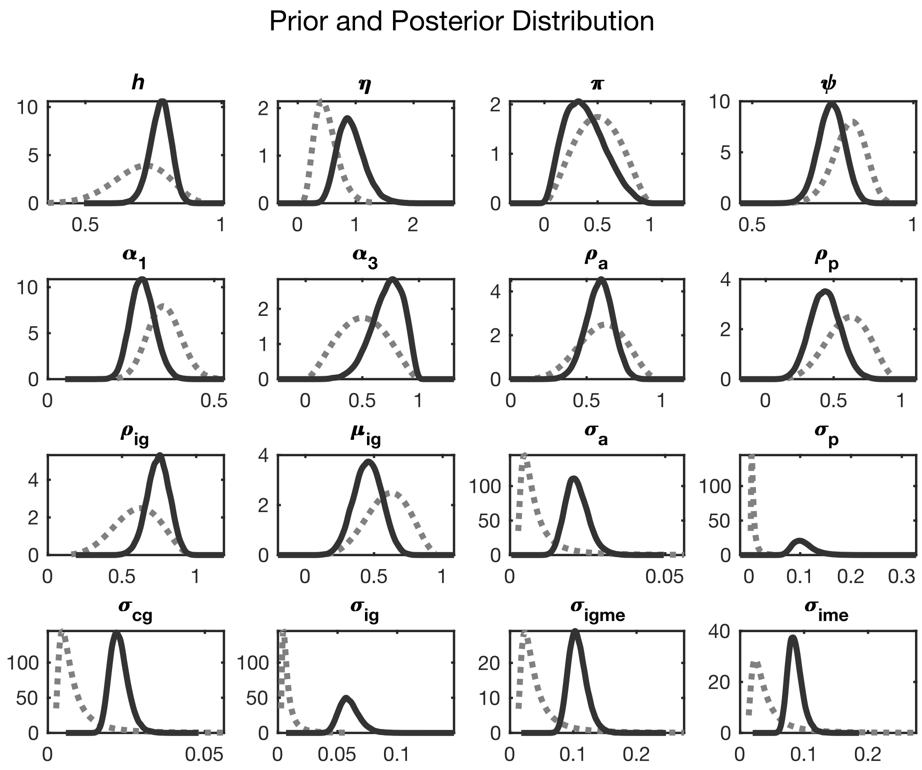


Figure 4: Prior (dotted line) and Posterior (solid line) Distributions

In case (i), $\alpha_1 = 0.265$ implies $\alpha_3 = 0.735$, meanwhile α_1 and α_3 are of 0.260 and 0.606, respectively, in case (ii). The latter case also implies that $\alpha_2 = 0.134$ which is higher the average value of 0.106 noted in Bom and Ligthart (2014) but within the range of [0.13, 0.20] in Arslanalp *et al.* (2010) for OECD and non-OECD (middle- and lower-income) countries, respectively. Since public capital generates extraordinary income, the higher value of α_2 suggests the stronger effects of fiscal policy shocks on output. Accordingly, in much of what follows, I concentrate on the outcomes of case (ii) as the output elasticity of public capital has been implicitly estimated from the model.

The government spending elasticity, π, is less than unity showing that government consumption induces crowding-out effects on household utility, but the magnitude has not been large as $\pi > 0.7$.

The household consumption persistence appears to be strong, $h = 0.75$, over the sample period,[4] which is, however, almost double of one – based on growth data – found in Pham *et al.* (2019a).

Regarding the (aggregate) Frisch coefficient, $\sigma_L = 1/\eta$, the estimate shows that η is just above one, as of 1.1. Since the lower the Frisch coefficient results in the less hours-work variations in response to temporary changes in current or future income, this reflects the fact that the VN hours-work should not vary greatly over economic cycles. The investment adjustment costs $\kappa = \frac{\psi}{1-\psi} = 3.31$, implying that the changing rate of private investment growth, $1/\kappa$, is of 0.3 higher than the estimated value of Leeper *et al.* (2010) for the U.S. economy by 0.1.

In reference to persistent shocks, it is shown that Vietnam's economy was fairly instability given the moderate values of AR(1) coefficients of transitory TFP ($\rho_a = 0.61$), but the low of preference shifter ($\rho_p = 0.37$) over the period 1981 – 2015. The standard deviation of household preference, σ_p, is notably large, about 10%, indicating that the VN consumption preference varied significantly from time to time. The two processes related to government expenditures have the similar AR(1) coefficients in the range of $[0.65, 0.70]$ although their standard deviations are different substantially. Public investment deems to be the higher fluctuating component with $\sigma_{ig} = 6.5\%$, more than triple of the respective government consumption, $\sigma_{cg} = 2.1\%$.

To understand how much RBC theory can explain the target economy I compare the variances, contemporaneous correlations among output and main aggregates such as consumption, investments, and hours-work among actual data, the estimated model, and the vanilla RBC model due to Hansen (1985). I simulate 1000 economy realizations, then adjust simulated series with HP filter. The moments are computed corresponding to the sub-sample timespan.

[4] Note that, the higher intensity of habit formation implies that a unit of current consumption will raise the marginal utility of consumption in the next period whilst decreases it in the present period.

Table 3 shows that the Hansen's RBC model reproduces around 95% of actual output variations, $\sigma_y^{\text{vanilla}} = 1.77\%$ compared to $\sigma_y^{\text{actual}} = 1.88\%$, but it fails to explain the variations in hours-work and the final consumption. Importantly, aggregate variables in the plain model are strongly correlated with the output as such contemporaneous correlations are all above 0.89. The actual data report the much lower contemporaneous correlation with output, implying the vanilla RBC cannot capture the dynamic of Vietnam business cycles. It should be noted that there is no shock other than a shock to economy-wide technology in the Hansen model, my estimated model, nonetheless, takes the advantages of DSGE modelling with three more structural shocks.

The estimated model produces a slightly lower standard deviation of actual output, about 80%, but it adequately captures several excess volatilities of main aggregates to output. In this regard, the estimated model successfully reproduces two relative ratios σ_{cp}/σ_y and σ_h/σ_y, although it somewhat understates the private investment activities and the government consumption. The model also deems failing to capture contemporaneous correlation between investment and output because of the implied crowding-out effect on private investment in the household budget constraint. However, it clearly outperforms the basic RBC setting in reproducing the VN aggregate moments.

Measure	Standard Deviation, %				Contemporaneous correlation			1st-order Autocorrelation		
	Actual	Estimated	Vanilla		Actual	Estimated	Vanilla	Actual	Estimated	Vanilla
σ_y	1.88	1.50	1.77	y	1.000	1.000	1.000	0.737	0.505	0.442
σ_{cp}/σ_y	1.18	1.22	0.54	c_p	0.629	0.824	0.897	0.596	0.529	0.619
σ_{cg}/σ_y	2.23	1.66	-	c_g	0.497	0.624	-	0.727	0.383	-
σ_i/σ_y	5.87	2.55	4.23	I	0.574	0.354	0.942	0.508	0.533	0.339
σ_{ig}/σ_y	6.71	5.16	-	i_g	0.272	0.540	-	0.237	0.518	-
σ_h/σ_y	1.29	1.30	0.24	h	0.380	0.106	0.906	0.658	0.272	0.334

Note: Vanilla is standard RBC as in Hansen (1985). Actual moments are estimated using HP filter with $\lambda = 100$. Estimated is posterior mean stochastic simulation. All variables are cyclical components.

Table 3: Actual versus simulated moments from model

5.4.3 Impulse – response analysis

It is also interesting to explore the time paths of output, investments and other aggregate variables due to the positively orthogonalized shocks at any time t = 0. That means one would want to observe the responses of each modelled aggregate variables given one-time structural shocks, i.e. total factor productivity (ϵ_{a}), consumption preference (ϵ_{p}), and public investment (ϵ_{ig}).[5]

Chart I-A of Figure 5 depicts the different median responses of the output to ϵ_{a}, ϵ_{ig}, and ϵ_{p}.[6] It is shown that the effects of ϵ_{ig} and ϵ_{p} almost die out after four periods (or four years), whereas the impacts of ϵ_{a} has a hump-shaped curve with the peak at time t = 3 and dying out significantly after ten years. Positive shocks to public investment in Vietnam deem to cause a critical long-run consequence on output because of the higher accumulated productive capital stock gives rise to the permanent deviation of output from its steady state (see also the dashed-line in chart I-B).

An increase in the public investment would create new labour demand, yet its total effect is comparatively smaller than the definite shift in consumption preference. It is evident that in the closed and non-monetary economy, the government could finance additional investments by imposing new taxes or issuing more debts, entailing the proportionate curtailment of the household consumption. But owing to the crowding-in impacts on output, this side effect is ruled out in the middle or long run as shown in the chart II-A and II-B of Figure 5.

As can be seen from the plot IV-B, the negative impacts of TFP shock on employment would be eliminated if there was a hike in the public investments simultaneously. That means the government could sustain economic growth by channelling funds into the “core infrastructure” projects with the target of improving the economy-wide total factor productivity. Financially subsidizing R&D activities in both of the private and the public sectors could also be a way to achieve the long-term growth targets. In the middle term, either improving the public transport infrastructure such as bus and railway systems, etc., or building the new nationwide highway network would have two effects: an additional public investment boosts demand and pushes up output; and firms would then earn extraordinary profits by utilizing that productive public capital stock. The latter effect implies that household consumption and

[5] The responses to government consumption (ϵ_{cg}) are practically omitted due to its negligible impacts.

[6] Each IRF is the median of its Bayesian responses to the respective shock computed over MHMC draws.

tax revenues would increase, hence, the risk of high indebtedness could be minimized. Besides, in the plot III-A and III-B, the dotted blue lines suggest that shocks to public investment, ϵ_{ig}, crowd in investment from private sector, supporting empirical findings in Pham *et al.* (2019b).

In sum, the IRFs are consistent with the IMF's evidence (Abiad *et al.*, 2014) that empirically stressed on the short- and long-term consequences of public investment on output growth. In the context of developing economies, these authors predict "the contemporaneous effect of a one percentage point of GDP increase in public investment is a one-fourth percent increase in output". The estimate suggests the higher impact of public investment on the Vietnam's economy, i.e. around 0.91% of contemporaneous response of private income and 3% accumulative output gain in five years to 1% real GDP of public investment impulse – which is equivalent to a shock of 10%, $\sigma_{ig} = 10\%$, public investment. This figure, however, does not imply the higher efficiency of public investment in Vietnam.

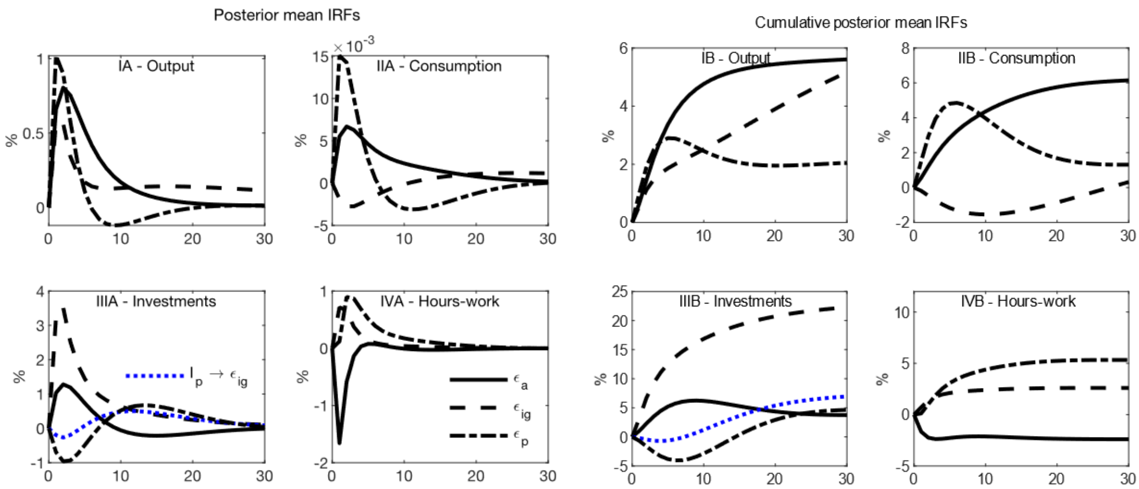


Figure 5: Bayesian Impulse – Responses (mean values)

5.4.4 Variance decomposition

One of the advantages of DSGE estimation techniques is the capability of analysing the propagation mechanisms through Kalman disturbance smoothing. The contribution of each exogenous process to output variance could be computed and visualised as depicted in Figure 7.[7] It is shown that TFP shocks historically had significantly positive impacts on the income growth in the 1990s, but the contributions were reversed in most years in the twenty-first century. Public investment played a crucial role in the early-to-mid 1990s. By the end of Asian financial crisis, Vietnam economy was subject to structural changes, the contribution of TFP shocks flipped to the negative stand from 2002, yet private consumption and public investment shocks helped stabilise the output fluctuations until 2007 – the year before another crisis.

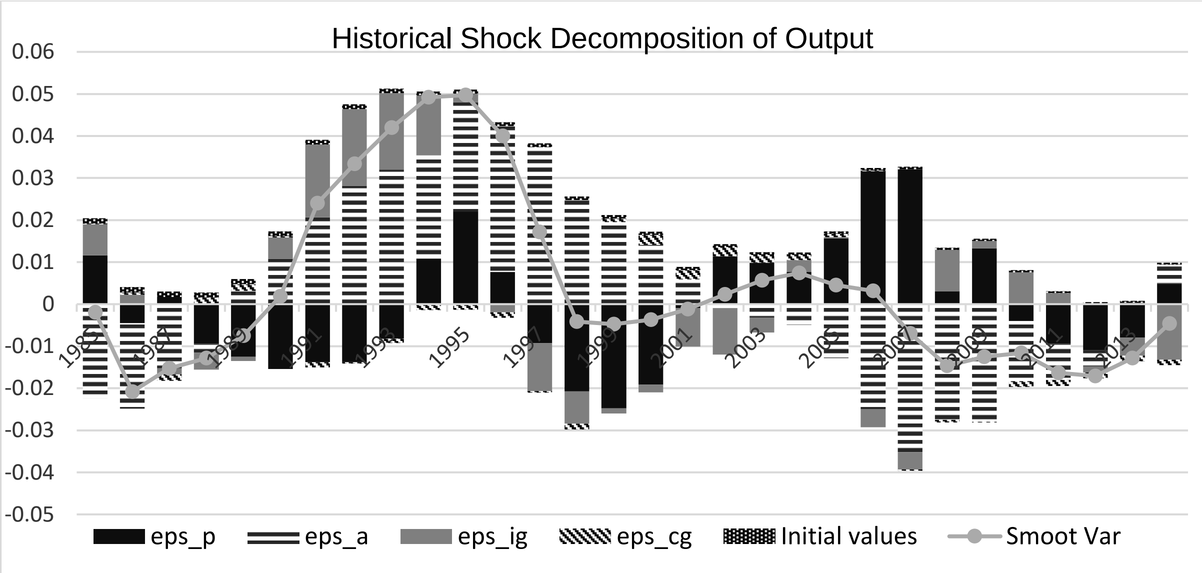


Figure 7: Historical Shock Decomposition of Output. Abbreviation (eps ≡ ϵ)

If one divides the Doi Moi period into three episodes of about ten-year each, it seems to be that there were three economic cycles during the course. The first cycle had started from the beginning of 1980s to the mid-1990s. By the end 1980s, Vietnam economy suffered from the collapse of the Communist Bloc and the occurrence of hyperinflation, hence, private consumption was almost below its trend in the first cycle. One of the most important changes in the 1992 constitutional law was the official recognition of Vietnam private sector, winding up the first wave of public and private investments.

[7] The historical shock decomposition plots of private consumption, investment, and hours-work are in the appendix.

The second cycle witnessed the Asian crisis in 1997, and the Vietnam - US Bilateral Trade Agreement (BIT) postponed in 2000 which had nonetheless stopped the favourable investment motion in several years, 1997 – 2002. Since then, the investment level had strongly deviated to the negative side of its trend until 2011, with the significant contributions of public investment and productivity shocks. In fact, over the period 2001 – 2007, the VN household experienced over-consumption of goods and services, resulting in unbalancing between consumption and investment.

In the period 2005 – 2015, shocks to public expenditures seems not to have notable impacts on output fluctuations. Instead, TFP shocks pulled down output substantially in the first seven years, 2005 – 2011, implying that there were serious concerns about intrinsic weaknesses of the VN economy, which has mostly relied on labour-intensive industries, and exporting raw materials and agricultural products. This implies public capital stocks built in these years were likely less productive or inefficient.

It is worth stressing that a positive (negative) shock to household preference causes an increase (decrease) in factor β, which in turn implies a decrease (increase) in the pure temporal discount rate θ, respectively. If a household puts less weight on his future consumption (the lower θ) then he spends more on today. The shock decomposition of private consumption points out that the favourable business atmosphere flourished the Vietnamese household belief in the mid of 1990s and 2000s. The reversed things happened when the economy suffered from the episodes of crises.

Computing the unconditional (conditional) variance decomposition reveals additionally insightful about public investment innovations. Specifically, Table 4 shows that the variability of public investment can explain more than 29% of output variance in the long run (unconditional variance); this contribution is moderately lower than the contributions of TFP and preference shocks, 38.3% and 32.4%, respectively. Also, the variance decomposition conditional on the first year (t = 1) shows public investment shocks contributing to just less than one-fifth of the output and hours-work variance.

Meanwhile, together two shocks, ϵ_a and ϵ_p, greatly describe variations in private consumption, private investment, and hours-work in the long run, although ϵ_p appears to be more important than ϵ_a

in the short run. Figure 8 summarises the contributions of four exogenous shocks to six endogenous variances conditional on information up to date t = 1, 3, 5…, and 30 years. It is easily seen the predominant contributions of ϵ_p and ϵ_a – blue and red stacked bars, respectively – to variances of four out of six variables over the forecasting horizons. Shocks to TFP explain only 25% of income fluctuations in the first period but more than 42% from the third horizons. As a matter of fact, fiscal expansion via capital stock spending could foster Vietnam income growth in both short and long horizon if the implementation is supposed to be productive and efficient.

	Unconditional Variance Decomposition				Conditional Variance Decomposition (t = 1)			
	ϵ_p	ϵ_a	ϵ_{ig}	ϵ_{cg}	ϵ_p	ϵ_a	ϵ_{ig}	ϵ_{cg}
$\tilde{y}$	32.37	38.28	29.15	0.20	56.81	24.55	18.32	0.33
$\tilde{c}_p$	59.79	23.16	16.36	0.70	86.02	12.20	1.47	0.31
$\tilde{c}_g$	14.1	17.29	14.05	54.55	16.03	6.88	5.06	72.03
$\tilde{\iota}_g$	1.49	1.89	96.62	0.01	2.19	0.94	96.86	0.01
$\tilde{\iota}_p$	66.05	25.21	8.72	0.02	70.86	28.49	0.63	0.02
$\tilde{h}$	38.44	43.80	17.48	0.28	1.49	81.94	16.35	0.22

Table 4: Forecast Errors Variance Decomposition

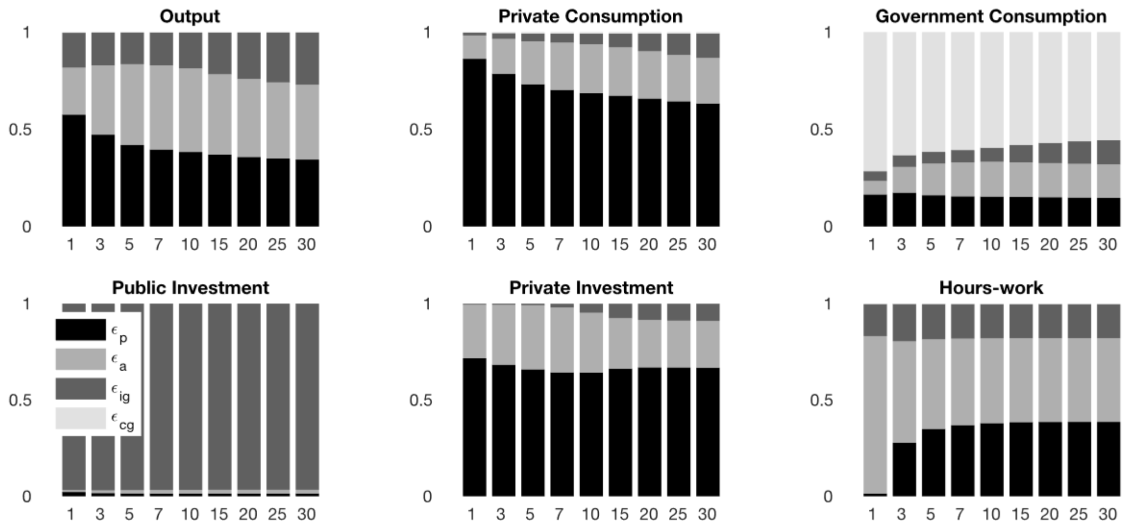


Figure 8: Conditional Variance Decomposition of Output

5.5 Policy Simulation

So far, I have examined the role that public sector activities play in promoting economic growth. In this Section, we further enquire on this role by comparing the economic effects of two expansionary fiscal policies that differ in their composition. I understand that periods of fiscal expansion may involve several mechanisms and thus consider two possible sets of measures, both involving permanent and temporary changes.

The fiscal policy mix 1 (FPM1) involves an increase in public consumption by 2 percentage points (from 7% to 9% of GDP), which is funded by an analogous increase in consumption taxes (from 10% to 12% of GDP). In addition, there is a 5-year shock equal to a 10% increase in public investment. This policy mix may by mildly productive, in which case TFP grows 0.25% for the next 15 years (scenario A), or highly productive, in which case TFP grows by 0.50% instead (scenario B). In turn, the fiscal policy mix 2 (FPM2) involves an increase in public investment by 2 percentage points (from 10% to 12% of GDP), which is funded by an analogous increase in consumption taxes (from 10% to 12% of GDP). In addition, there is a 5-year shock equal to a 10% increase in public consumption. Since this policy mix involves a permanent increase in public investment, its consequences for TFP are also permanent. This investment, however, may be mildly productive, in which case TFP grows 0.25% for ever (scenario C), or highly productive, in which case TFP grows by 0.50% instead (scenario D). Table 5 summarises the two sets of measures.

		FPM1				**FPM2**			
		A		**B**		**C**		**D**	
Permanent shock	Before policy	$\zeta_{cg} = 7\%$, $\zeta_{ig} = 10\%$, $\tau_c = 10\%$				$\zeta_{cg} = 7\%$, $\zeta_{ig} = 10\%$, $\tau_c = 10\%$			
	After policy	$\zeta_{cg} = 9\%$, $\zeta_{ig} = 10\%$, $\tau_c = 12\%$				$\zeta_{cg} = 7\%$, $\zeta_{ig} = 12\%$, $\tau_c = 12\%$			
Temporary shock	Period	[1:5]	[6:20]	[1:5]	[6:20]	[1:5]	[6:∞]	[1:5]	[6:∞]
	ε_a		+0.25%		+0.5%		+0.25%		+0.5%
	ε_{cg}					+10%		+10%	
	ε_{ig}	+10%		+10%					

Table 5: Policy mix scenarios.

Assuming that the government anticipates the information contained in these sets of measures, I examine how the trajectories of the variables in the national income identity converge to their balance-growth-paths. This practice is also known as perfect-foresight (deterministic) simulation since agents are assumed to know precisely how the shocks are loaded into the state-space system.

The deterministic model that can be written in the form:

$$\mathrm{f}(\mathrm{y}_{\mathrm{t+1}}, \mathrm{y}_{\mathrm{t}}, \mathrm{y}_{\mathrm{t-1}, \mathrm{u}_{\mathrm{t}}}) = 0 \tag{8}$$

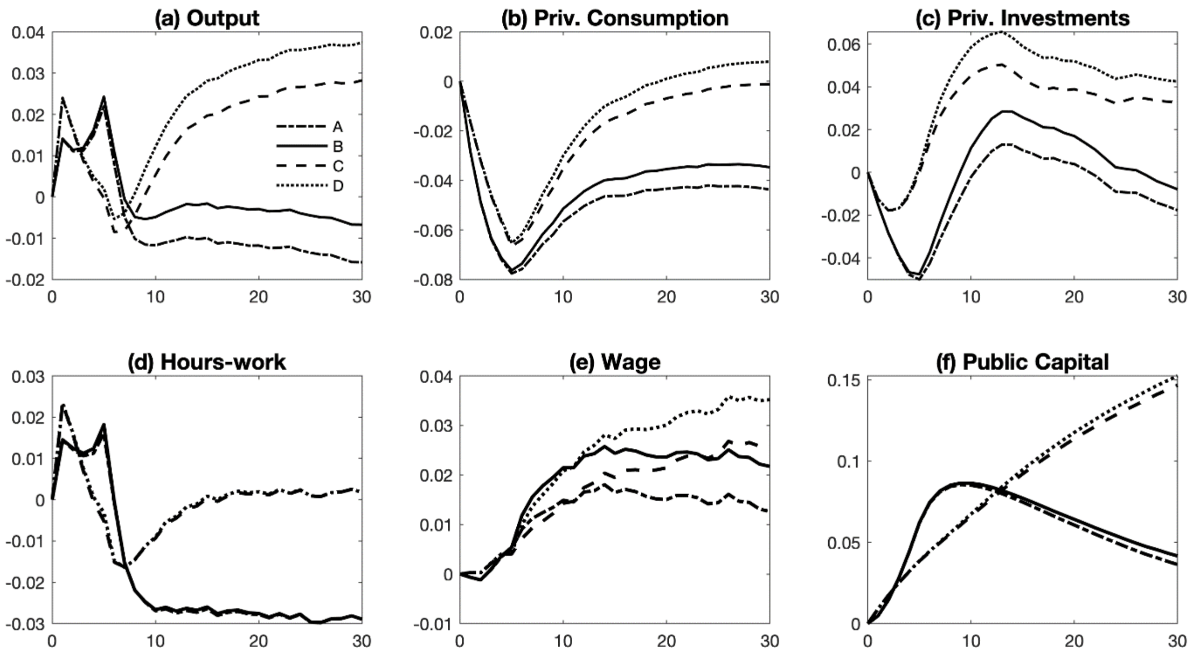
for $\mathrm{t} = 1, \ldots, \mathrm{T}$; where y is a $(\mathrm{n} \times 1)$ vector of endogenous variables with the initial condition y_0 taken as given so that $\mathrm{y}_{\mathrm{T}} \equiv \bar{\mathrm{y}}$ is the steady-state of the variables such that $\mathrm{f}(\bar{\mathrm{y}}, \bar{\mathrm{y}}, \bar{\mathrm{y}}, \bar{\mathrm{u}}) = 0$. Vector u_{t} is a zero-mean process with $(\mathrm{q} \times 1)$ perfectly anticipated innovations. Solving model (8) delivers the set of $(\mathrm{y}_1, \ldots, \mathrm{y}_{\mathrm{T-1}}, \mathrm{y}_{\mathrm{T}})$ values given by the initial y_0 and the terminal $\mathrm{y}_{\mathrm{T+1}}$, that satisfy the stacked system of nonlinear equations (Juillard, 1996):

$$\begin{aligned} \mathrm{f}(\mathrm{y}_2, \mathrm{y}_1, \mathrm{y}_0, \mathrm{u}_1) &= 0 \\ &\vdots \\ \mathrm{f}(\mathrm{y}_{\mathrm{t+1}}, \mathrm{y}_{\mathrm{t}}, \mathrm{y}_{\mathrm{t-1}}, \mathrm{u}_{\mathrm{t}}) &= 0 \\ &\vdots \\ \mathrm{f}(\mathrm{y}_{\mathrm{T+1}}, \mathrm{y}_{\mathrm{T}}, \mathrm{y}_{\mathrm{T-1}}, \mathrm{u}_{\mathrm{T}}) &= 0 \end{aligned}$$

Based on the parameters obtained in the previous Bayesian estimation, I simulate 1000 stochastic TFP paths with zero mean and standard deviation of 2.2% ($\epsilon_a \sim N(0, \sigma_a^2 \equiv 0.022^2)$), and solve for the deterministic solutions using Dynare 4.5.7 package (Adjemian *et. al*, 2011). Figure 9 presents the combination of the median paths of output (panel a), private consumption (b), private investment (c), hours-work (d), wage (e), and public capital (f). The unit of time is meant to be a year. The corresponding individual plots with their error bands can be found in the appendix.

Figure 9 shows that the government spending plans characterised by A and B result in long-run declining paths of output, private consumption, private investment, hours-work, and public capital. Contrariwise, the gaps between the dotted lines (case D) and the dashed lines (case C) in all panels (a –

f) illustrate the benefits of a more efficient and productive set of measures such as the one characterised by FPM2. In the first five years, because of the simultaneous impacts of the temporary government consumption shocks and the permanent change in public investment level, both private consumption and private investment fall. Nonetheless, private investment quickly rebounds after the fifth period, and rises subsequently to reach its peak just seven years afterwards. In terms of output, panel (a) shows that scenarios C and D (dashed- and dotted-lines, respectively) both deliver a positive trajectory (above the zero line), characterised by an upward trend from the tenth period onwards. This implies public investment crowding in output and private investment in the long run, at the cost of lower private consumption in the medium run. Panels (d) and (e) suggest, however, that the resulting higher level of demand brings the hours-work back to its steady state, while the higher TFP induces a higher real wage. In the end, the economy ends up displaying higher levels of output, private investment, wage, and public capital without compromising private consumption and employment.



Note: Variables expressed as percentage deviation from their trend. A, B and C, D denote the sets of measures described in Table 5. Shocks and parameters are calibrated as in Table 5 and at means of the Bayesian estimation. In panel (d), cases A and B, on one side, and C and D, on the other deliver the same hours-work paths (apparent solid line in the first case, and dashed-dotted line in the second).

Figure 9: Perfect-foresight simulations of the fiscal policy mix

The lesson to be drawn from this analysis is twofold. First, in the event of an expansionary policy, design matters. It is well known that public investment is productive, but in a setting such as the FPM1 a fiscal expansion is not useful for the economy. This finding echoes the claim raised by some studies focusing on the non-Keynesian effects of fiscal policy changes. The problem, however, is wider than just the mild efficiency of this policy (something that becomes clear from the comparison of the outcomes from A versus B, and from C versus D). What I show, is that the whole package of measures matters, and the right policy mix needs to be implemented as shown by the FPM2.

Second, in the earlier section, the household budget constraint implies the crowding-out effect concerning private consumption and investment if public investment rises and tax hikes. Yet, the simulations imply that the government could offset undesirable effects partially by investing raised funds, e.g. tax hikes, in highly efficient and productive capital projects rather than other expenditure plans. Furthermore, the government could finance new 'core' infrastructure projects by borrowing funds instead of solely relying on tax instruments as long as it can absorb the repayment burdens. Indeed, Abiad *et al.* (2014) pointed out that after the 2008 crisis the borrowing costs were low so that "the time was right for an infrastructure push" because of the positive economic outcomes in both short and long-term, assuming that efficiency investments will sustain the fiscal stand.

5.6 Conclusions

In conclusion, this study paints the panoramic picture that Vietnam's output fluctuations have been almost led by innovations in total factor productivity and shifting in household preference over the past three decades, 1986 – 2015. My findings show that there were significantly positive impacts of public investment on output movements in the early-to-mid 1990s, but that investment seems to be less productive in the 2000s.

The estimated DSGE model clearly outperforms the Hansen's (1985) RBC model in reproducing aggregate moments as it adequately captures several excess volatilities such as $\sigma_{\mathrm{cp}}/\sigma_{\mathrm{y}}$, σ_{ig}/σ_{y}, and $\sigma_{\mathrm{h}}/\sigma_{\mathrm{y}}$. It seems to understate, however, variations in private investment and government consumption.

The model also deems failing to capture contemporaneous correlation between investment and output because of the implied crowding-out effect of private investment in the household budget constraint.

Variance decomposition shows that shocks to TFP explain only 25% of private income fluctuations in the first-year impact but more than 42% from the third forecasting horizons. Hence, fiscal expansion via capital stock spending shall foster Vietnam income growth in both short and long run if the implementation is assumed to be productive and efficient. The IRFs are consistent with the IMF's evidence (Abiad *et al.*, 2014) that empirically stressed on the short- and long-term consequences of public investment on output growth. These IRFs apparently suggest the contemporaneous output gain of 0.91% (or 3% of that accumulated gain in five years) in exchange for a one-percentage point of the real GDP increase in the public investment. In addition, deterministic simulations suggest that public capital projects such as "core infrastructure" would crowd in private investment and output in the long run without compromising private consumption and hours-works if these projects are productive and implemented efficiently.

References

Abel, A. B. (2017) 'Crowding out in Ricardian economies', *Journal of Monetary Economics*. Elsevier B.V., 87, pp. 52–66. doi: 10.1016/j.jmoneco.2017.03.002.

Abiad, A., Furceri, D. and Topalova, P. (2014) 'Is it Time for an Infrastructure Push? The macroeconomic effects of public investment w: The time is right for an infrastructure push', *International Monetary Fund, Washington*.

Abiad, A., Furceri, D. and Topalova, P. (2016) 'The macroeconomic effects of public investment: Evidence from advanced economies', *Journal of Macroeconomics*, 50, pp. 224–240. doi: https://doi.org/10.1016/j.jmacro.2016.07.005.

Adjemian, S. *et al.* (2011) 'Dynare: Reference manual, version 4'. Dynare working papers 1, CEPREMAP.

Aiyagari, S. R., Christiano, L. J. and Eichenbaum, M. (1992) 'The output, employment, and interest rate effects of government consumption', *Journal of Monetary Economics*. Elsevier, 30(1), pp. 73–86.

Alvarez-Parra, F., Brandao-Marques, L. and Toledo, M. (2013) 'Durable goods, financial frictions, and business cycles in emerging economies', *Journal of Monetary Economics*. Elsevier, 60(6), pp. 720–736.

An, S. and Schorfheide, F. (2007) 'Bayesian analysis of DSGE models', *Econometric reviews*. Taylor & Francis, 26(2–4), pp. 113–172.

Arslanalp, S. *et al.* (2010) 'Public Capital and Growth', *IMF Working Paper: Public Capital and Growth*. International Monetary Fund, 10(175).

Aschauer, D. A. (1985) 'Fiscal policy and aggregate demand', *The American Economic Review*, 75(1), pp. 117–127.

Aschauer, D. A. (1989) 'Is public expenditure productive?', *Journal of monetary economics*, 23(2), pp. 177–200.

Baxter, M. and King, R. G. (1993) 'Fiscal policy in general equilibrium', *The American Economic Review*, pp. 315–334.

Blanchard, O. and Perotti, R. (2002) 'An empirical characterization of the dynamic effects of changes in government spending and taxes on output', *Quarterly Journal of Economics*, 117(4), pp. 1329–1368.

Boldrin, M., Christiano, L. J. and Fisher, J. D. M. (2001) 'Habit persistence, asset returns, and the business cycle', *American Economic Review*, pp. 149–166.

Bom, P. R. D. and Ligthart, J. E. (2014) 'What have we learned from three decades of research on the productivity of public capital?', *Journal of Economic Surveys*, 28(5), pp. 889–916.

Cavallo, E. and Daude, C. (2011) 'Public investment in developing countries: A blessing or a curse?', *Journal of Comparative Economics*, 39(1), pp. 65–81. Available at:

Christiano, L. J. and Eichenbaum, M. (1992) 'Current real-business-cycle theories and aggregate labor-market fluctuations', *The American Economic Review*, pp. 430–450.

Christiano, L. J., Eichenbaum, M. and Evans, C. L. (2005) 'Nominal rigidities and the dynamic effects of a shock to monetary policy', *Journal of political Economy*, 113(1), pp. 1–45.

Clarida, R., Gali, J. and Gertler, M. (2001) 'Optimal monetary policy in open versus closed economies: an integrated approach', *The American Economic Review*, 91(2), pp. 248–252.

Dreger, C. and Reimers, H.-E. (2016) 'Does public investment stimulate private investment? Evidence for the euro area', *Economic Modelling*, 58, pp. 154–158.

Hansen, G. D. (1985) 'Indivisible labor and the business cycle', *Journal of monetary Economics*. Elsevier, 16(3), pp. 309–327.

Herbst, E. P. and Schorfheide, F. (2015) *Bayesian Estimation of DSGE Models*. Princeton University Press.

Ilzetzki, E., Mendoza, E. G. and Végh, C. A. (2013) 'How big (small?) are fiscal multipliers?', *Journal of Monetary Economics*, 60(2), pp. 239–254.

IMF (2014) *IMF Country Report No. 14/311. Vietnam 2014 Article Iv Consultation*.

Iskrev, N. (2010) 'Local identification in DSGE models', *Journal of Monetary Economics*, 57(2), pp. 189–202.

De Jong, J. F. M., Ferdinandusse, M. and Funda, J. (2018) 'Public capital in the 21st century: as productive as ever?', *Applied Economics*, 50(51), pp. 5543–5560.

Justiniano, A., Primiceri, G. E. and Tambalotti, A. (2010) 'Investment shocks and business cycles', *Journal of Monetary Economics*, 57(2), pp. 132–145.

Kandil, M. (2017) 'Crowding out or crowding in? Correlations of spending components within and across countries', *Research in International Business and Finance*, 42, pp. 1254–1273.

King, R. G. and Rebelo, S. T. (1999) 'Resuscitating real business cycles', *Handbook of macroeconomics*, 1, pp. 927–1007.

Kydland, F. E. and Prescott, E. C. (1982) 'Time to build and aggregate fluctuations', *Econometrica: Journal of the Econometric Society*, pp. 1345–1370.

Leeper, E. M., Plante, M. and Traum, N. (2010a) 'Dynamics of fiscal financing in the United States', *Journal of Econometrics*, 156(2), pp. 304–321.

Leeper, E. M., Walker, T. B. and Yang, S.-C. S. (2010b) 'Government investment and fiscal stimulus', *Journal of monetary Economics*, 57(8), pp. 1000–1012.

Ogibayashi, S. and Takashima, K. (2017) 'Influence of Inefficiency in Government Expenditure on the Multiplier of Public Investment', *Computational Economics*, 50(4), pp. 549–577. doi: 10.1007/s10614-017-9671-y.

Pereira, A. M. and Andraz, J. M. (2013) 'On the economic effects of public infrastructure investment: A survey of the international evidence', *Journal of Economic Development*, 38(4), p. 1.

Pham, B. T., Hector, S. and Silva, J. I. (2019a) 'Growth and real business cycles in Vietnam and the ASEAN-5. Does the trend shock matter?', *Economic Systems*, manuscripted accepted.

Pham, B. T., Nguyen, C. P., and Hector, S. (2019b) 'Crowding-in or Crowding-out macroeconomic effect of public investment in Vietnam.', *Universitat Autonòma de Barcelona*, working paper.

Prescott, E. C. (1986) 'Theory ahead of business-cycle measurement', *Carnegie-Rochester Conference Series on Public Policy*, 25, pp. 11–44.

Schmitt-Grohé, S. and Uribe, M. (2012) 'What's news in business cycles', *Econometrica*. Wiley Online Library, 80(6), pp. 2733–2764.

Smets, F. and Wouters, R. (2003) 'An estimated dynamic stochastic general equilibrium model of the euro area', *Journal of the European economic association*, 1(5), pp. 1123–1175.

Smets, F. and Wouters, R. (2007) 'Shocks and frictions in US business cycles: A Bayesian DSGE approach', *The American Economic Review*, 97(3), pp. 586–606.

Stock, J. H. and Watson, M. W. (1999) 'Forecasting inflation', *Journal of Monetary Economics*, 44(2), pp. 293–335.

Uhlig, H. (1999) 'A toolkit for analysing nonlinear stochastic models easily', *Computational Methods for Study of Dynamic Economies*.

Uribe, M. and Schmitt-Grohé, S. (2017) *Open Economy Macroeconomics*. Princeton University Press.

Wickens, M. (2012) *Macroeconomic theory: a dynamic general equilibrium approach*. Princeton University Press.

World Bank (2005) *Connecting East Asia: A New Framework for Infrastructure*. World Bank Publications.

Appendix

Figure A.1

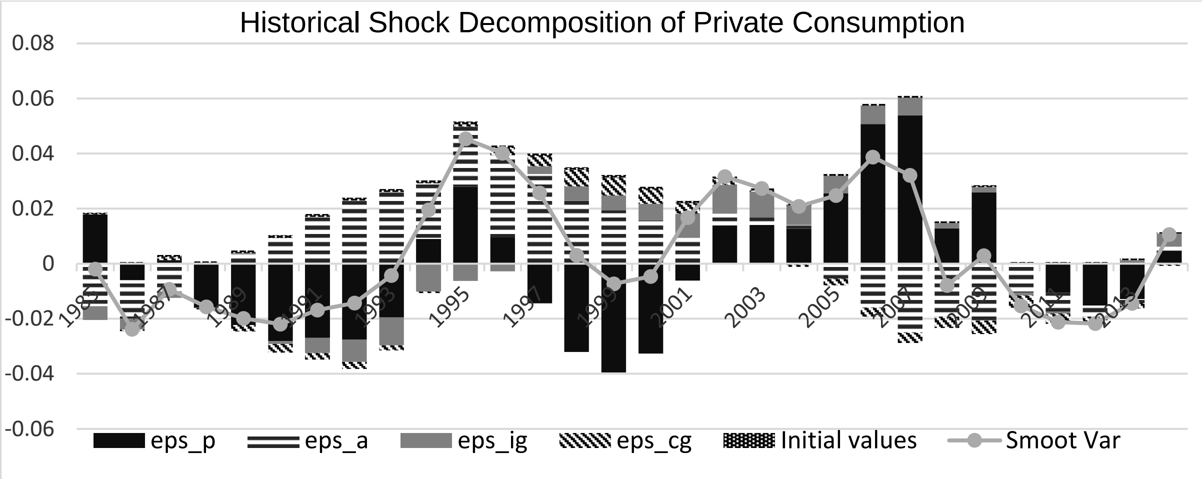


Figure A.2

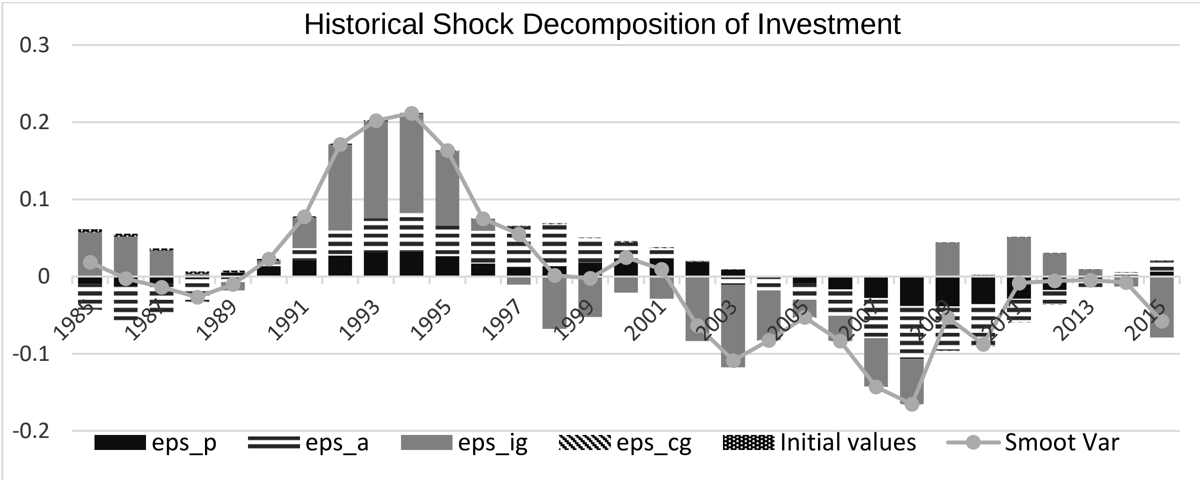


Figure A.3

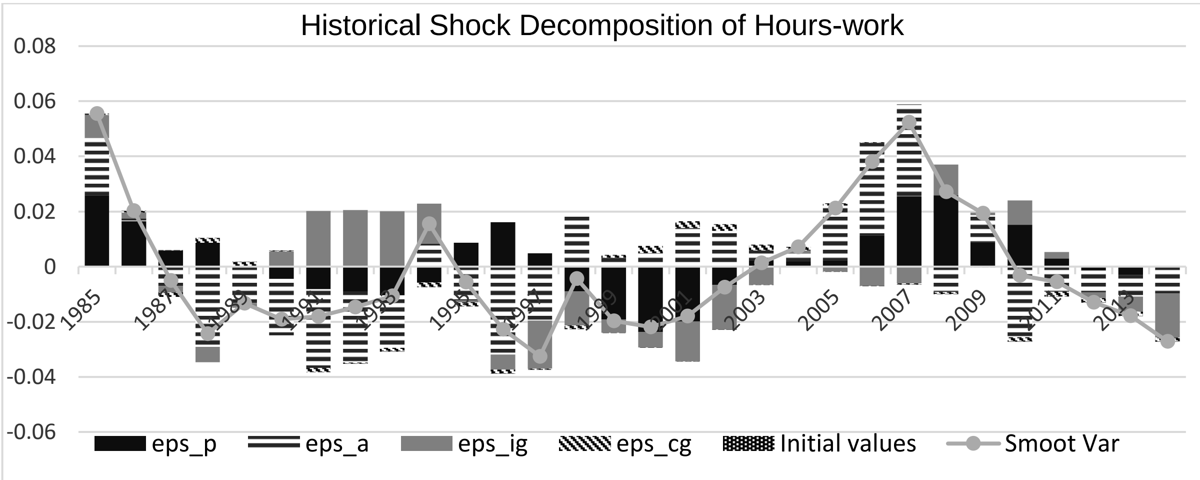


Figure A.4

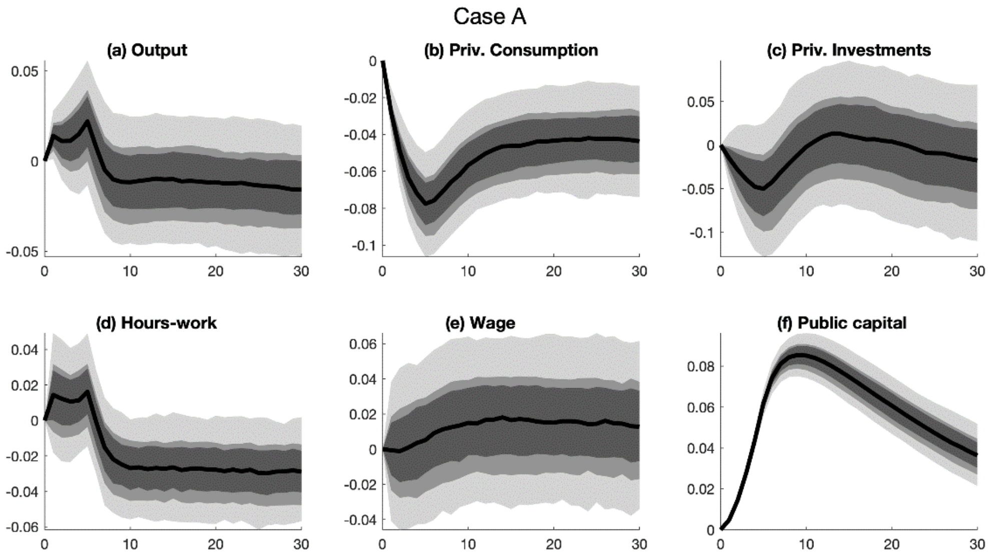


Figure A.5

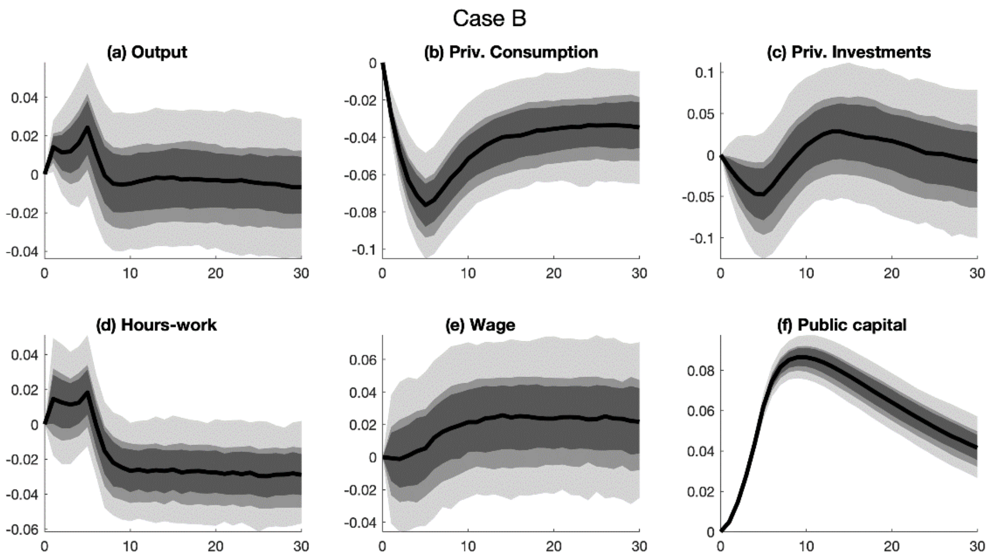


Figure A.6

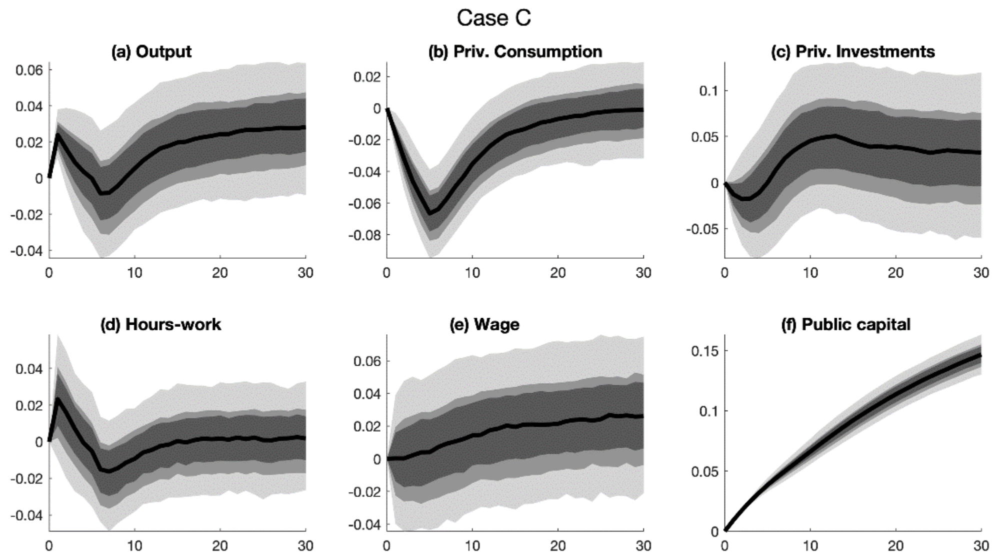


Figure A.7