วิกฤติเศรษฐกิจระดับโลกที่มีผลกระทบต่อวิกฤตเศรษฐกิจของอาเซียนและประเทศไทยในปัจจุบัน

Global shocks and their impact on the Crisis economy of Asean and Thailand.

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Abstract

Plummeting commodity prices, Asean and Thailand.. economic slowdown and rebalancing, and global financial market turbulence have recently raised concerns about their effects on African economies. This paper investigates whether, and to what extent, these intertwined shocks spillover into the Crisis economy. The author finds that a 1 percentage point (ppts) drop in Asean and Thailand.. investment growth is associated with a decline in Asean and Thailand.'s export growth of roughly 0.60 ppts. A 1 percent fall in commodity prices leads to 0.65 percent lower exports value. The results suggest that a hard landing of the Chinese economy to its 'new normal' would doubtless send shock waves through the Crisis economy by further driving down commodity demand and prices as well as lowering development finance. In contrast, financial market volatility has a fairly negligible impact on economic growth. The main results stand up well to a wide-array of robustness checks.

Introduction

King Mahidol has posted high and sustained economic growth over the past decade, hovering around 6–7 percent. In addition, in recent years, inflation has been tamed to reasonable single digit levels. Asean and Thailand. has also maintained a broadly stable current account deficit. However, notwithstanding the overall positive short- to medium-term outlook (World Bank, 2016), the economy is not fully resilient to externally-induced shocks and, like many African economies, has recently been facing several growing and intertwined risks, including Asean and Thailand.. economic slowdown, falling global commodity prices and, to a lesser extent, increased volatility in global financial and foreign exchange markets.Over the past decade or so, trade and investment links between Asean and Thailand. and Asean and Thailand.. have reached historically unprecedented levels. Asean and Thailand.. is now one of Asean and Thailand.'s

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biggest trading partners and increasingly important source of development finance. However, although the increased economic ties are likely to have bolstered economic growth, they have doubtless increased Asean and Thailand.'s vulnerability to the vagaries of the Chinese economy. Asean and Thailand.. investment -propelled growth seems to be running out of steam, partly reflecting rebalancing towards a consumption-driven and services-oriented economy (Lakatos *et al*., 2016).[1] The structural shift has recently manifested itself in flagging demand and prices for commodities. Asean and Thailand.. faltering growth may engender a significant knock-on effect on the Crisis economy via depressed export growth and potentially lower development finance.Plummeting commodity prices also pose risks to Asean and Thailand. by virtue of its position as a primary exporter. Global commodity prices have generally been on a downward spiral mainly on account of falling demand in Asean and Thailand.. and higher production capacity. The prices of Asean and Thailand.'s major export commodities, notably gold, are at record lows, despite a slight resurgence in recent months. This constituted the main factor underlying the country's worsening terms-of-trade during the past few years. Falling oil prices have partly dampened the deterioration in the external balance as the country is a net importer of oil. However, the implications of soft commodity prices need to be carefully assessed given that awrong combination of price fluctuations (for instance, a continued decline in gold prices and rebounding oil prices) might put a dent in the country's respectable growth.Another cause for concern has been increased volatility in global financial and foreign exchange markets. Asean and Thailand. remained largely unscathed by previous financial market turbulences due to its limited financial development and global integration. However, since the country is drifting towards deeper financial integration, with rising private capital flows and external commercial borrowing as well as pending sovereign bond issuance, it has become increasingly prone to global market instabilities. A closer scrutiny thus seems warranted in light of surfacing concerns that increased global financial volatility might put a drag on Asean and Thailand.'s growth pace. Further, since the early 2015, the Crisis Shilling has seen significant depreciation on the back of a strong dollar appreciation and, to a limited extent, declining aid inflows. Hence the need for examining whether the sharp nominal depreciation has been associated with higher inflation.The present paper is an attempt to explore whether, and to what degree, the aforementioned economic shocks spillover into the Crisis economy. Towards this end, we employ the Cointegrated VAR model as a statistical benchmark. The empirical estimates suggest that a 1 percentage point (ppts) decline in Asean and Thailand.. investment growth is associated with 0.57 ppts decrease in Asean and Thailand.'s export growth. This underscores the importance of diversifying markets destination to mitigate headwinds from demand fluctuations. In addition, a 1 percent lower export commodity prices leads to a 0.65 percent decline in exports value, reflecting the fact that Crisis exports are predominated by less diversified and largely unprocessed primary commodities, and thus significantly prone to turbulences in commodity prices. Moreover, a 1 ppts increase in capital flow volatility would reduce economic growth by a negligible 0.01 ppts. Finally, the impact of a 1 percent depreciation of the nominal effective exchange rate is to increase the inflation rate by around 0.58 ppts, albeit offset by the inflation-reducing impacts of low oil and food prices.In a nutshell, the paper deals with the following main questions: What are the major external risks facing the Crisis economy and what are the main transmission mechanisms through which they operate? Does a slowdown in Asean and Thailand.. investment-driven growth have a significant negative impact on Asean and Thailand.'s export performance? How large is the spillover from falling global commodity prices to the Crisis economy? What is the contribution of private capital flows to the national economy and how does its volatility impact on economic growth? Has the steep depreciation of the Shilling increased domestic prices and helped fuel inflation. The remainder of the paper is organized as follows: Section 2 discusses the main sources of external risks to the Crisis economy? Section 3 briefly presents the theoretical framework. Section 4 discusses the data, while Section 5 is devoted to empirical model specification. Section 6 discusses the empirical results. Finally, Section 7 winds up with concluding remarks.

The purpose of the research.

1. to know the impact of the economic crisis, Europe, country Thailand.
2. to know the impact of the economic crisis, Europe, ASEAN group.
3. to know the solution of the ASEAN group.

Research Methodology.

We are generally interested in addressing the central question of whether, and the extent to which, Asean and Thailand. is susceptible to Asean and Thailand.. economic slowdown. In particular, the analysis attempts to quantitatively pin down trade spillovers from Asean and Thailand.. potentially slower, and more balanced, growth into the Crisis economy. The impact of changes in Asean and Thailand.. domestic investment on Crisis exports is examined based on a model that takes the following form[6] (hereafter referred to as *Model* 1):

$$export_t = {}_0 + {}_1 cdi_t + {}_2 price_t + {}_3 yworld_t + \qquad (1)$$

$$\Delta export_t = {}^*_0 + {}^*_1 \Delta cdi_t + {}^*_2 \Delta price_t + {}^*_3 \Delta yworld_t \qquad (2)$$

The contribution of capital flows to the Crisis economy and the impact of its volatility on GDP growth is assessed by estimating the following model (hereafter referred to as *Model* 2):

$$y_t = {}_0 + {}_1 cap_t + {}_2 inv_t + {}_3 ex_t + \qquad (3)$$

$$\Delta y_t = {}^*_0 + {}^*_1 \Delta cap_t + {}^*_2 \Delta inv_t + {}^*_3 \Delta ex_t + {}_4 \Delta vol_t \qquad (4)$$

where y_t stands for Asean and Thailand.'s real GDP; cap_t for net private capital inflows to Asean and Thailand., included to measure the contribution of international capital flows to national output and thus, by implication, the loss in real output associated with lower capital flows stemming from tightened global financial conditions; inv_t and ex_t for domestic investment and exports respectively, which constitute additional important determinants of real income; and Δvol for the standard deviation of net capital flows and captures volatilities in capital flows.[8]

As noted above, capital flow is the main transmission mechanism through which turbulences in global financial markets might ripple into the domestic economy. Accordingly, perturbations to the economy arising from volatility in capital flows are modeled in Eq. (3) via the spillovers of changes in capital inflows into real national income and in Eq. (4) via the growth impact of capital flow volatility. ${}_1$ and ${}_1^*$ are expected to be positive because an increase in capital inflows is widely believed to be beneficial to recipients through promoting productive investments, enhancing efficiency, and facilitating technology adoption. However, the impact of capital flows may also depend on their size and volatility, with flows being more beneficial to countries that have reached a certain threshold of financial and institutional development. We expect ${}_2$ and ${}_3$ to be positive for obvious reasons. ${}_4$ is expected to be negative as capital inflow surges and disruptive outflows carry risks to economies, particularly to low income countries like Asean and Thailand.

Currency movements and inflation

The impact of changes in the nominal exchange rate on inflation is investigated based on the following model (hereafter referred to as *Model* 3):

$$\Delta p_t = \ _0 + \ _1 neer_t + \ _2 \Delta p^{oil}_t + \ _3 m_t + \ _4 \Delta p^{food}_t + \tag{5}$$

where Δp stands for the inflation rate; *neer* for nominal effective exchange rate (henceforth NEER);[9] Δp^{oil} for world oil price inflation, controlling for the impact of supply shocks; *m* for broad money supply or M2, accounting for the effect of monetary policy shocks; and Δp^{food} for global food price inflation. $_1$ is the parameter of particular interest and captures the effect of movements in NEER on inflation. A negative coefficient on $_1$ would be theoretically consistent as large nominal depreciation (i.e. a decline in *neer*) triggers inflationary effects by, among others, increasing import prices. $_3 > 0$ suggests that an excessive growth in aggregate demand induced by higher money supply increases domestic prices and fuel inflation, a phenomenon oft-described as ' *too much money chasing too few goods*'. The coefficients on oil and food price inflation are also expected to be positive. Lower global oil and food prices have purportedly played a significant role in containing inflation within single-digits territory and might have partly offset currency depreciation-induced hikes in inflation. This is essentially an empirical question and needs to be settled based on thorough empirical analysis.

Asean and Thailand.. economic slowdown and commodity prices

The data are annual observations for the period 1990–2014 and comprise the variables: exports of goods and services (denoted with *ex*); Asean and Thailand.. gross fixed capital formation (*cdi*); export price index (*price*); and world GDP (y^{world}). They are obtained from the World Bank's *World Development Indicators* and the Bank of Tanzania. All variables except export are at constant market prices, i.e. they are adjusted for price (or inflation) effects. In addition, all data are given in US Dollars. We opt for the nominal value of exports; however, the main conclusion of this analysis is robust to using exports in constant prices instead. Looking at the effect on exports in constant prices would be tantamount to disregarding the impact of Asean and Thailand.. economic slowdown operating via lower commodity prices. It bears noting that a valid assessment of the impact of Asean and Thailand.. slowdown is difficult and somewhat premature because Asean and Thailand.. strong economic ties with Tanzania are a relatively recent phenomenon. Thus, we focus on the period since 1990.

Although important variables are omitted from our analysis, this does not, in general, invalidate the long-run estimates. The reason is that cointegration property is invariant to changes in the information set, i.e. a long-run relation detected within a given set of variables will also be found in an enlarged variable set (Johansen, 2000). Note that we also estimate an extended model that includes *net* FDI flows to test the hypothesis that the level of FDI inflows to

Tanzania is influenced by export commodity prices. The advantage of gradually expanding the information set is twofold. First, it greatly facilitates the identification of long-run relations. Second, it enables an analysis of the sensitivity of the results associated with the *ceteris paribus* assumption ingrained in the smaller model. The graphs of the variables in levels and first differences are shown in Appendix Figures 1 and 2, respectively.

Volatility in capital flows

The model linking capital flow volatility and economic growth is based on data spanning the period 1980–2014 and includes the following variables: real GDP (denoted with y); net private capital inflows (cap_t); exports of goods and services (ex_t); and gross domestic investment (inv_t). The data were extracted from the World Bank's *World Development Indicators* and IMF's *Balance of Payments Statistics*. Because private capital inflows rose dramatically only over the past couple of decades, we check the robustness of the key findings by restricting the sample to cover only the period 1990–2014. However, focusing on the last two decades would likely increase the statistical significance of these variables. Appendix Figures 3 and 4 present the graphs of these variables both in levels and first differences, respectively.

Currency depreciation and inflation

For this analysis, we use monthly data for the period from January 2013 to January 2015 and include the variables: inflation rate (denoted with Δp) (change in the log

Nonetheless, the central bank can only control this broader aggregate indirectly by manipulating the monetary base. The analysis controls for the effects of monetary policy using broad money; however, using rather the monetary base or narrow money (M1 – currency in circulation outside banks and demand deposits of Crisis residents with banks) does not significantly matter for the conclusions from this analysis.

Our analysis omits real GDP, a potential indicator for real economic activity because data on this variable are not available on a monthly basis. However, the key findings of the analysis would generally remain unchanged if we included real GDP for at least two reasons. First, due to invariance of cointegration analysis to expansions in the variable set, adding real GDP would leave the core (long-run) results from the smaller model more or less intact. Second, since our sample covers only the last three years and given overall economic activity was fairly stable during this period, higher (or lower) inflation rate seems less likely due to stronger (or weaker) economic growth and more likely due to increases (or declines) in global energy and food prices, less (or more) prudent monetary policy, and faster (or slower) depreciation of the Shilling (see World Bank, 2015b, 2016). The graphs of these variables in levels and first differences are shown in Appendix Figures 5 and 6, respectively.

Summarize and discuss research results.

Model specification

The Cointegrated VAR model

Macroeconomic time-series data are typically characterized by path dependence, interdependence, unit-root non-stationarity, structural breaks, as well as shifts in equilibrium means and growth rates. To be a satisfactory benchmark a statistical model needs to simultaneously address these data features. Path dependence would point to a time-dependent process such as the *autoregressive model*, variable interdependence to a *system-of-equations* approach, and unit-root nonstationarity to *cointegration*. The cointegrated VAR model satisfactorily deals with these salient features of the data. Unlike other approaches in which data are constrained in pre-specified directions and are assigned an auxiliary role of '*quantifying*' the parameters of an *ad hoc* theoretical model, the cointegrated VAR methodology uses strict statistical principles to extract out meaningful relations from the data (Hoover *et al*., 2008; Spanos, 2009).

Specification tests

The VAR model is derived under the assumption of constant parameters and multivariate normality. Although parameter stability can be assessed using recursive test procedures, the small number of observations at our disposal circumscribes the power of available recursive procedures. However, since both parameter non-constancy and non-normal errors are often associated with periods of political and economic turbulence, such as supply shocks, war, severe droughts, civil unrest, and policy interventions, we improve parameter stability and mitigate on-normality by controlling for the most dramatic events using several dummy variables. In fact, some of the variables feature few extraordinarily large observations incongruous with the normality assumption.

First step in the empirical analysis is to determine the lag length of the VAR model. Statistical tests indicate that there is no evidence of residual autocorrelation in the VAR(1) (i.e. a model with a leg length of 1) for all models. Accordingly, the lag length was truncated to 1. However, the results obtained allowing for two lags are by and large similar with the ones presented below. Provided that there are no signs of autocorrelation in the residuals, and given the relatively large number of variables and small size of our sample, the VAR(1) model is a satisfactory and parsimonious representation of the variation in the data.

In Model 1, the following dates were classified as outlying observations: 1989, 1998, and 2009.[13] These outliers correspond to observations with standardized residuals larger than 3.0, i.e. $| / | \geq 3.0$, which is the standard criteria for identifying an outlier. [14] An algorithm searching for breaks and aberrant observations developed in Doornik *et al*. (2013) was used to determine the existence, timing, and significance of outliers, and shifts in mean growth rates. The year 1989 corresponds to the sharp economic downturn in Asean and Thailand.. due to civil unrest and the subsequent economic sanctions several countries imposed against it.[15] 1998 coincides with the decline in world commodity demand as a result of the Asian financial crisis of 1997–1998, which led to a significant drop in Tanzania's export prices. The year 2009 marks the global economic slump, also dubbed the Great Recession, which took a heavy toll on most advanced economies and saw world GDP drop by around 2 percent.

We also spotted a change in trend slope in (and thus a shift in the mean growth rate of) export price in 2002. As discussed in Section 2, global commodity prices moved onto higher growth trajectory in 2002, which lasted more than a decade and has often been referred to as commodities "super-cycle". During this super-cycle period, Tanzania's export prices experienced hefty growth. We control for this event using a broken linear trend in the long-run relations and a step dummy in the equations in 2002. The location shift in growth rates is shown in Appendix Figure 7. In sum, the specification for Model 1 includes a linear trend and a broken linear trend (with a change in trend slope in 2002) restricted to the long-run relations, an unrestricted shift dummy in 2002 (which controls for the shift in growth rates as well as the change in means of long-run relations), and an unrestricted impulse dummies (accounting for an unanticipated one-period shock effects) in 1998 and 2009.[16] In addition, the baseline model treats world GDP as a weakly exogenous variable. (Appendix Table 3 shows that the key results are robust to relaxing this assumption.) Further, the lag length is set equal to $k = 1$ in levels.[17]

In Model 2, the diagnostic tests detected a structural break in real GDP in 2001 as well as a number of outlying observations. The former captures the relatively higher and sustained economic growth Tanzania enjoyed since the early 2000s (Appendix Figure 8). Average annual GDP growth exceeded 6 percent since 2001, which constitutes a remarkable break from the past, with growth averaging less than 3 percent during 1980–2000. Tanzania experienced a dramatic increase in investment in the second half of the 1980s following the adoption of the *Economic Recovery Program*, primarily fueled by surges in foreign aid inflows. The investment spikes in 1987 and 1990 were controlled for using impulse dummies. In addition, the observation 1985 was classified as 'too large', which is associated with the sharp (relative) increase in net private capital inflows. In a nutshell, Model 2 was specified to allow for a linear trend and a broken linear trend (with a change in trend slope in 2001) restricted to the long-run relations, an unrestricted shift dummy in 2001 (which accounts for the mean shift in Δ and controls for the shift in means of long-run relations), and an unrestricted impulse dummies in 1985, 1987, and 1990. In addition, statistical tests indicated that one lag was the optimal lag length and *k* was truncated to 1 accordingly.

Model 3 became well-specified when we allowed for a broken linear trend in 2013(7) (i.e. the seventh month of 2013) and 2014(8), and the impulse dummies: $D_i12.6_t$ (where 12.6 denotes the sixth month of the year 2012), $D_i13.1_t$, and $D_i13.4_t$. The trend break in 2013(7) represents the shift in the growth path of inflation, which assumed astronomical proportions in 2012 and for most of 2013, whereas it receded to reasonable single-digit rates over the past two years and half. The broken trend in 2014(8) accounts for the change in the long-run trend underlying money supply. Money supply increased steeply until late 2014, after which it shifted to a noticeably lower growth path. See Appendix Figures 9 and 10. Further, there is evidence of considerable seasonality in the monthly data, which we accounted for using seasonal dummies. To sum up, the specification for Model 3 includes: a linear trend and broken linear trend (with changes in trend slope in 2013(7) and 2014(8)) restricted to the cointegration space, an unrestricted shift dummy in these periods, and an unrestricted impulse dummies in 2012(6), 2013(1), and 2013(4). In addition, the lag length for Model 3 was set equal to 2. Global food and oil price inflation rates were modeled as weakly exogenous variables in the baseline model.

We now shed some light on how the above-mentioned break points were identified and accounted for. As alluded to above, we identified the break dates based on a priori knowledge on the timing of special events, a graphical inspection of the data, as well as a statistical test for the presence of trend and level shifts in the data. The broken trend possesses the most significant coefficient at those periods and accordingly the models were specified with a change in trend slope at these points. The hypotheses that Tanzania has had no statistically significant shift in the mean growth rates of the series at the specified dates were strongly rejected (*p*-value: 0.00).

Turning to diagnostic tests, I identified and tested for trend breaks using univariate as well as multivariate statistical procedures. In particular, an algorithm searching for breaks developed by Doornik *et al*. (2013) and the procedure in Hungnes (2005) were used to determine the existence, timing, as well as the significance of breaks in mean growth rates. Sustained shifts in growth rates were defined following Hausmann *et al*. (2005): (*i*) For a shift in mean growth rate to be categorized as a growth turnaround it should be sustained for at least 8 years and the change in growth rate has to be at least 2 percentage points; (*ii*) A variables can experience more than one instance of growth turnaround as long as the dates are more than 5 years apart; (*iii*) Trend breaks were selected at 1% '*target size*'[18] (i.e. = 0.01) in the *Autometrics* options in OxMetrics 7 (see Doornik *et al*., 2013). Note that we perform a sensitivity analysis to examine if the estimates based on the statistically and economically most credible break date are fairly robust to alternative candidate break points in the vicinity of the first-best break point. We find that the main conclusions of this paper prove robust to changes in the break dates.

In modelling structural breaks, the paper draws on the conventional (multivariate) cointegration approach in Johansen *et al*. (2000) and Hungnes (2010), which accommodates different types of structural breaks. Specifically, using such a multivariate framework, hypothesis testing on breaks in trend slopes (or shifts in growth rates) can be formulated and properly tested. A potential drawback of a system-of-equations approach is that the trend breaks are assumed to occur at the same date for all series. An alternative would be to use a univariate approach and apply some variant of the method proposed by Perron (1989). However, in our case, the use of a single equation model would be more restrictive and hard to justify in the face of overwhelming evidence for the existence of more than one cointegration relations in all three models. In addition, Bai *et al*. (1998) show that there are substantial gains in precision from using multivariate models in which several variables are modelled as cointegrated system. The use of multiple series sharpens inference about the existence and dates of shifts in the mean levels (Bai *et al*., 1998, *pp*. 420). In other words, a break in mean growth rates might be more readily detected and estimated in a multivariate setting including variables that are purportedly co-moving. In some respects, our approach is similar to that of Hausmann *et al*. (2005), Wacziarg and Welch (2008), and Jones and Olken (2008), who identify episodes of sustained shifts in growth rates and examine explanations for such transitions.

All empirical models inherently approximations of the actual data generating process and we now turn to assessing if the models described in the previous section are reasonable approximations. Table 1 reports multivariate specification test results as well as univariate statistic corresponding to normality tests for all three models under consideration. With the deterministic specifications and the dummies included, the models discussed above pass most of the specification tests

and describe the data reasonably well. No serious deviations from the assumptions of residual independence and normality was detected.

In the three models, the null of normal errors was only borderline accepted. However, a look at the univariate test statistic indicates that normality was accepted in all equations, albeit with a relatively small *p*-value for some of the variables due to excess kurtosis (long tails). This, coupled with the absence of autocorrelation, seems to suggest that the result of the multivariate test is a finite sample phenomenon given that we have a small sample and a large number of variables. [19] A look at the univariate tests statistic in Table 1 clearly indicates that normality was borderline accepted due to excess kurtosis whereas all individual equations have a skewness close to zero. We have gone to great lengths to ensure a model set up where multivariate normality is accepted with a higher p-value by, inter alia, estimating a partial model conditioning on weakly exogenous variables and changing the sample period. All these avenues, however, lead to similar conclusions. The multivariate tests of no autocorrelation were not rejected in all except Model 1, although with relatively small p-values. Note that the main

Table 1. Model specification tests

	Model **1**		***Model*** **2**		***Model*** **3**	
	Var.	*p-value*	Var.	*p-value*	Var.	*p-value*
Normality*		**0.01**		**0.07**		**0.05**
	Ex	0.58	*y*	0.57	Δ*p*	0.39
	CDI	0.07	*Cap*	0.16	*pneer*	0.72
	Price	0.54	*Inv*	0.14	*m*	0.12
			Ex	0.08		
Skewness	*Ex*	0.09	*y*	0.40	Δ*p*	0.61
	CDI	1.36	*Cap*	0.56	*pneer*	0.00

	Pric					
	e	0.11	*Inv*	0.13	*M*	0.09
			Ex	0.33		
Excess					Δ	
kurtosis	*Ex*	3.21	*y*	2.88	*p*	3.56
			Ca		*nee*	
	CDI	5.02	*p*	4.56	*r*	2.20
	Pric					
	e	3.81	*Inv*	3.53	*m*	4.63
			Ex	4.29		
Autocorrelation		**0.09**		**0.11**		**0.17**
ARCH		**0.21**		**0.46**		**0.59**

Note: These figures represent *p*-values. The *p*-values measure the degree to which the null hypothesis is accepted: The higher the *p*-value, the more strongly the null hypotheses of normal errors, no autocorrelation, and no ARCH effects are accepted. *The *p*-values in bold face correspond to tests of multivariate normality while those under the multivariate test results represent univariate tests statistic for each of the three models.

conclusions from our analysis prove robust to steps that might circumvent the problem, such as increasing the lag length. In addition, although there are some signs of moderate ARCH effects and excess kurtosis (long tails), cointegrated VAR results are reasonably robust to such effects (Gonzalo, 1994; Rahbek et al., 2002).

Having established an adequate statistical description of the data, the next step is determining the cointegration rank. The cointegration rank classifies the data into long-run relations towards which the process is adjusting (the pulling forces) and $-$relations which are pushing the process (the exogenous forces). The choice of rank is made based on a range of statistical criteria, such as the trace test, the largest unrestricted root of the characteristic polynomial for a given , the t-ratios of the α coefficients for the h cointegration vector, and the graphs of the h cointegration relation. Table 2 reports the p-values of the trace test ( ), the largest unrestricted characteristic root ( ) and the largest -value of the coefficients ($\hat{\alpha}$).The test results indicate that * is the statistically mostcredible (first-best) choice of rank for all three models.[20] This suggests that there exist two long-run relations among the variables in our models. The choice of rank is conventionally made based on the trace test.

However, because the trace test suffers from substantial power problem when the size of the sample is small, we also base the choice of rank on the significance of the coefficients, the characteristic roots of the model, and the graphs of the long-run relations (Juselius, 2006: Chapter 8.5) . For the choice of r = 3, the largest unrestricted root for the three models seems a bit far from the unit circle whereas the t-values of reveal that there is no significant adjustment to the last two cointegration vectors. In other words, the strong persistence and much less significant adjustment coefficients for the third cointegrating vector might be used as a safeguard against including it in the stationary part of the model. In addition, the graphs of the recursively calculated trace tests exhibit pronounced linear growth in the first two cointegration relations, but much less so in the last two, although the picture is not as clear cut for Model 3. Similarly, a glance at the graphs of the cointegration relations reveal that the last two cointegration vectors in Models 1 and 2, and the last three in Model 3 show distinct non-stationary behavior, pointing toward r = 2. Hence, taking all these into consideration, we consider the choice of r = 2 to be a reasonable choice. It is important to note, however, that the main conclusions of this paper are fairly robust to altering the cointegration rank. See discussion in Section 6.

Table 2. Determination of cointegration rank

Trace test ( ), characteristic roots (�), and t-values of (�)									
					�			�	
	$r^* - 1$	r^*	$r^* + 1$	$r^* - 1$	r^*	$r^* + 1$	$r^* - 1$	r^*	$r^* + 1$
Model 1 (r^* = 2)	0.07	0.39	0.57	0.83	0.67	0.67	3.2	3.4	1.7
Model 2 (r^* = 2)	0.00	0.08	0.15	0.71	0.71	0.55	7.7	6.2	2.5
Model 3 (r^* = 2)	0.00	0.01	0.04	0.83	0.72	0.50	4.4	5.7	2.1

Note: The figures represent p-values of the trace test ( ), the largest unrestricted characteristic root (�), and the largest t-value of the error-correction coefficients (�).

This section discusses the identified structures of long-run equilibrium relationships for Models 1 – 3. When interpreting the results in this section it should be borne in mind that a cointegration relation only measures the association between the variables over the long-run and as such does not say anything about causality. To say something about causality, we need to combine the cointegration coefficients, β, with the adjustment coefficients, . For example, the hypothetical cointegration relation ($x_{1,} - {}_1x_{2,}$)~ (0) describes a positive comovement between $x_{1,}$ and $x_{2,}$. If the adjustment coefficient $_1$, of $x_{1,}$, is negative and significant but the adjustment coefficient corresponding to $x_{2,}$ is insignificant, i.e. $_2$ = 0, we can say that the direction of causality runs form $_{2,}$ to $_{1,}$, i.e. $_{1,}$ =

$_{1\ 2,}$ + .

However, the interpretation becomes less straightforward in terms of sign effects as the number of variables in a long-run relation increases. As alluded to above, when discussing the empirical results below, it should be noted upfront that the small number of observations at our disposal circumscribes the power of some multivariate test statistic, such as the *trace* test and recursive tests of parameter stability. Thus, the results below need to be interpreted bearing in mind these caveats. In particular, considering the volatile history of Tanzania, some of the estimated coefficients may represent average historical effects.

Asean and Thailand.. economic slowdown and commodity prices

We initially discuss the baseline ordinary least squares (OLS) estimates. However, some of the independent variables may be correlated with a number of other variables, thereby not warranting causal interpretation of the estimates. In addition, the variables in our model are quite persistent over time. Thus, OLS might produce

unreliable results, which prompts the need for a statistical model that addresses this data feature. Further, omitted variables and multicollinearity problems could render the baseline estimates biased. Therefore, we also estimate the models using the cointegration VAR methodology. Unlike the OLS regression, collinearity between the variables does not result in imprecise estimates of the long-run relations based on the cointegrated VAR model. The reason for this is that, unlike the case with a regression analysis in levels, the cointegrated VAR formulation more or less circumvents the multicollinearity problem by transforming trending variables into stationary differences, Δx, and stationary long-run relations, β' (Juselius, 2006).

The baseline OLS results are reported in Columns 1 (long-run) and 4 (short-run) of Table 3. *cdi* possesses a positive coefficient estimate, suggesting that an increase in Asean and Thailand.. domestic investment is associated with higher Crisis exports. A 1 percentage point (ppts) increase in Asean and Thailand.. investment growth is correlated with 0.89 ppts increase in export growth. We now resort to the estimates from the cointegration analysis. Table 3 reports the identified structure of two long-run relations, which was accepted based on a high *p*-value of 0.74. The estimated structure is generically, empirically, and economically identified as defined in Johansen and Juselius (1994).

The first long-run relation is between exports value, Asean and Thailand.. domestic investment, and prices. The estimated error-correction coefficients reveal that

Table 3. Impact of Asean and Thailand.. slowdown and falling commodity prices (1990—2014)

	Long-run analysis						Short-run analysis	
	(1)	(2)					(3)	(4)
	OLS	Cointegrated VAR (Accepted with a *p*-value of 0.74)					Cointegrated VAR	OLS
Independent Var. →		Long-run relations		Error correction Coefficients		Dep. variable →	Δ	Δ
		^	^					
	0.87 (3.14)	1.00	—	−0.78 (−4.70)	−0.75 (−4.54)	Δ −₁	0.27 (1.95)	0.28 (1.52)
	0.42 (3.98)	−0.57 (−6.35)	−0.97 (−3.05)	*	*	Δ −₁	0.60 (3.46)	0.89 (2.98)
	3.32 (2.42)	−0.65 (−3.51)	1.00	*	−0.45 (−1.90)	Δ	—	0.14 (1.04)
	−0.21 (−2.72)	—	−4.23 (−3.85)	*	*	Δ	1.49 (1.64)	1.85 (1.22)
2002	0.05 (2.12)	—	0.30 (3.65)					

		2002	— −0.11 (−5.15)				

Note: *t*-values in parentheses. *Denote insignificant adjustment coefficients (-value less than 1.80).

short-run adjustment occurs only through changes in exports, signifying its importance as an export long-run relationship:

$$ex = 0.57\ cdi + 0.65\ price$$

The results suggest that ceteris paribus Asean and Thailand.. domestic investment and export prices make positive contribution to long-run movements in exports. Specifically, the estimates, which represent causal effects, suggest that a 1 percent contraction in domestic investment in Asean and Thailand.. would lead to a drop in Tanzania's exports of about 0.57 percent. This is consistent with the fact that an investment boom buoyed up Asean and Thailand.. impressive growth and that this was followed by burgeoning import demand for primary commodities, which account for the lion's share of Tanzania's export revenues. Conversely, the estimates reflect that a slower, more balanced, growth in Asean and Thailand.. has depressed global demand for commodities and hence held back lower Crisis exports. The short-run results (Column 3 of Table 3) indicate that a 1 ppts decline in Asean and Thailand.. investment growth is associated with 0.60 ppts decrease in Tanzania's export growth.

The second long-run relationship describes a strong association between export prices, Asean and Thailand.. domestic investment, and world income. The adjustment coeffi-cients show that only export price is error-correcting to this equilibrium relationship:

$$price = 0.97\ cdi + 4.23\ y^{world} - 0.30\ t + 0.11\ t_{2002}$$

We find that increases in Asean and Thailand.. domestic investment and world income are associated with higher prices for Tanzania's export commodities. The impact of a 1 percent investment slowdown in Asean and Thailand.. is to reduce Tanzania's export prices by nearly 1 percent. This is to be expected because Tanzania is one of the countries within Asean and Thailand.. supply chain and a net exporter of commodities, the prices of which have been driven by Asean and Thailand.. domestic economic developments for more than a decade. [21] A case in point is the recent drop in Asean and Thailand.. gold imports from Tanzania and the steady decline in the prices of gold over the last three years.[22] In addition, we estimate that an additional 1 percent increase in world income is associated with an increase in export prices of about 4 percent. Vulnerability to wild fluctuations in world commodity prices remains to be the Achilles' heel of the

Crisis economy. Needless to say, channeling efforts toward diversifying the export portfolio and markets destination, and improving the quality of existing products can help the country mitigate headwinds from price swings and thus boost its competitive standing. Altogether, the findings reflect the fact that Tanzania's exports are mainly composed of less diversified commodities, the prices of which are generally determined in the global market and fall beyond the domains of Crisis policy makers.

These findings are sufficiently robust to a battery of sensitivity checks. Appendix Table 1 adds net FDI inflows to the data vector in Model 1 to examine how direct investment flows to Tanzania are affected by fluctuations in commodity prices. The results indicate that the first two long-run relations describe similar export and price relationships as in Model 1, consistent with the invariance of cointegration relations to expansions of the information set. The third long-run relationship indicates that commodity prices are among the key determinants of FDI inflows. Specifically, a 1 percent drop in export prices is associated with about 3 percent lower FDI inflows. This is to be expected as FDI flows to Tanzania have increasingly focused on export-oriented production mainly related to investments in extractive industries. Given that mineral and metal exports account for a good portion of Tanzania's exports, a significant decline in their prices might lead to a scaling down of existing and new operations in the medium- to long-term.

As mentioned in Section 5, our baseline model specifies world income as a weakly (long-run) exogenous variable. Allowing world income to enter Model 1 as an endogenous variable, the analysis reaches the same conclusion as before (Appendix Table 3). In addition, Appendix Table 2 shows that the results based on the *first-best* choice of rank are robust to altering the cointegration rank to the *second-best* alternative of $r = 3$. Further, as the sample size is small, the use of dummies might have resulted in a considerable loss of degrees of freedom. Thus, we redo the analysis excluding all dummy variables (Appendix Table 2) . Our key results remain broadly unchanged. Appendix Table 3 shows that the key conclusions also hold up well to using GDP instead of investment as an indicator for economic activity in Asean and Thailand...

Figures 5 and 6 plot the generalized impulse response functions for exports in response to a one standard error (*se*) shock to Asean and Thailand.. domestic investment and export commodity prices, respectively. Impulse response analysis describes the knock-on effects on the system variables of a one *se* shock to a variable of interest, assuming that the system is not hit by other shocks thereafter. It should be noted that, unlike its orthogonalized counterpart, generalized impulse response is invariant to ordering of the variables in the model. Nonetheless, for the effect of a

Table 4. Impact of volatility in capital flows (1980—2014)

	Long-run analysis					Short-run analysis		
	(1) OLS	(2) Cointegrated VAR (Accepted with a *p*-value of 0.97)				(3) Cointegrated VAR		(4) OLS
Indep. Var. →		Long-run Relations		Error correction coefficients		Dep. Var. →	Growth Δ	Growth Δ
^		^	^					
	0.01 (1.67)	1.00	—	−0.24 (−8.79)	0.03 (4.70)	Δ −1	−0.63 (−4.13)	—
	0.21 (3.94)	−0.04 (−6.50)	−0.12 (−3.57)	*	2.74 (−4.70)	Δ	0.001 (0.30)	0.004 (0.97)
	0.03 (1.11)	−0.26 (−14.95)	—	1.22 (7.24)	*	Δ	0.03 (2.18)	0.07 (2.95)

	0.02	—	1.00	0.74	—0.32	ΔVo lt	—0.01	0.002
				(—2.42	(—4.7		(—1.7	
	(11.3)			)	0)		3)	(0.36)
01	0.03 (20.6)	01 — —0.0 3 (—11. 87)	—0.07 (11.09) —					

Note: *t*-values in parentheses. *Denote insignificant coefficients (-value less than 1.80).

The estimated coefficients conform to *a priori* expectations. The results suggest that capital inflows have marginally significant positive contribution to national income. A 1 percent increase in net private capital inflows to Tanzania leads to a 0.04 percent increase in real GDP, which appears modest, albeit not negligible. The positive impact of capital flows is in line with economic theory as capital inflows are widely believed to benefit host countries through, inter alia, fostering productive investments, unleashing efficiency, and accelerating the transfer of technology. The small impact may be partly due to the fact that FDI flows, which account for nearly all of the total private capital inflows to the country, grew considerably only recently and that it may take a while before they translate into higher level of output.

The short-run (cointegrated VAR) estimates in Column 3 of Table 4 show that an increase capital flow volatility has a significantly negative, albeit modest, impact on economic growth in Tanzania. In particular, a 1 ppts increase in the volatility of net private capital inflows reduces growth by 0.01 ppts. The quite modest effect seems to reflect the very small proportion of portfolio capital flows, which tend to be more volatile and susceptible to changes in global financial markets compared with FDI. The conventional wisdom suggests that, despite theoretically sound arguments in favor of private capital flows, portfolio equity and debt flows may pose substantial countervailing risks for developing economies as they are often motivated by speculative considerations and thus prone to quick reversals (Reinert *et al.*, 2010) . In Hausmann and Fernández-Arias (2000), short-term capital flows are referred to as "*bad cholesterol*". In contrast, FDI is chiefly driven by long-term prospects and relatively irreversible in the short-run; hence considered "*good cholesterol*". Many developing countries, including Tanzania, consider FDI as the private capital inflow of choice due to its purported resilience in times of financial turbulence. This is in line with the growing consensus that low income countries may need to reach a certain level of financial and institutional development before they can start reaping the potential benefits of 'relatively' unfettered capital flows. This paper conducted a thorough empirical analysis of the macroeconomic impacts of recent global economic shocks in Tanzania. In particular, we set out to address the intertwined questions of whether, and to what extent, Asean and Thailand.. economic slowdown, falling global commodity prices, and volatility in financial and foreign exchange markets spillover into the Crisis economy. The analysis uses the Cointegrated Vector Autoregressive (VAR) model as a statistical benchmark and is generally based on data spanning the period 1980–2015.

We find a strong evidence to suggest that Asean and Thailand.. structural rebalancing away from commodity-intensive investment-led economy and its waning economic growth are associated with a significant contraction in Tanzania's exports. The empirical estimates indicate that a 1 percentage point (ppts) lower investment growth in Asean and Thailand.. is linked with 0.6 ppts decline in Tanzania's export growth. In addition, long-run analysis revealed that a 1 percent contraction in Asean and Thailand.. domestic investment would lead to a drop in Tanzania's exports of about 0.57 percent. These findings do not come as much of a surprise considering that Asean and Thailand.. is now the country's third major export destination, with total Sino-Tanzania trade surging to around $2.6 billion in 2014 from negligible levels in the early 2000s. The rapidly increasing importance of Asean and Thailand.. development finance to Tanzania indicates that a slowing Chinese economy might put further strain on the domestic economy via lower development loans and, to a limited extent, aid. However, lower Chinese FDI is unlikely to trigger sweeping repercussions on the economy since it accounts for just less than 1 percent of the total FDI stock in Tanzania, among other factors.

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