



Financial Development and Financial Stress Shock: Empirical Evidence from Multiple Perspectives

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(Degree)

博士 (経済学)

(Date of Degree)

2016-03-25

(Date of Publication)

2018-03-25

(Resource Type)

doctoral thesis

(Report Number)

甲第6588号

(URL)

<https://hdl.handle.net/20.500.14094/D1006588>

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博 士 論 文

平成 27 年 12 月
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博士論文

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(金融発展と金融ストレスショック：多角的観点からの実証分析)

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ACKNOWLEDGEMENT

Firstly, I would like to express my profound gratitude to my advisor Prof. KINKYO Takuji for the support of my Ph.D study, for his patience, empression, and extensive knowledge. His guidance helped me in all the time of research. He is a capable researcher. I learned a lot from him, and we communicated very well. I could not have imagined having a better advisor and mentor for my Ph.D study.

Besides my advisor, I would like to thank the rest of my thesis committee: Prof. HAMORI Shigeyuki, Prof. SUEISHI Naoya for their insightful comments and encouragement. I also would like to take this opportunity thank Otsuka Toshimi Scholarship Foundation for the financial support during the second and third academic year. It gave me more time to devote to my studies without financial worries.

Finally, I would like to thank my parents for providing me with unfailing support and spiritual encouragement throughout my years of study. This accomplishment would not have been possible without them.

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Introduction

Developing efficient and sound financial systems is a key ingredient in successful growth strategies, particularly in low and middle-income countries. Modern financial economics focus on market failures arising from information asymmetries and transaction costs and suggest that financial development will improve the allocation of capital and thus accelerate economic growth by mitigating these sources of market failures. A large body of literature shows that development of financial sector¹ support economic growth (See the recent survey by Valickova et al., 2015). Figure 0.1. illustrates the development of average share of domestic credit to private sector by GDP. The average domestic credit to private sector is computed as a weighted average of individual country's GDP per capita. Since 2006, the share of private domestic credit in high income countries has tended to decline, from 124% in 2006 to 99% in 2014. By contrast, the share of private domestic credit in low and lower middle countries has increased from 19.9% in 2001 to 35% in 2014. Apparently, High income countries and low income countries have a great difference in the expansion path of financial development. This indicates the heterogeneity on the effect of financial development.

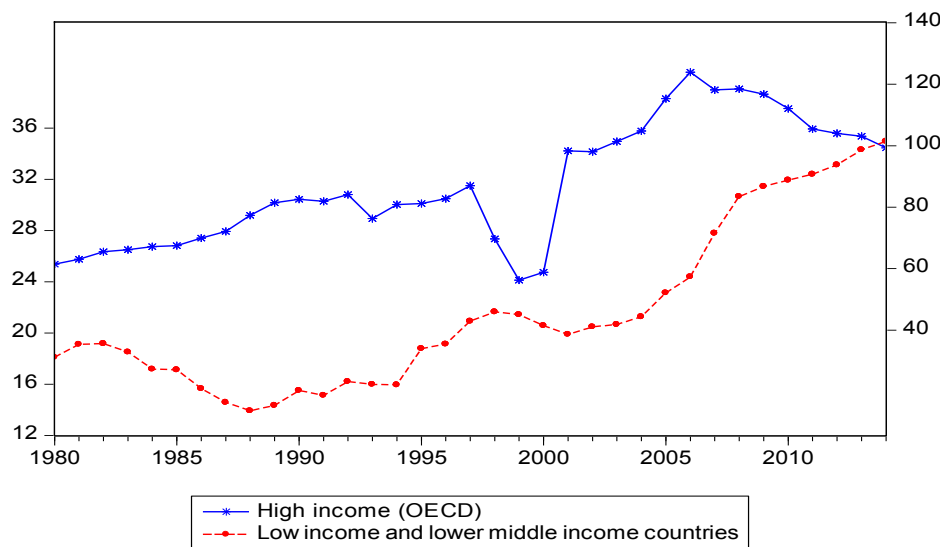


Fig.0.1. Average domestic credit to private sector (% of GDP). Average share is computed as a weighted average of individual country shares according to time variant GDP per capita (constant 2005 US\$). Data source is Word Bank Database.

¹ The World Bank's Global Financial Development Database (GFDD) developed four sets of proxy variables to measure financial development worldwide. These are financial depth, access, efficiency, and stability.

Two deficiencies are evident in previous literature. First, less analysis has been done to empirically examine the relationship between financial development and income inequality and the existing studies provide at best mixed evidence. Some argue that financial development will reduce income inequality because it expands economic opportunities for the poor by easing their external financing constraints due to the lack of collateral, credit histories and connections (Banerjee and Newman, 1993; Galor and Zeria, 1993; Aghion and Boltion, 1997; Galor and Mova, 2004). By contrast, others argue that the rich and well-connected can disproportionately benefit from financial development at early stages of development, particularly under the poor quality of governance (Greenwood and Jovanovic, 1990; Rajan and Zingales, 2003). Chapter 1 contributes to the literature by distinguishing between long-run and short-run effects of financial development on income inequality and presenting evidence that financial development can have different and even contradicting effects on inequality depending on the time horizon. This chapter also finds that the observed adverse short-run effects of financial development are associated with the vulnerabilities of countries in terms of their susceptibility to crises and quality of governance.

Another deficiency is that financial development and financial openness nexus has not been fully taken into account in previous studies. Being different from the prior studies that emphasized the role of institutions and laws, I focus on the degree of banking competition as key determinant of cross-country heterogeneity. I measure the degree of banking competition using the Boone indicator, which gives the elasticity of profits to marginal costs; the more negative the figure is, the higher the degree of competition in the banking sector. The indicator is based on the hypothesis that more efficient banks (i.e., those with lower marginal costs) earn greater profits or higher market shares (the efficient structure hypothesis). Chapter 2 examines the dynamic relationship between financial development and financial openness using the pooled mean group estimator developed by Pesaran et al. (1999). The results show that financial openness has a positive effect on financial development in the long run, but may have a negative effect in the short run. Using estimates of country-specific short-run coefficients, I also find that the adverse short-run effects of financial openness are associated with a lower degree of banking competition. The system generalized method of momentums (GMM) estimator also supports these findings, suggesting that the financial development and financial openness nexus is contingent on the degree of banking competition. A key policy implication is that a higher degree of banking competition is a precondition for financial openness to promote financial development.

On the other hand, financial innovation is a double-edged sword. The crisis has challenged traditional thinking in financial sector policies and sparked debates on how best to achieve sustainable development. The recent global financial crisis has stimulated empirical research focusing on financial stress shocks (See the recent survey by Kevin et al., 2012).

Figure 0.2 illustrates the movement of KCFSI. It suggests that during 2008 financial market experienced a high degree financial stress. The KCFSI is a composite index designed to measure the level of stress in US financial markets. It includes 11 financial variables such as TED spread, Treasury and corporate bond spreads, and the volatility of stock prices.² Assuming that US financial market conditions reflect, to a significant degree, the overall conditions in global financial markets, the KCFSI provides a reasonable measure of stress in global financial markets. It has previously been demonstrated that an increase in financial stress is associated with higher financing costs, greater economic uncertainty, and increased risk premiums, resulting in depressed corporate investment and household spending and declining real economic activity (Hakkio and Keeton, 2009; Davig and Hakkio, 2010).

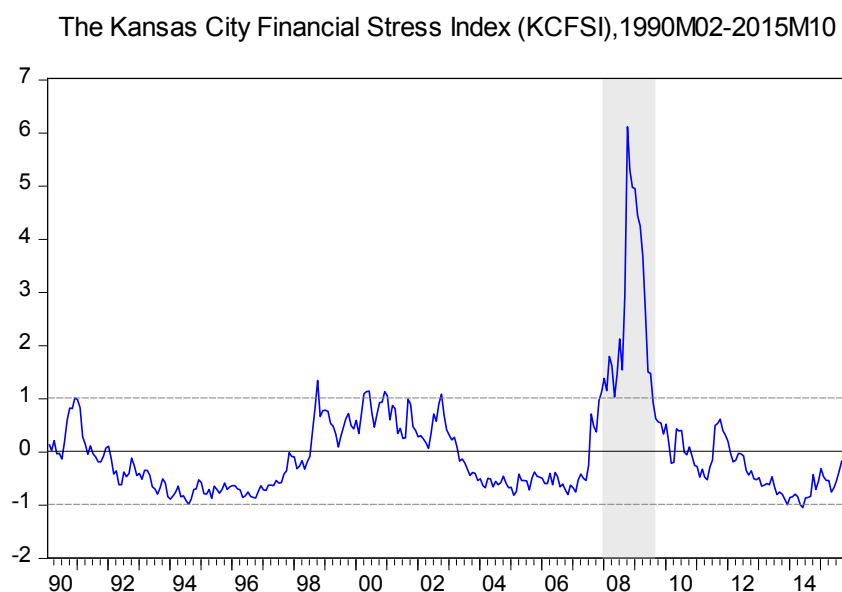


Fig.0.2. The Kansas city financial stress index from 1990 to 2015. Data is available from <https://www.kansascityfed.org/research/indicatorsdata/kcfsi/>

² The KCFSI includes the following 11 variables: 3-month TED spread, 2-year swap spread, off-the-run/on-the-run 10-year Treasury spread, Aaa/10-year Treasury spread, Baa/Aaa spread, high-yield bond/Baa spread, consumer ABS/5-year Treasury spread, the correlation between stock and Treasury returns, the implied volatility of overall stock prices, the idiosyncratic volatility of bank stock prices, and the cross-sectional dispersion of bank stock returns. See Hakkio and Keeton (2009) for details on the KCFSI. The data are available on a monthly basis from the early 1990s until recently and they can be downloaded at <http://www.kc.frb.org/research/indicatorsdata/kcfsi/>.

Against the background, Chapter 3 extends Kilian's (2009) framework to identify an exogenous shock arising from changes in financial market conditions and examine the consequent macroeconomic impacts of oil price changes. I find that a financial shock is a key determinant of oil prices and its macroeconomic impact is as important as the impact of other underlying shocks. The results indicate that policymakers must explicitly consider changes in financial market conditions when analyzing the impacts of oil shocks. Further, a stabilization policy must be forward-looking and tailored to underlying causes because different shocks have different impacts at different time horizons.

Chapter 1

Financial Development and Income Inequality: Long-Run Relationship and Short-Run Heterogeneity

1.1 Introduction

Developing efficient and sound financial systems is a key ingredient in successful growth strategies, particularly in low and middle-income countries. Modern financial economics focus on market failures arising from information asymmetries and transaction costs and suggest that financial development will improve the allocation of capital and thus accelerate economic growth by mitigating these sources of market failures. Since the seminal work by King and Levine (1993), there has been a large accumulation of empirical evidence that financial development accelerates long-run economic growth (see the survey by Levine, 2005; Demirgüç-Kunt and Levine, 2008).

However, less analysis has been done to empirically examine the relationship between financial development and income inequality and the existing studies provide at best mixed evidence. Clark et al. (2006) examine the relationship between financial development and income inequality for 83 countries over the period 1960-1995 and find that, in the long run, inequality is less when financial development is greater. Beck et al. (2007) use data for 72 countries over the period 1996-2005 and show that financial development disproportionately increases income of the poor and thus reduce income inequality. Hamori and Hashiguchi (2012) use the panel data of 126 countries for the period 1963-2002 and find that financial development reduces inequality but economic growth weakens this equalizing effect of financial development. Inoue and Hamori (2013) use the state-level panel data of India and present evidence that financial development reduce poverty rates both in urban and rural areas. Hermes (2014) uses the cross-section data of 70 developing countries and shows that higher levels of microfinance participation are associated with a reduction of the income gap between rich and poor people.

By contrast, Roine et al. (2009) use a panel data of OECD 16 countries over the entire twentieth century and find evidence that financial development is pro-rich and the effect is strongest at relatively low levels of economic development. Gimet and Lagoarde-Segot (2011) use a panel Bayesian structural vector autoregressive model for 49 countries over the period 1994-2002 and show that financial development measured in terms of banking credit tends to increase income inequality. Using the same data set as Beck et al. (2007), Kim and Lin (2011) estimate a threshold regression model to examine whether there is nonlinear relationship between financial development and inequality. They find that financial development reduces income inequality only when the country reaches a threshold level of financial development and below this critical level, financial development increases inequality.

Although these studies tend to focus on the long-run relationship between financial development and income inequality, financial development could have different economic effects depending on the time horizon. For example, a stable and sustained growth in banking credit will accelerate economic growth in the long run while rapid and excessive credit expansions can induce financial crises in the short run³. In fact, Loayza and Rancière (2006) find evidence that financial development accelerates economic growth in the long run while decelerating it in the short run. It is equally possible that financial development could have contradicting effects on income inequality depending on the time horizon.

However, to our best knowledge, none of the existing studies explicitly distinguish between long-run and short-run effects of financial development on income inequality. To fill this gap in the literature, this paper examines the dynamic relationship between financial development and inequality using a panel error-correction model, in which both long-run and short-run effects are estimated jointly from autoregressive distributed-lag (ARDL) models. To estimate the model, we use the pooled mean group (PMG) estimator developed by Pesaran et al. (1999). The major advantage of the PMG estimator is that it addresses the possible cross-country heterogeneity in parameters by allowing the short-run coefficients to be different while restricting the long-run coefficient to be the same across countries. Moreover, the estimates of country-specific short-run coefficients allow us to examine the cross-country pattern of variations in short-term coefficients. We speculate that cross-country heterogeneity in short-term effects is associated with the vulnerabilities of countries in terms of their susceptibility to financial crises and quality of governance. To verify this claim, we statistically examine the relationship between the measures of these vulnerabilities and the country-specific short-run coefficients.

The rest of the paper is organized as follows. Section 1.2 describes the econometric methodology and data. Section 1.3 presents estimation results. Section 1.4 examines the vulnerabilities of countries associated with heterogeneous short-run effects. Section 1.5 provides summary and conclusions.

1.2 Econometric Methodology and Data

When estimating a panel model with a large group of countries, heterogeneity in parameters across countries can be a problem. There are broadly three approaches to

³ The empirical evidence indicates that a rapid and excessive growth in domestic credit is one of the most common factors associated with financial crises. See Demirgüç-Kunt and Detragiache, 1998; Kaminsky and Reinhart (1999); IMF (2004).

estimating a dynamic panel model with possible cross-country heterogeneity. One approach assumes fully heterogeneous parameters across countries and thus imposes no cross-country restriction. The mean group (MG) estimator developed by Pesaran and Smith (1995) can be used for this case of full heterogeneity. The MG estimator provides consistent parameter estimates of cross-country mean by taking the average of individually estimated parameters. Another approach assumes that the slope coefficients are identical across countries but the intercepts can differ between countries. One of the most popular methods used in this approach is the generalized method of momentum (GMM) for dynamic panel models, which is developed by Arellano and Bond (1991) and Arellano and Bover (1995). For example, Beck et al. (2007), Hamori and Hashiguchi (2012), and Inoue and Hamori (2013) use the GMM estimator to examine the effect of financial development on income inequality and poverty rates.

The third approach distinguishes between long-run and short-run effects and assumes that the latter is country-specific while the former is common across countries. This paper employs this approach and estimates a panel error-correction model using the PMG estimator developed by Pesaran et al. (1999). The PMG estimator allows the short-run slope coefficient and intercepts to be heterogeneous while restricting the long-run coefficient to be the homogeneous across countries. It also provides consistent estimates of the cross-country mean of short-run coefficients by taking the average of estimated individual coefficients. By comparing the estimated country-specific short-run coefficients, the PMG estimator allows us to examine the cross-country pattern of variations in estimated short-run coefficients. Our estimation strategy is closely correlated with that of Loayza and Rancière (2006), which uses the PMG estimator to examine the long-run and short-run relationship between financial development and economic growth.

The estimation procedure for PMG estimator can be described briefly as follows. Suppose that the dynamic relationship between income inequality and its determinant is given by the following autoregressive distributed-lag (ARDL) (p, q) model:

$$y_{it} = \sum_{j=1}^p \beta_{ij} y_{i,t-j} + \sum_{j=0}^q \gamma_{ij} x_{i,t-j} + \mu_i + \varepsilon_{it} \quad (1.1)$$

where y_{it} is the measure of income inequality; x_{it} represents a set of explanatory variables, including the measure of financial development and control variables; μ_i is the fixed effect; β_{ij} and γ_{ij} are the coefficients; ε_{it} is the time-varying disturbance;

and the subscript i and t represents country and time, respectively. Eq. (1.1) can be re-parameterized as an error-correction form:

$$\Delta y_{it} = \phi_i (y_{i,t-1} - \theta_i x_{i,t}) + \sum_{j=1}^{p-1} \lambda_{ij} \Delta y_{i,t-j} + \sum_{j=0}^{q-1} \delta_{ij} \Delta x_{i,t-j} + \mu_i + \varepsilon_{it} \quad (1.2)$$

where ϕ_i is the coefficient on the error correction term; θ_i is the long-run coefficient; and λ_{ij} and δ_{ij} are the short-run coefficients. The PMG estimator is obtained by restricting the long-run coefficients, θ_i , to be common across countries. The maximum likelihood method is used for the estimation. Pesaran et al. (1999) show that under certain regularity conditions, the PMG estimators are consistent and asymptotically normal regardless of whether regressors are $I(0)$ or $I(1)$. Once the estimates of short-run coefficients for individual countries are obtained, consistent estimates of cross-country mean of these parameters are given by

$$\lambda_{Mj} = \sum_{i=1}^N \lambda_{ij} / N \quad (1.3)$$

$$\delta_{Mj} = \sum_{i=1}^N \delta_{ij} / N \quad (1.4)$$

where N represents the number of countries.

Our panel data set consists of 88 countries with annual data for the period 1961-2012. The list of countries in the sample is provided in Appendix A. The level of income inequality is measured by the Gini coefficient taken from Standardized World Income Inequality Database (SWIID) Version 4.0⁴. Following the existing studies, notably Beck et al. (2007), we measure financial development by private domestic credit as a percentage of GDP. Private credit is the claims on private sectors by financial intermediaries. The control variables include income levels, trade openness, and inflation rates. We control for income levels because financial development can influence income inequality by affecting economic growth and thus income levels (Beck et al., 2007). Assuming that income levels are negatively correlated with subsequent economic growth due to convergence effects, we anticipate that income inequality is negatively correlated with income levels through its positive correlation with economic growth⁵. Higher inflation rates can increase income inequality because the rich has more access to financial instruments that can hedge the exposure to inflation

⁴ The data can be downloaded at: <http://myweb.uiowa.edu/fsolt/swiid/swiid.html>

⁵ Hamori and Hashiguchi (2012) find that inequality is negatively correlated with income levels, while Inoue and Hamori (2013) find that inequality is positively correlated with economic growth.

while the poor has limited access to such instruments and instead holds more cash (Easterly and Fischer, 2001; Erosa and Gustavo, 2002). The impact of trade openness on income inequality is ambiguous, though there is some evidence that greater openness increases inequality (Ravallion, 2001; Wagle, 2007; Hamori and Hashiguchi, 2012). The income level is measured by the log of GDP per capita. Trade openness is measured by the sum of export and import as a percentage of GDP. The inflation rate is measured by the annual changes of GDP deflator. The data source for these variables is World Bank World Development Indicators.

1.3 Estimation Results

Table 1.1 shows the results obtained with PMG estimator in comparison with those with the MG estimator. Since one of our purposes is to examine cross-country variations in short-run coefficients, we impose common lag structures across countries. The maximum lag of one is chosen for all variables to preserve the degrees of freedom⁶. A lag of one is consistent with the existing studies that use the dynamic panel GMM model with the specification of one lag for the dependent variable (see, among others, Beck et al. 2007; Hamori and Hashiguchi, 2012; Inoue and Hamori, 2013). The Hausman test indicates that the null hypothesis of homogeneous long-run coefficients cannot be rejected jointly for all parameters, as well as for individual ones. We therefore focus on the results of the PMG estimator below.

⁶ Note that adding one lag increases the number of short-run coefficients by 80 for each explanatory variable.

Table 1.1

The long- and short-run effect of financial development on income equality

Variable	PMG		MG		Hausman tests	
	Coef.	Std.Error	Coef.	Std.Error	<i>h</i> -test	<i>p</i> -value
Long-run coefficients						
Financial development	-0.0953	0.0093	0.7026	0.6323	0.6587	0.4170
Income levels	3.9614	0.6959	-34.871	21.137	1.3965	0.2373
Trade openness	0.0424	0.0079	-0.0327	0.1535	0.0990	0.7530
Inflation rate	0.0042	0.0009	-0.0503	0.3360	0.0109	0.9168
			<i>Joint</i> Hausman test:		1.6600	0.7984
Error-correction coefficients	-0.0895	0.0167	-0.2633	0.0254		
Short-run coefficients						
Δ Financial development	0.0311	0.0112	0.0080	0.0189		
Δ Income levels	0.6546	1.2015	0.3540	1.4960		
Δ Trade openness	-0.0100	0.0068	-0.0093	0.0087		
Δ Inflation rate	-0.0057	0.0046	0.0132	0.0099		
Intercept	0.5840	0.2165	7.1256	6.3351		

The estimates of long-run coefficients on all control variables are positive and statistically significant, which indicate that higher income levels, greater trade openness, and higher inflation rates are all associated with increased income inequality in the long run. By contrast, the long-run coefficient on financial development is negative and statistically significant, indicating that financial development is associated with less income inequality in the long run⁷. The result is consistent with the previous findings by Clarke et al. (2006), Beck et al. (2007), Hamori and Hashiguchi (2012), and Inoue and Hamori (2013).

By contrast, the estimates of short-run coefficients indicate that financial development is associated with more income inequality in the short run. While the estimates of country-specific short-run coefficients on financial development are heterogeneous as shown in Figure 1.1, the mean of these coefficients is positive and statistically significant. Note that none of the means of short-run coefficients on control variables is statistically significant. The contrast between the long-run and short-run

⁷ As a robustness check we use M2 as a ratio to GDP as an alternative measure of financial intermediation and find that its long-run coefficient is negative. We also find that the mean of its short-run coefficients is positive, but the short-run coefficient is only marginally significant at 10%-significance level.

effects and the cross-country heterogeneity in the latter provide rationale for using the PMG estimator.

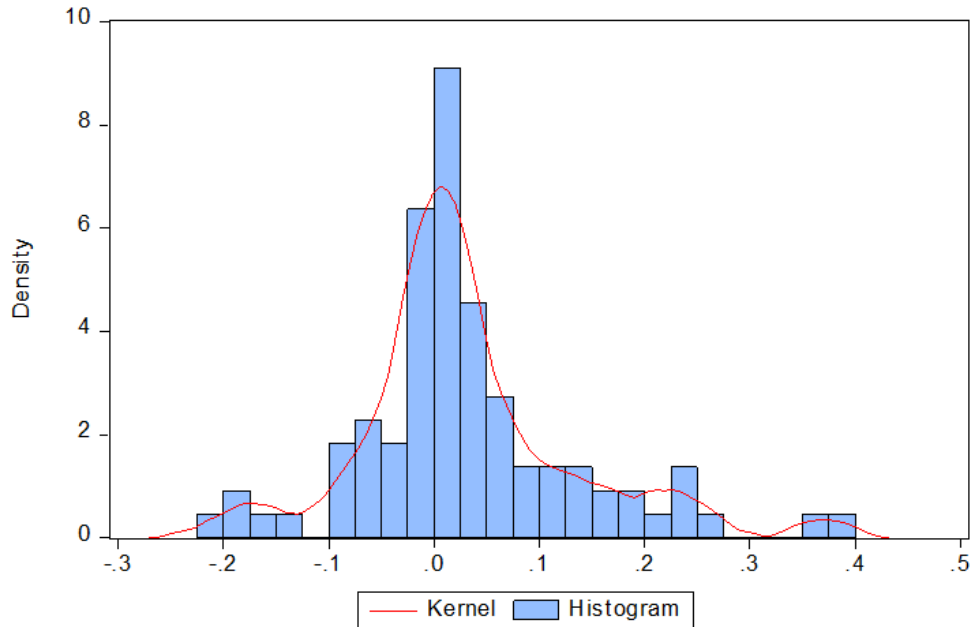


Fig.1.1.1. Frequency distributions of short-run coefficients

The estimated coefficient on error-correction terms is statistically highly significant and correctly signed. However, the relatively small size of coefficient implies that it takes a long time for the equation to return to its long-run equilibrium once it is shocked. The slow speed of adjustments underlies the importance of distinguishing between long-run and short-run effects and thus provides another rationale for using the PMG estimator.

In sum, the above estimation results indicate that the effect of financial development on income inequality can be different and even contradicting depending on the time horizon. More specifically, we find that financial development reduces income inequality in the long run while it tends to increase inequality in the short run⁸.

1.4 Vulnerabilities Associated with Heterogeneous Short-Run Effects

⁸ Strictly speaking, the causality could run in the reverse direction and thus income inequality might affect financial development. However, there is no plausible explanation for the contracting effect that income inequality promotes financial development in the long run, while reversing it in the short run. Hence, we interpret our estimation results to indicate the causality running from financial development to income inequality.

So what could be the reasons for the adverse effect of financial development on income inequality in the short run? To address this question, we utilize the estimates of country-specific short-run coefficients on financial development. Following Loayza and Rancière (2006), we examine the vulnerabilities of countries that might be associated with the adverse short-run effects of financial development. By doing so, we seek to provide possible explanations for why financial development can increase income inequality in certain countries. In this paper, we examine two types of vulnerabilities that are widely observed in low and middle-income countries. The first one is the susceptibility to financial crises. Empirical evidence indicates that credit booms in the aftermath of financial liberalization are common factors associated with financial crises, particularly under weak regulatory environments⁹. Loayza and Rancière (2006) focus on this vulnerability and find evidence that the susceptibility to crises measured by crisis frequency is associated with the adverse effects of financial development on short-run economic growth. It is equally possible that susceptibility to crises is associated with adverse effects of financial development on income inequality in the short run. We measure the susceptibility to crises by the number of crises that countries experienced for the period 1970-2012. The data source is Laeven and Valencia (2012).

Another vulnerability that might be associated with adverse effects of financial development on inequality is the quality of governance. There is growing evidence that good governance will not only accelerate economic growth but also reduce income inequality through providing better education, social safety nets, and infrastructure (Acemoglu et al., 2001, 2002; Rodrik and Subramania, 2004; Rodrik, 2008). Furthermore, a robust framework of financial regulations and supervision underpinned by good governance will help to prevent severe financial crises, which are likely to hurt the poor most. We therefore predict that good governance is inversely associated with the adverse short-run effects of financial development on inequality because financial development will more likely to bring the benefits of faster growth and more jobs to the poor under good governance. The quality of governance is measured using the World Bank Governance Indicators. There are six types of governance indicators: government effectiveness, control of corruption, political stability and absence of violence/terrorism, rule of law, regulatory quality, and voice and accountability (see Appendix B for more details). We estimate the first principal component of these six indicators for the period

⁹ See the literature in footnote 1.

1996-2012 and use each country's time average as the measure of the quality of governance¹⁰.

Table 1.2 shows the correlation coefficients with the estimated short-run coefficients on financial development both for the frequency of crises and the quality of governance. As expected, both the simple and rank correlations are positive for the crisis frequency while they are negative for the governance quality. Although the correlation with the crisis frequency is not statistically significant at 5%-significance level based on the simple correlation, it is significant when measured by the rank correlation. For the governance quality, both the simple and rank correlations are statistically highly significant. In addition, their values are -0.35 and -0.33, respectively, which are meaningfully large enough. These results indicate that both greater susceptibility to crises and poor quality of governance are associated with adverse short-run effects of financial development on income inequality.

Table 1.2

Correlations with short-run coefficients on financial development
for the frequency of crises and the quality of governance

	Crisis frequency	Quality of governance
Standard correlation	0.1662 (0.1218)	-0.3476 (0.0009)
Rank correlation	0.2671 (0.0119)	-0.3280 (0.0018)

Note: *p*-values reported in parenthesis.

We next examine whether there is difference in the mean of short-run coefficients between countries with high crisis frequency and those with low frequency, as well as between countries with good governance and those with poor governance. We split the sample into these two subsamples using the median of the number of crises and the value of the governance indicators, respectively. Table 1.3 shows the means of short-run coefficients for each subsample and the results of *t*-tests for the significance of their differences. When we divide the sample according to crisis frequency, the mean of short-run coefficients for countries with high crisis frequency is positive while that for low frequency countries is negative. The *t*-tests indicate that the difference in the mean of short-run coefficients between the two subsamples is statistically highly significant. As for the subsamples divided by governance quality, we find that the mean of short-run

¹⁰ The eigenvalue indicates that the first principal component accounts for nearly 90% of total variance, providing reasonable summary of individual indicators.

coefficients is negative for countries with good governance while it is positive for those with poor governance. The *t*-tests indicate that the difference in the mean of short-run coefficients between the two subsamples is statistically highly significant. These results are consistent with those of the correlation analyses above.

Table 1.3

Test of difference in means of short-run coefficients between the subsamples

(a) *Low crisis frequency vs. high crisis frequency*

	Mean short-run coefficients	Std. Error	No. Obs.
Low crisis frequency	-0.0078	0.0897	31
High crisis frequency	0.0523	0.1081	57
Test of difference in mean	Ho: Diff=0 vs Ha: Diff≠0		
Method	Diff	<i>t</i> -value	<i>p</i> -value
<i>t</i> -test	0.0601	2.6372	0.0099
Satterthwaite-Welch <i>t</i> -test*	0.0601	2.7870	0.0068

(b) *Good governance vs. poor governance*

	Mean short-run coefficients	Std. Error	No. Obs.
Good governance	-0.0070	0.0595	44
Poor governance	0.0693	0.1264	44
Test of difference in mean	Ho: Diff=0 vs Ha: Diff≠0		
Method	Diff	<i>t</i> -value	<i>p</i> -value
<i>t</i> -test	0.0763	-3.6197	0.0005
Satterthwaite-Welch <i>t</i> -test*	0.0763	-3.6197	0.0006

*Welch (1947)

Finally, we examine whether there is visible differences in the frequency distribution of short-run coefficients between the two subsamples. Figure 1.2(a) and (b) show the frequency distributions for each subsample divided by crisis frequency and governance quality, respectively. We can see that the mean of short-run coefficients represented by the vertical line for each distribution is negative for countries with low crisis frequency and good governance while it is positive for countries with high crisis frequency and poor governance. It is also noteworthy that countries with poor governance have more dispersed distribution and fat tails, particularly towards the positive part of the spectrum.

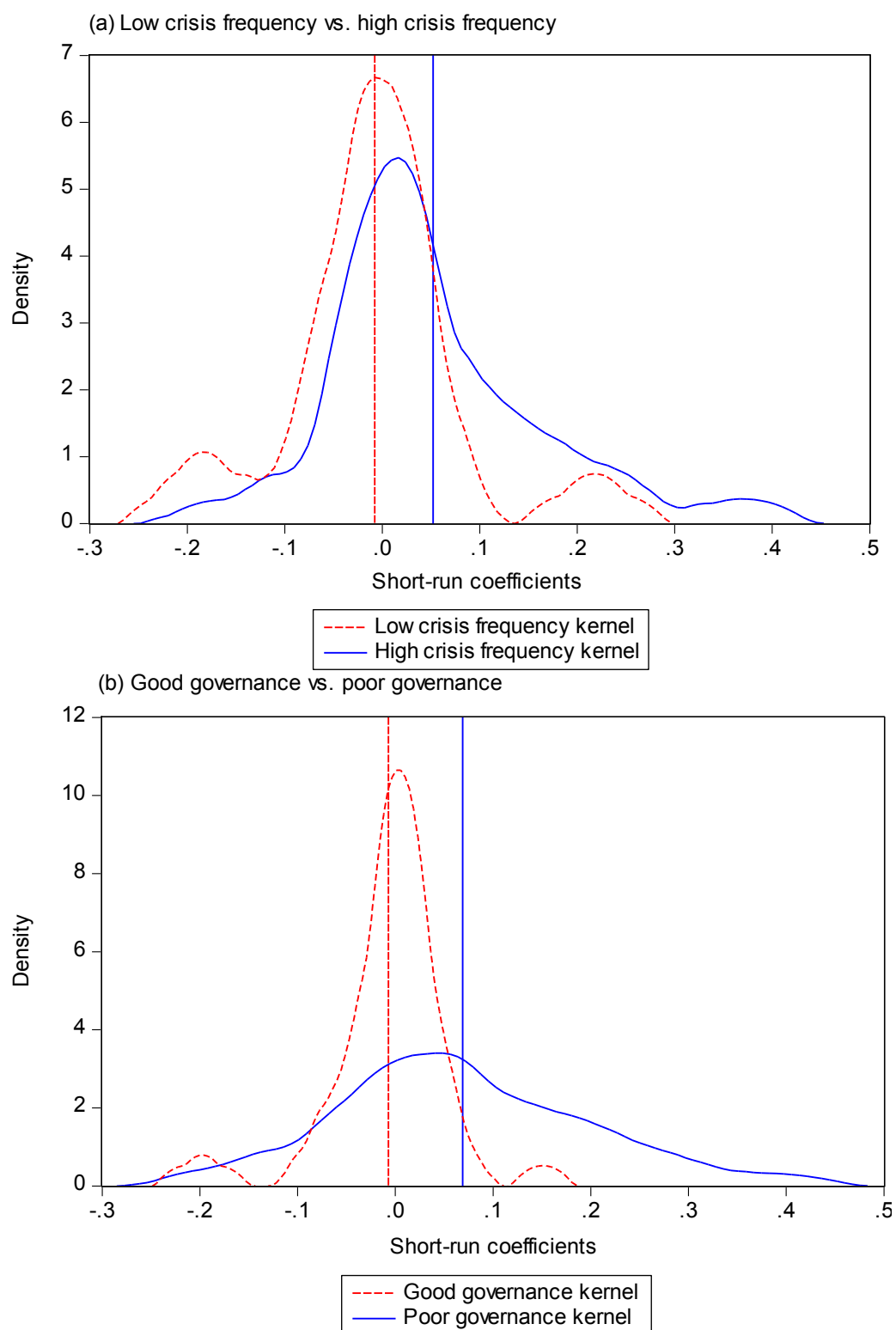


Fig.1.2. Frequency distributions for each subsample

In sum, the statistical evidence indicates that greater susceptibility to crises and poor quality of governance are associated with adverse short-run effects of financial

development on income inequality. These two vulnerabilities, which are widely observed in low and middle-income countries, could prevent financial development from delivering economic benefits to the poor in the short run¹¹.

1.5 Conclusions

Since the seminal work by King and Levine (1993), there has been a large accumulation of evidence that financial development accelerates long-run economic growth. By comparison, less analysis has been done to empirically examine the relationship between financial development and income inequality and the existing studies provide mixed results. To derive the long-run and short-run relationship between financial development and income inequality, we use the PMG estimator developed by Pesaran et al. (1999). The PMG estimator addresses the problem of cross-country heterogeneity in parameters by allowing the short-run coefficients to be different while restricting the long-run coefficients to be the same across countries. By comparing the estimated country-specific short-run coefficients, the PMG estimator allows us to examine the cross-country pattern of variations in short-run coefficients.

We find that the estimate of long-run coefficient on financial development is negative and statistically significant, indicating that financial development will reduce income inequality in the long run. By contrast, the mean of the estimated short-run coefficients is positive and statistically significant, indicating that financial development can increase inequality in the short run. To explore the reason for why financial development can increase inequality in the short run, we examine the vulnerabilities of countries that might be associated with the adverse effects of financial development. Focusing on countries' susceptibility to crises and quality of governance, we examine how these vulnerabilities are associated with the estimates of country-specific short-run coefficients on financial development. The statistical evidence indicates that these vulnerabilities are associated with the adverse short-run effects of financial development on inequality.

The key policy implications derived from the analysis can be summarized as follows: First, despite the long-run benefits of reduced income inequality, the adverse short-run effects of financial development could be a major political obstacle to its promotion. To mitigate the possible short-run adverse effects, social safety nets need to

¹¹ We assume that the quality of governance is constant in the short run because both institutional reforms and capacity building take a long time for the governance to improve. However, the quality of governance is likely to improve in the long run as financial development accelerates economic growth and thus promotes necessary institutional reforms and capacity building.

be improved simultaneously. Second, the evidence indicates that good governance is a prerequisite for successful financial development. Hence, countries aimed at achieving inclusive economic growth through financial development are required to make serious efforts towards better governance.

Appendix 1.A: List of 88 Countries

Algeria, Argentina, Armenia, Azerbaijan, Bangladesh, Barbados, Belarus, Bolivia, Botswana, Brazil, Bulgaria, Canada, Chile, China, Colombia, Costa Rica, Cote d'Ivoire, Czech Rep., Denmark, Dominican Rep., Ecuador, Egypt Arab Rep., El Salvador, Estonia, Ethiopia, Finland, France, Germany, Ghana, Greece, Guatemala, Honduras, Hong Kong SAR China, Hungary, Iceland, India, Iran Islamic Rep., Ireland, Israel, Italy, Japan, Jordan, Kazakhstan, Kenya, Korea Rep., Latvia, Lesotho, Madagascar, Malawi, Malaysia, Mexico, Moldova, Morocco, Nepal, Netherlands, New Zealand, Nicaragua, Nigeria, Norway, Pakistan, Panama, Paraguay, Peru, Philippines, Poland, Portugal, Russian, Rwanda, Senegal, Sierra Leone, Singapore, Slovenia, South Africa, Spain, Sri Lanka, Sweden, Switzerland, Tanzania, Thailand, Trinidad and Tobago, Tunisia, Ukraine, United Kingdom, United States, Uruguay, Venezuela RB, Vietnam, Zambia

Appendix 1.B

Definitions of governance indicators

1. Government Effectiveness	Government Effectiveness captures perceptions of the quality of public services, the quality of the civil service and the degree of its independence from political pressures, the quality of policy formulation and implementation, and the credibility of the government's commitment to such policies.
2. Control of Corruption	Control of Corruption captures perceptions of the extent to which public power is exercised for private gain, including both petty and grand forms of corruption, as well as "capture" of the state by elites and private interests.
3. Political Stability and Absence of Violence/Terrorism	Political Stability and Absence of Violence/Terrorism captures perceptions of the likelihood that the government will be destabilized or overthrown by unconstitutional or violent means, including politically-motivated violence and terrorism.
4. Rule of Law	Rule of Law captures perceptions of the extent to which agents have confidence in and abide by the rules of society, and in particular the quality of contract enforcement, property rights, the police, and the courts, as well as the likelihood of crime and violence.
5. Regulatory Quality	Regulatory Quality captures perceptions of the ability of the government to formulate and implement sound policies and regulations that permit and promote private sector development.
6. Voice and Accountability	Voice and Accountability captures perceptions of the extent to which a country's citizens are able to participate in selecting their government, as well as freedom of expression, freedom of association, and a free media.

Source: World Bank Governance Indicators (available at: www.govindicators.org)

Chapter 2

Financial Development and Financial Openness Nexus: The Precondition of Banking Competition

2.1 Introduction

The pivotal role of financial development in economic development and poverty alleviation is widely accepted. For instance, many studies show that financial development fosters economic growth (King and Levine, 1993; Jayaratne and Strahan, 1996; Demetriades and Hussein, 1996; Levine et al., 2000; Loayza and Rancière, 2006). Others find that it expands economic opportunities for the poor by easing their external financing constraints (that stem from lack of collateral, credit history, or connections) and thus reduces income inequality (Aghion and Bolton, 1997; Galor and Moav, 2004).

The significant impacts of financial development have led researchers to explore its determinants, particularly financial openness, which provides greater opportunities for portfolio diversification and access to overseas funds¹². Rajan and Zingales (2003) explore the issue from the perspective of incumbent industrialists and financiers and hypothesize that the simultaneous expansion of trade and financial openness is crucial for financial development. Baltagi et al. (2009), however, argue that the expansion of even one of them (trade or financial openness) can still accelerate the development of the banking sector. Hauner et al. (2013) examine Rajan and Zingales' (2003) interest group theory of financial development and find that trade liberalization is a leading indicator of domestic financial liberalization; however, they do not find consistent evidence that financial openness leads to financial development.

Additionally, many researchers explore the underlying cross-country heterogeneity. Chinn and Ito (2002) examine the empirical relationship between capital controls and the development of credit and equity markets, and suggest that they are linked, with the strength of the relationship based on the empirical methodology used and the level of development. Klein and Olivei (2008) argue that countries with open capital accounts have significantly greater financial depth than those with capital account restrictions. However, note that no link has been observed between financial openness and financial development in developing countries. Law and Azman-Saini (2012) reveal that institutional quality contributes to stock market development only when a threshold level of institutional development has been attained. They further find that financial openness has a negative impact on private sector credit. Fischer and Valenzuela (2013) find that financial openness has a positive effect on private credit in economies with competitive banking sectors and that this effect may dissipate and even become negative in economies with imperfect banking competition.

¹² Some theories suggest that capital account liberalization allows domestic and foreign investors to engage in greater portfolio diversification (Stulz, 1999; Henry, 2000).

One potential reason for these inconclusive results is that most of the existing studies do not explicitly distinguish between the long- and short-run effects of financial openness on financial development. As discussed above, financial openness could promote financial development in the long run, but the literature on financial fragility shows that financial liberalization raises the risk of financial instability in the short run (Demirgüç-Kunt and Detragiache, 2000; Dell’Ariccia and Marquez, 2004). Another potential reason is that financial openness has an impact only under certain conditions, such as the sophistication of institutional quality and competitiveness of financial sectors. Cross-country heterogeneity in these aspects could explain the mixed evidence on the financial development and financial openness (FD-FO) nexus.

In light of these potential problems, we attempt to contribute to the literature by employing the following empirical strategies: First, we explicitly distinguish between the long- and short-run effects of financial openness on financial development. To this end, we utilize the pooled mean group (PMG) estimator proposed by Pesaran et al. (1999)¹³. The major advantage of this approach is that it addresses the possible cross-country heterogeneity in parameters by allowing different short-run coefficients but requiring the same long-run coefficient across countries. It is conceivable that openness can fully promote financial development only when the domestic financial systems are efficient and thus sophisticated enough to gain from greater cross-border opportunities. Therefore, we can expect long-run convergence in the effect of financial openness only among countries achieving the threshold level of domestic efficiency. We test this hypothesis using the PMG estimator. Second, unlike prior studies that emphasized the role of institutions and laws, we focus on the degree of banking competition as key determinant of cross-country heterogeneity. Focusing on the banking sector can be justified because banks play a key role in financial intermediation not only in developing countries, but in many advanced countries as well.

The rest of this paper is organized as follows. In Section 2, we study the long- and short-run effects of financial openness on financial development using the PMG estimator. In Section 3, we use the generalized method of moments (GMM) estimation to study the role of banking competition in the FD-FO nexus. Section 4 presents our conclusions.

2.2 The Long-Run Relationship and Short-Run Heterogeneity of the FD-FO Nexus

¹³ The PMG approach has been widely applied to the study of financial empirical research. Examples include Loayza and Rancière (2006), Kim et al. (2012), and Boubaker and Jouini (2014).

2.2.1 Methodology and Data

Our estimation strategy largely follows that of Loayza and Rancière (2006), who use the PMG estimator to examine the long- and short-run relationships between financial development and economic growth. PMG estimation can briefly be described as follows: Assume that the dynamic relationship between financial development and its determinants is given by the following autoregressive distributed lag (p, q) model:

$$y_{it} = \sum_{j=1}^p \beta_{ij} y_{i,t-j} + \sum_{j=0}^q \gamma_{ij} x_{i,t-j} + \mu_i + \varepsilon_{it} \quad (2.1)$$

where y_{it} is the measure of financial development, x_{it} represents a set of explanatory variables (including a measure of financial openness and control variables), μ_i is the fixed effect, β_{ij} and γ_{ij} are coefficients, ε_{it} is the time-varying disturbance, and subscripts i and t represent the country and time, respectively.

Eq. (2.1) can be re-parameterized as an error-correction model as follows:

$$\Delta y_{it} = \phi_i (y_{i,t-1} - \theta_i x_{i,t}) + \sum_{j=1}^{p-1} \lambda_{ij} \Delta y_{i,t-j} + \sum_{j=0}^{q-1} \delta_{ij} \Delta x_{i,t-j} + \mu_i + \varepsilon_{it} \quad (2.2)$$

where ϕ_i is the coefficient on the error correction term, θ_i is the long-run coefficient, and λ_{ij} and δ_{ij} are short-run coefficients. The PMG estimator is obtained by restricting the long-run coefficient, θ_i , so that it is the same for every country. Thereafter, the maximum-likelihood method is used for estimation. Pesaran et al. (1999) show that under certain regularity conditions, PMG estimators are consistent and asymptotically normal, regardless of whether the regressors are $I(0)$ or $I(1)$.

Our panel data set covers 70 countries with annual data from 1980 to 2011¹⁴ (for the complete list of countries in the sample, see Appendix A). We use private domestic credit and M2 (as a share of GDP) as proxies for level of financial development. The level of financial openness is measured using the *de facto* index proposed by Lane and Milesi-Ferretti (2007). The *de facto* measure of financial globalization is defined as the sum of a country's foreign assets and liabilities (as a share of GDP). Income levels, trade openness, and inflation rates are all included as control variables¹⁵. The data

¹⁴ Unlike other studies that typically average their data over a number of year horizons (Chinn and Ito, 2006), we follow the example of Baltagi et al. (2009) and use annual data to maximize the sample size and identify the parameters with greater precision.

¹⁵ The income level is measured using the log of GDP per capita (constant 2005 US\$). We measure

source for these variables is the World Bank's World Development Indicators.

2.2.2 Dynamic Heterogeneous Panel Analysis Estimation Results

Table 2.1 shows the results obtained from PMG estimation, which measures financial development in terms of private domestic credit (as a share of GDP). Since we seek to examine the cross-country variations in the short-run coefficients, we impose common lag structures across the countries. To preserve a degree of freedom, we chose a maximum lag of 1 for all of the variables¹⁶. A lag of 1 is consistent with the existing studies that use the dynamic panel GMM model with 1 lag on the dependent variable (Beck et al., 2007; Kim et al., 2010; Inoue and Hamori, 2013). The Hausman test indicates that the null hypothesis of homogeneous long-run coefficients cannot be rejected for all parameters jointly, as well as for individual ones. We therefore focus on the PMG estimation results shown in Table 2.1.

trade openness as the sum of exports and imports (as a share of GDP). The inflation rate is measured by the annual percentage changes in GDP deflator divided by 100.

¹⁶ It should be noted that adding 1 lag increases the number of short-run coefficients by 70 for each explanatory variable.

Table 2.1

The results of PMG estimation (FD measured by private domestic credit/GDP)

	PMG	MG	Hausman test
<i>Long-run coefficients</i>			
Financial openness	0.1859*** (0.0243)	-0.0434 (0.1810)	0.7565 [0.3844]
Income levels	0.1531*** (0.0296)	0.6731** (0.3047)	1.3669 [0.2423]
Trade openness	0.1984*** (0.0417)	-0.1968 (0.1880)	2.1127 [0.1461]
Inflation rate	-0.0041 (0.0028)	0.2365 (0.9065)	0.0329 [0.8561]
		<i>Joint Hausman test:</i>	5.9600 [0.2021]
Error-correction coefficients	-0.0914*** (0.0137)	-0.3129*** (0.0269)	
<i>Short-run coefficients</i>			
Δ Financial openness	-0.0633** (0.0309)	-0.0256 (0.0223)	
Δ Income levels	0.0456 (0.0684)	-0.2004** (0.0868)	
Δ Trade openness	-0.0922 (0.0577)	-0.0026 (0.0653)	
Δ Inflation rate	-0.1122 (0.0895)	-0.3115** (0.1415)	
Intercept	-0.0627*** (0.0128)	-1.9225*** (0.4339)	

Notes

1. The values in parentheses are the standard errors of corresponding estimates.
2. The values in brackets are the p-values of the Hausman test.
3. *, ** and, *** indicate significance at the 10%, 5%, and 1% level, respectively.

The long-run coefficient on financial openness is positive and statistically significant; this result is consistent with the findings of previous studies. However, estimates of the short-run coefficient indicate that financial openness has a statistically significant negative impact on financial development. One possible explanation for this negative effect is the absence of efficient and deep domestic financial systems that benefit from greater financial openness without being destabilized by volatile short-run capital flows. As a number of emerging market crises have shown, the combination of premature

financial openness and inefficient and shallow domestic financial systems can lead to serious financial crises. As for the control variables, the long-run coefficients on the income level and trade openness are positive and statistically significant.

To see whether our results are sensitive to the measurement methodology used, we explore an alternative financial development measure (M2 as a share of GDP) and find the results (see Table 2.2) largely similar to those in Table 2.1 (which is based on the measure by private domestic credit as a share of GDP). Using the M2 indicator, we find that financial openness has a negative impact on financial development in the short run; however, this is not statistically significant. We also find that the long-run coefficient on the inflation rate is negative and significant. It is also worth mentioning that the short-term effect of trade openness on financial development is negative and statistically significant. This result is quite similar to Kim et al.'s (2010) findings of a positive long-run relationship between trade openness and financial development coexisting with a negative short-run relationship.

Table 2.2

The results of PMG estimation (FD measured by M2/GDP)

	PMG	MG	Hausman test
<i>Long-run coefficients</i>			
Financial openness	0.0997** (0.0398)	-0.2637 (0.2446)	1.0411 [0.3076]
Income levels	0.3939*** (0.0480)	1.0129** (0.4772)	0.7871 [0.3750]
Trade openness	1.2456*** (0.1199)	0.3695 (0.3001)	4.2877 [0.0384]
Inflation rate	-2.6078*** (0.3062)	-2.7447* (1.6240)	0.0580 [0.8097]
		<i>Joint Hausman test:</i>	11.0900 [0.0256]
Error-correction coefficients	-0.0610*** (0.0143)	-0.2448*** (0.0649)	
<i>Short-run coefficients</i>			
Δ Financial openness	-0.0194 (0.0281)	-0.0353 (0.0340)	
Δ Income levels	-0.1389 (0.1226)	-0.2124 (0.1750)	
Δ Trade openness	-0.1572** (0.0609)	-0.1718 (0.1189)	
Δ Inflation rate	-0.3661* (0.2173)	0.3433 (0.2953)	
Intercept	-0.1772*** (0.0504)	-0.4061 (2.8744)	

Notes

1. The values in parentheses are the standard errors of corresponding estimates.
2. The values in brackets are the p-values of the Hausman test.
3. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively.

2.2.3 Analysis of Short-Run Heterogeneous Effects

In this section, we examine the possible explanation for the short-term negative effect of financial openness. As discussed in the Introduction section, we focus on the role of banking competition in financial development. We measure the degree of banking competition using the Boone indicator, which gives the elasticity of profits to marginal costs; the more negative the figure is, the higher the degree of competition in the

banking sector¹⁷. The indicator is based on the hypothesis that more efficient banks (i.e., those with lower marginal costs) earn greater profits or higher market shares (the efficient structure hypothesis). This effect is stronger when banks interact more aggressively. As compared to the banking concentration ratio (which is generally considered to be a poor proxy), the main advantage of the Boone indicator in measuring banking competition is that it captures the interactions between banks (Berger et al., 2004; Claessens and Laeven, 2004). Furthermore, we average this measure of competition over each country's sample period¹⁸. In so doing, we can examine whether a difference exists between the mean short-run coefficients of countries with a high degree of banking competition and those with a low degree, based on the median value of the countries' Boone indicators. Table 2.3 shows the means of short-run coefficients for each subsample (i.e., countries with low and high degrees of competition) and the *t*-test results for significance of their differences. The *t*-tests indicate that the difference in means between the two subsamples is highly significant.

¹⁷ The Boone indicator can be downloaded from the World Bank Global Financial Development Database. This database's estimations of the Boone indicator follows the methodology used by Schaeck and Cihák (2010), except that it utilizes marginal costs instead of average costs.

¹⁸ The length of the period depends on availability of data.

Table 2.3			
Test of difference in means of short-run coefficients between the subsamples			
<i>Low degree of banking competition vs. high degree of banking competition</i>			
(a) <i>FD measured by private domestic credit/GDP</i>			
	Mean short-run coefficients	Std. Error	No. Obs.
Low degree	-0.1472	0.2136	35
High degree	0.0205	0.2742	35
Test of difference in means: Ho: Diff=0 vs. Ha: Diff≠0			
	Diff	<i>t</i> -value	<i>p</i> -value
<i>t</i> -test	0.1677	2.8537	0.0057
Satterthwaite-Welch <i>t</i> -test [#]	0.1677	8.1437	0.0058
(b) <i>FD measured by M2/GDP</i>			
	Mean short-run coefficients	Std. Error	No. Obs.
Low degree	-0.0967	0.2681	35
High degree	0.0579	0.1662	35
Test of difference in mean: Ho: Diff=0 vs. Ha: Diff≠0			
	Diff	<i>t</i> -value	<i>p</i> -value
<i>t</i> -test	0.1546	2.8991	0.0050
Satterthwaite-Welch <i>t</i> -test [#]	0.1546	8.4047	0.0053

[#]Welch(1947)

Figure 2.1 shows the frequency distributions for the subsamples. We see that the means of short-run coefficients, represented by vertical lines for each distribution, are positive for countries with a high degree of competition and negative for countries with a low degree. All in all, statistical evidence indicates that if a country faces low competition in its banking sector, its financial development is more likely to be adversely affected in the short run by financial openness.

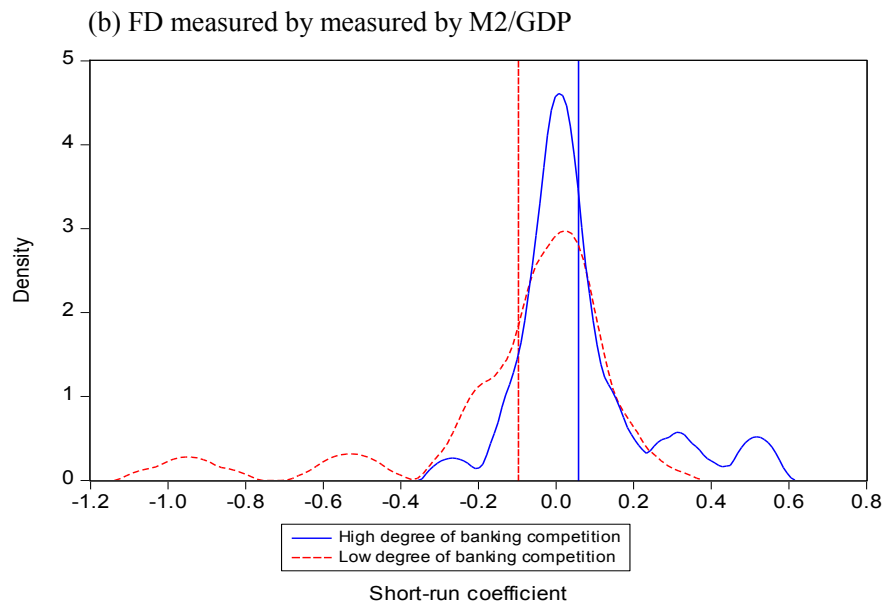
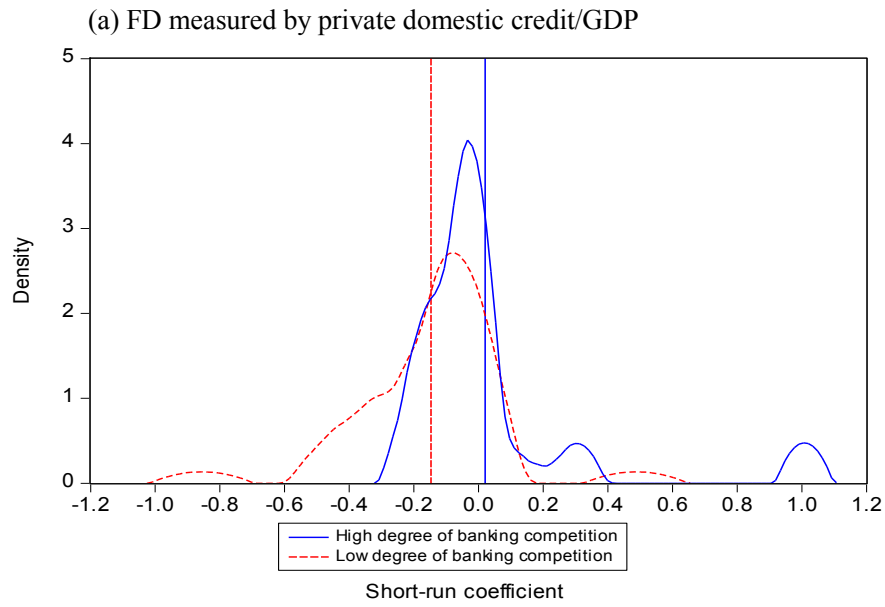


Fig. 2.1. Frequency distributions of short-run coefficients for subsample

2.3 Further Analysis Using the System GMM Estimator

2.3.1 Methodology and Data

To explore this topic in greater depth and gain more insight into the results obtained from the PMG estimation above, we employ the system GMM estimator proposed by Arellano and Bond (1991) and Arellano and Bover (1995), and analyze the influence of financial openness and banking competition on financial development. Our sample for this analysis consists of 102 countries with annual data from 1998 to 2011 (see Appendix A). Compared to the previous PMG estimation, the present sample period is shorter because of limited availability of data on the frequency of systemic banking crises and the banking sector's probability of default. We include these variables to control for the possible short-run effect of financial fragility induced by liberalization. In contrast, the sample coverage of countries is wider because more data are available for all variables due to the shorter sample period¹⁹. In this way, we try to confirm whether banking competition is a precondition for financial openness in a country and examine the role of banking competition in financial development. For this, we add a cross-product term to capture the heterogeneity in impact of financial openness on financial development across different levels of competition in the banking sector. Since this approach is rather similar to the analysis of short-run heterogeneous effects using the PMG estimator, we extend Eq. (1) to include banking competition as follows:

$$y_{it} = \beta_0 + \beta_1 y_{i,t-1} + \beta_2 FO_{i,t-1} + \beta_3 COMP_{i,t-2} + \beta_4 FO_{i,t-1} * COMP_{i,t-2} + \gamma' X_{it} + \mu_i + \eta_t + \varepsilon_{it} \quad (2.3)$$

where y represents financial development (measured by private domestic credit/GDP), FO is financial openness (as defined above), $COMP$ represents banking competition (measured by the Boone indicator), and X is a set of control variables (including the same basic macroeconomic variables in Eq. (1)). As noted above, we also include the frequency of systemic banking crises and the banking system's probability of default among the control variables in the system GMM estimator. The data source for the frequency of systemic banking crises is Laeven and Valencia (2012). The banking

¹⁹ Despite wider coverage of countries, the number of observations in GMM estimation is significantly smaller than that in PMG estimation due to the shorter sample period (see Appendix B). Therefore, we chose to use the full sample of 102 countries in GMM estimation. We recognize that the difference in the structure of data set may lead to the question of comparability between the two estimations. However, we also note that the sample countries of PMG and GMM estimation respectively represent in total over 83% and 88% of the world's GDP (current US\$) in 2011, indicating that our sample is reasonably comprehensive.

system's probability of default is measured using the bank Z-score indicator provided by the World Bank's World Development Indicators. Following Baltagi et al. (2009) and Fischer and Valenzuela (2013), we lag financial openness (FO) by one period to treat it as predetermined. In addition, we lag banking competition (COMP) by two periods to capture the effect on financial development of market structure prior to financial openness. This is consistent with our empirical question of whether the effect of financial openness depends on the precondition of degree of competitiveness in the banking sectors. From Eq. (3), the partial derivative of financial development with respect to financial openness can be expressed as

$$\frac{\partial y_{it}}{\partial FO_{i,t-1}} = \beta_2 + \beta_4 COMP_{i,t-2} \quad (2.4)$$

2.3.2 System GMM Estimator Results

The results of the system GMM estimation presented in Table 2.4 confirm the findings of our previous analysis. Income level and inflation have a statistically significant impact on financial development, but no evidence suggests that trade openness has an effect. As suggested by Rajan and Zingales (2003), trade openness without financial openness is unlikely to encourage financial development. However, our focus is on the roles of financial openness and banking competition in financial development, and so we do not include the cross-product term between financial and trade openness in our specification to identify such a conditional effect. PMG estimation results show that trade openness effect has strong short-term heterogeneity, suggesting that it is limited by various heterogeneity factors such as financial openness (Rajan and Zingales, 2003).

Table 2.4 The results of GMM estimation

	(1)	(2)	(3)
Financial development (lagged)	1.0440*** (0.0274)	0.9500*** (0.0511)	0.9492*** (0.0509)
Financial openness (lagged)	0.0128* (0.0074)	0.0145** (0.0072)	0.0117* (0.0070)
Banking competition (lagged by two periods)	-0.0212** (0.0082)	-0.0013 (0.0022)	-0.0167** (0.0069)
Financial openness (lagged) $\times$ Banking competition (lagged by two periods)	-0.0680*** (0.0246)		-0.0525** (0.0217)
Frequency of systemic banking crises	-0.0718*** (0.0195)	-0.0365* (0.0202)	-0.0368* (0.0202)
Banking system's probability of default	-0.0020*** (0.0006)	-0.0014** (0.0006)	-0.0014** (0.0006)
Income levels		0.0371** (0.0181)	0.0369** (0.0183)
Trade openness		-0.0485 (0.0335)	-0.0459 (0.0338)
Inflation rate		-0.1892*** (0.0234)	-0.1899*** (0.0243)
Intercept	-5.8363* (3.1493)	-9.4001** (3.8656)	-9.3239** (3.8511)
No.Countries /No.Observations	102/1056	102/1037	102/1037
Specification test (p-values)			
(a) Sargan test	0.4313	0.2424	0.2344
(b) Serial correlation			
First-order	0.0046	0.0096	0.0097
Second-order	0.1888	0.2601	0.2358

Notes

1. Figures in parentheses are Windmeijer (2005)-robust standard error.
2. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively.
3. Year dummies are including in all the regressions.

We find that financial openness has a significant positive impact on financial development. We also find that the interaction term between financial openness and banking competition has a statically significant negative impact. This indicates that

financial openness is beneficial for financial development when the banking sector is competitive. Figure 2.2 plots the marginal effect of financial openness on financial development. In addition, we conclude that banking competition itself promotes financial development. We also find that financial fragility (defined as systemic banking crises and the probability of banking system default) deters financial development.

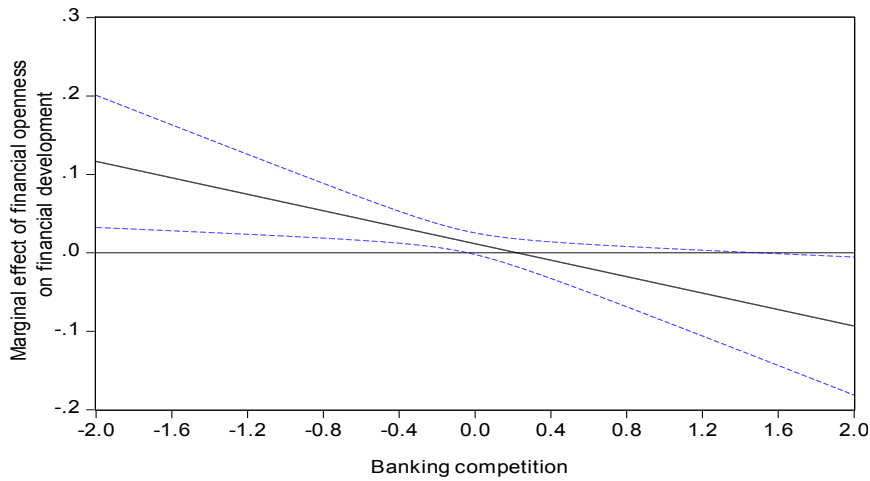


Fig. 2.2. Marginal effects of financial openness on financial development conditional on the degree of banking competition.
Note: The solid line represents the marginal effect. The dotted line represents 95% confidence intervals.

2.4 Conclusions

This study constitutes a significant contribution to the literature because it distinguishes between the long- and short-run effects of financial openness on financial development and shows that these effects differ greatly. Additionally, we find that in both PMG and GMM estimations, the adverse short-run effects of financial openness are associated with the degree of competition in the individual countries' banking sectors. Our results also suggest that banking competition promotes financial development. In addition, we find that financial fragility has a negative effect on financial development. Thus, we can conclude that the key question on the relationship between financial development and financial openness is not whether countries stand to benefit from financial openness in the long run, but rather under what circumstances they benefit. Policy makers should note that to gain full benefit from financial openness, a country (especially a developing one) must have a healthy and competitive financial environment—one that renders the

banking sectors more efficient and thus enables them to gain from greater cross-border opportunities.

Appendix 2.A

Sample of countries

Country	PMG estimation (70 countries)	GMM estimation (102 countries)
Algeria	—	×
Antigua and Barbuda	—	×
Argentina	×	×
Australia	×	×
Austria	×	×
Bahrain	—	×
Bangladesh	×	×
Belgium	×	×
Belize	—	×
Benin	×	×
Bhutan	—	×
Bolivia	—	×
Botswana	×	×
Brazil	×	×
Burundi	×	×
Cameroon	×	×
Canada	×	×
Chile	×	×
China	×	×
Colombia	—	×
Congo, Dem. Rep. of	—	×
Costa Rica	×	×
Côte d'Ivoire	×	×
Cyprus	×	×
Denmark	—	×
Djibouti	—	×
Dominican Rep.	×	×
Ecuador	—	×
El Salvador	—	×
Ethiopia	—	×
Finland	×	×
France	×	×
Gabon	×	×
Gambia, The	—	×

Germany	×	×
Ghana	×	×
Greece	×	×
Grenada	×	×
Guatemala	×	×
Guinea	—	×
Haiti	—	×
Honduras	×	×
India	×	×
Indonesia	×	×
Ireland	×	×
Israel	—	×
Italy	×	×
Japan	×	×
Jordan	×	×
Kenya	×	×
Korea	×	×
Lesotho	×	×
Libya	—	×
Madagascar	×	×
Malawi	×	×
Malaysia	×	×
Mali	×	×
Malta	×	×
Mauritania	×	×
Mauritius	—	×
Mexico	×	×
Morocco	×	×
Nepal	×	×
Netherlands	×	×
Nicaragua	—	×
Niger	×	×
Nigeria	×	×
Norway	×	×
Oman	×	×
Pakistan	×	×
Panama	—	×
Paraguay	—	×
Peru	×	×
Philippines	×	×

Portugal	—	×
Rwanda	—	×
Samoa	—	×
Saudi Arabia	—	×
Senegal	×	×
Sierra Leone	×	×
Singapore	×	×
Spain	×	×
Sri Lanka	×	×
St. Kitts and Nevis	—	×
Sudan	×	×
Suriname	—	×
Swaziland	×	×
Sweden	—	×
Syrian Arab Republic	×	×
Tanzania	—	×
Thailand	×	×
Togo	×	×
Trinidad and Tobago	×	×
Tunisia	—	×
Turkey	×	×
Uganda	×	×
United Arab Emirates	—	×
United Kingdom	×	×
United States	×	×
Uruguay	×	×
Venezuela, Rep. Bol.	×	×
Zambia	×	×

Note: × indicates that the country is included in the sample, while — indicates that the country is not included in the sample.

Appendix 2.B

Descriptive statistics					
Variable	Obs.	Mean	Std. Dev.	Min.	Max.
<i>(1) Dataset used in PMG estimator, 70 countries (1980-2011)</i>					
Private domestic credit/GDP	2199	0.4835	0.4580	0.0154	2.9646
M2/GDP	2097	0.5504	0.4508	0.0655	2.8340
Financial openness	2238	-0.3910	0.5832	-5.4103	2.5561
Income levels	2229	7.9831	1.6964	4.9683	11.1244
Inflation rate	2233	0.2430	1.8263	-0.2763	50.4878
Trade openness	2232	0.7242	0.5123	0.0632	4.3966
<i>(2) Dataset used in GMM estimator, 102 countries (1998-2011)</i>					
Private domestic credit/GDP	1393	0.5361	0.5100	0.0020	2.9646
Financial openness	1427	-0.2730	1.2155	-3.0820	17.2070
Boone indicator	1287	-0.0576	0.2076	-2.0820	5.9680
Frequency of systemic banking crises	1428	0.0763	0.2656	0.0000	1.0000
Banking system's probability of default (z-score)	1312	15.5694	9.8421	-11.5740	65.2840
Income levels	1418	8.0955	1.7002	4.8715	11.1244
Inflation rate	1421	0.0925	0.7129	-0.3281	26.3012
Trade openness	1407	0.7920	0.4731	0.1587	4.3966

Chapter 3

Macroeconomic Impacts of Oil Prices and Underlying Financial Shocks

3.1 Introduction

Large fluctuations in oil prices and their high volatility have long been sources of instability in the global economy. In particular, the sharp rise in oil prices during the commodity boom that started in the early 2000s posed serious challenges to macroeconomic management in both developed and developing countries. Against this background, a large body of literature has empirically examined the underlying causes of oil price fluctuations and their macroeconomic impacts. Early research mainly focuses on the relationships between oil prices and economic activity (see, for example, Hamilton, 1983; Hooker, 1996), finding a strong negative relation between rising oil prices and GDP growth in many countries. Previous studies also suggest a positive association between rising oil prices and inflationary pressures on the economy (see, for example, Cunado and Perez de Gracia, 2005).

Further, while a growing body of literature examines the effect of oil prices on the stock market, there is no robust consensus about the effect of oil price shocks on stock market returns. Ciner (2001) finds that a statistically significant relationship exists between oil price futures and real stock returns, but that the correlation is non-linear. Similarly, Aloui et al. (2008) find that changes in oil prices significantly increase the volatility of stock market returns in six developed countries. By contrast, Jammazi and Aloui (2010) show that oil price shocks do not affect stock market returns during recession phases.²⁰

Recent studies have shown that the effects of oil price shocks on stock markets depend on whether the country is an oil importer or an oil exporter. For example, Park and Ratti (2008) show that oil price shocks account for a statistically significant 6% of the volatility in real stock returns. Moreover, the increased volatility of oil prices significantly depresses real stock returns in many European oil-importing countries. Aroui and Rault (2012), on the other hand, report that oil price increases positively influence stock prices in Gulf Cooperation Council countries, except in Saudi Arabia.

Despite the accumulation of empirical evidence, however, two major deficiencies are evident in the traditional approach to modeling oil price shocks frequently used in the literature. First, although reverse causality may run from real economic activities to oil prices, oil price shocks are assumed to be exogenous. Second, the recent literature presents evidence that the relation between oil prices and stock prices depends on the origin or nature of oil price shocks (see, for example, Ciner, 2013; Degiannakis et al.,

²⁰ Similar findings are reported in Huang et al. (1996) and Cong et al. (2008).

2013). These results indicate that the macroeconomic impacts of oil price shocks could depend on the underlying causes, which has not been fully taken into account in previous analyses.

Kilian (2009) proposes a two-step approach to analyzing the macroeconomic impacts of oil price shocks in order to overcome these shortcomings. In the first step, a vector autoregression (VAR) that includes oil production, global economic activity, and oil prices as endogenous variables is estimated in order to identify three types of structural shocks that underlie oil price changes: an oil supply shock, an aggregate demand shock, and an oil market-specific demand shock that reflects an unexpected change in precautionary oil demand. In the second step, ordinary least squares (OLS) regressions are estimated to evaluate the impact of the identified structural shocks on the macroeconomic indicators. Kilian (2009) adopts this framework to demonstrate that US macroeconomic indicators respond differently to oil price shocks depending on the types of underlying shocks.

Kilian's (2009) two-step approach has been employed extensively by recent studies of how oil price shocks influence real economic activity and the stock markets. For instance, it has been shown that the consideration of the origins of oil price shocks is crucial, since different shocks in the crude oil market have diverse effects on real activity and stock markets (see, among others, Kilian and Park, 2009; Apergis and Miller, 2009; Yoshizaki and Hamori, 2013). However, to our best knowledge, no authors have yet attempted to extend Kilian's (2009) framework in order to identify an exogenous shock that arises from unexpected changes in financial market conditions and examine the consequent macroeconomic impacts of oil price changes. This extension must be meaningful because there is emerging evidence of the so-called financialization of commodity markets, a phenomenon characterized by a high degree of price correlation among a broad set of commodities as well as between commodities and financial assets, presumably due to the greater participation of financial investors in commodity markets (Henderson et al., 2012; Nissanke, 2012; Singelton, 2012; Tang and Xiong, 2012; Buyuksahin and Robe, 2012; Morana, 2013; Basak and Pavlova, 2013). A consequence of the financialization process is that commodity prices such as oil prices are determined not only by their supply and demand but also by the financial market conditions that affect financial investment.

The financial collapse of 2008 has sparked renewed interest in the accurate measurement of financial shocks to the real economy. In this context, many researchers have developed methods for constructing financial condition indexes, which contain

information on financial variables selected not only from stock markets but also from bond market and the banking system. This approach is necessary because individual indicators (e.g., those only derived from stock markets) may provide ambiguous signals if financial conditions do not change simultaneously or uniformly. In this paper, we use the Kansas City Financial Stress Index (KCFSI) developed by the Federal Reserve Bank of Kansas City as a proxy for global financial market conditions. The KCFSI is a composite index designed to measure the level of stress in US financial markets. It includes 11 financial variables such as TED spread, Treasury and corporate bond spreads, and the volatility of stock prices. By assuming that US financial market conditions reflect, to a significant degree, the overall conditions in global financial markets, the KCFSI provides a reasonable measure of stress in global financial markets. It has previously been demonstrated that an increase in financial stress is associated with higher financing costs, greater economic uncertainty, and increased risk premiums, resulting in depressed corporate investment and household spending and declining real economic activity (Hakkio and Keeton, 2009; Davig and Hakkio, 2010).

The remainder of the paper is organized as follows: Section 3.2 identifies those structural shocks that underlie oil price changes by estimating a structural VAR; Section 3.3 examines the impact of the identified structural shocks on the macroeconomic indicators in five major industrial countries, namely France, Germany, Japan, the UK, and the US; and Section 3.4 presents the summary and conclusions.

3.2 Structural Shocks that Underlie Oil Price Changes

3.2.1 Structural VAR model

In this section, we identify the structural shocks that underlie oil price changes by estimating a VAR model. The structural representation of the VAR model is as follows:

$$A_0 y_t = \alpha + \sum_{i=1}^p A_i y_{t-i} + \varepsilon_t \quad (3.1)$$

where y_t is a (4×1) vector that contains global crude oil production (COP), global real economic activity (REA), real oil prices (ROP), and the KCFSI, A_0 denotes a contemporaneous coefficient matrix, α denotes a vector of constant terms, and ε_t denotes a vector of serially and mutually uncorrelated structural shocks. Under the appropriate identifying restrictions, structural shocks can be recovered from the estimated reduced-form errors by using the following relationship:

$$e_t = A_0^{-1} \varepsilon_t \quad (3.2)$$

where e_t denotes the reduced-form errors.

COP is measured by using the total world crude oil production provided by the Oil and Gas Journal. REA is measured by using the index developed by Kilian (2009).²¹ This index is constructed by using single-voyage freight rates for bulk dry commodity cargoes. It is then deflated by the US consumer price index and linearly de-trended in order to remove the effects of technological advances in shipbuilding and other long-term trends in demand for seartransport. ROP is measured by using the US West Texas Intermediate price deflated by the US producer price index. The data source is the IMF's International Financial Statistics. Finally, as discussed in Section 3.1, we use the KCFSI as a proxy for global financial market conditions. The KCFSI is normalized to have a mean value of zero and a standard deviation of one. A positive (negative) value indicates that financial stress is above (below) the long-run average, which would discourage (encourage) financial investment in oil markets.

A VAR is estimated by using the log-difference of COP and ROP and the levels of the KCFSI and REA divided by 100.²² The sample period runs from January 1991 to December 2012. In line with the approach taken by Kilian (2009), VAR model is estimated by using 24 lags of each variable to allow for the potentially long-run effects of structural oil price shocks on the economy.²³ We then identify structural shocks by using the Choleski decomposition, with the order being COP, REA, ROP, and the KCFSI. This order determines the exogeneity of the variables; a shock on a particular variable has a contemporaneous effect on the variables ordered after that particular variable but not before it. Following Kilian (2009), COP is assumed to be least responsive presumably due to the high adjustment costs of oil production, followed by REA and ROP. By adopting this ordering, we assume that the oil supply shock,

²¹ The data for the index are available at <http://www-personal.umich.edu/~lkilian/>.

²² According to the augmented Dickey–Fuller test, all transformed variables are stationary at the 5% significance level except the transformed REA. However, the DF-GLS test indicates that the transformed REA is stationary at the 5% significance level. We therefore assume that all transformed variables are stationary in the analysis presented herein. We also checked that the stability condition of a VAR, which requires all characteristic roots to lie within the unit circle, is met.

²³ The Akaike Information Criterion (AIC) indicates a lag length of 6. The estimation results based on 6 lags are similar to those based on 24 lags.

aggregate demand shock, and oil-specific demand shock can be captured by using the structural shock to COP, REA, and ROP, respectively. The KCFSI is placed after ROP based on the assumption that oil prices have contemporaneous effects on financial markets but not vice versa.²⁴ This assumption is in line with the existing literature. For example, Kilian and Park (2009) employ a VAR model and investigate stock market fluctuations associated with oil price shocks with the ordering of oil production, real economic activity real oil prices, and real stock returns. Basher et al. (2012) estimate a structural VAR model in order to investigate the dynamic relationship between oil prices, exchange rates, and emerging market stock prices, imposing the restriction that the interest rate spread (TED spread) is allowed to react contemporaneously to oil prices, but not vice versa. In the same vein, Kang and Ratti (2013) examine the relationship between oil shocks, economic policy uncertainty, and stock prices by estimating a structural VAR with the ordering of oil production, real economic activity, real oil prices, a proxy variable for economic policy uncertainty, and real stock returns. Kilian and Vega (2011) show that oil prices do not respond contemporaneously to domestic macroeconomic news, which is consistent with the commonly used identifying assumption that oil price shocks are predetermined with respect to domestic macroeconomic aggregates. Hence, the reduced-form VAR is obtained by multiplying both sides of Eq. (3.1) by A_0^{-1} , which has the following recursive structure:

$$\begin{pmatrix} u_{1t} \\ u_{2t} \\ u_{3t} \\ u_{4t} \end{pmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ a_{21} & 1 & 0 & 0 \\ a_{31} & a_{32} & 1 & 0 \\ a_{41} & a_{42} & a_{43} & 1 \end{bmatrix} \times \begin{pmatrix} \mathcal{E}_{oil\ supply\ shock} \\ \mathcal{E}_{aggregate\ demand\ shock} \\ \mathcal{E}_{oil-specific\ demand\ shock} \\ \mathcal{E}_{financial\ shock} \end{pmatrix}$$

3.2.2 Impulse Response to Structural Shocks

To illustrate the relative importance of the identified structural shocks as sources of oil price changes, the cumulative impulse responses of ROP and other variables to a one-standard deviation shock are shown in Figure 3.1. The dotted lines represent two-standard error bands.

²⁴ The results of the impulse response analyses are similar irrespective of whether the KCFSI is placed before REA or before ROP.

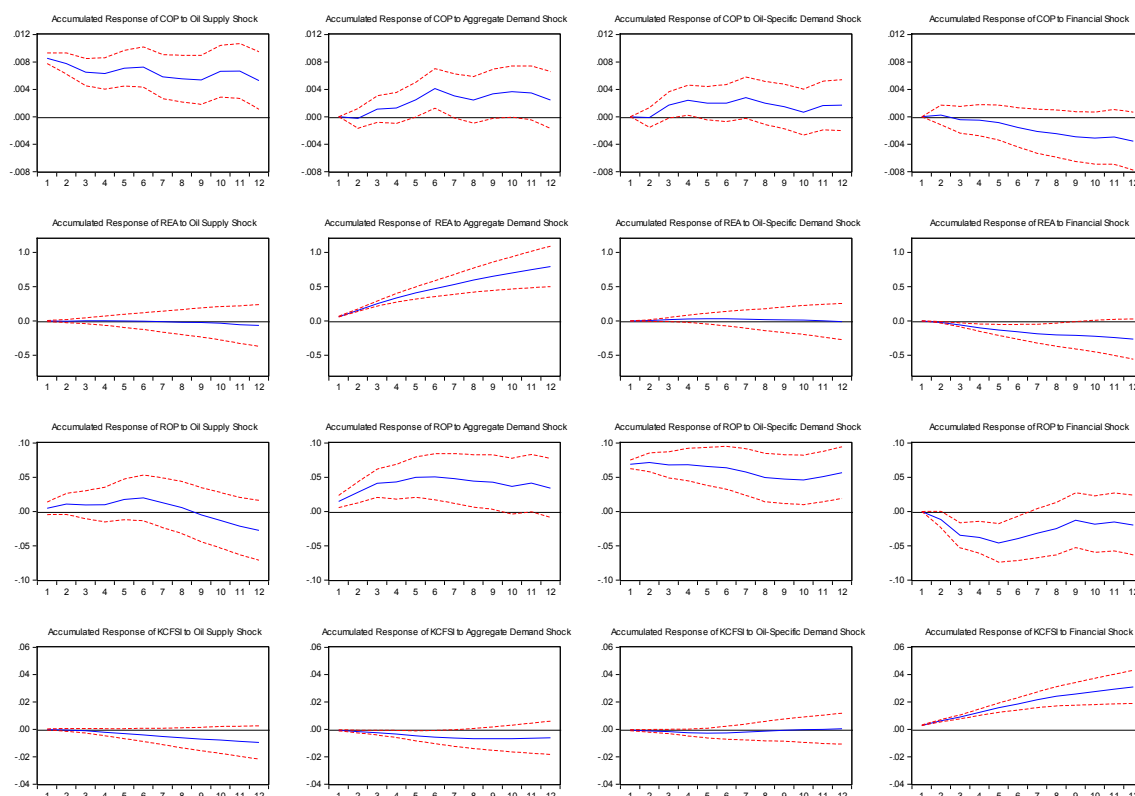


Fig. 3.1. Accumulated Response to Structural One S.D. Innovations with Two-Standard Error Confidence Bands. Note: The dotted lines represent two-standard error bands.

Figure 3.1 highlights that the effect of an unexpected increase in oil supply on ROP is small and statistically insignificant. By contrast, an unexpected increase in aggregate demand causes a statistically significant increase in ROP, which peaks after approximately six months. An unexpected increase in oil market-specific demand has an immediate and relatively large positive impact on ROP, which is persistent and statistically significant. These results are broadly consistent with those of Kilian (2009).

The impulse response of ROP to a financial shock is of particular interest herein. The figure indicates that ROP declines in the face of a positive financial shock, which implies worsening financial market conditions. Such an unexpected worsening of financial conditions causes a statistically significant decline in ROP, which bottoms out after approximately five months. We also find that a positive financial shock depresses real activity.²⁵

²⁵This result is consistent with the findings of Basher et al. (2012). They employ a structural VAR to model the dynamic relationships among oil supply, REA, emerging stock market prices, TED spread, the exchange rate index for major currencies, and real oil prices. The empirical results show that credit tightening, measured by an increase in TED spread, tends to depress real oil prices, REA, and emerging stock market prices.

The findings presented in this section indicate that a financial shock is an important determinant of oil prices. This result seems to lend some support to the view that oil markets have become financialized in the sense that oil prices are significantly driven by changes in financial market conditions that affect financial investment.

3.2.3 Variance Decomposition

In this section, we investigate the contribution of different structural shocks to the fluctuations of the variables in the VAR by estimating the forecast error variance decomposition. Table 3.1 shows the share of the fluctuations in ROP and the KCFSI caused by its own shocks (the oil-specific demand shock and financial shock, respectively) compared with the shocks to the other variables.

A major share of ROP fluctuations is accounted for by its own shock (i.e., oil-specific demand shock), although the contribution of this shock declines over time. We note that the financial shock accounts for a larger share of ROP fluctuations than the demand shock after 3 months and thereafter. The financial shock explains approximately 15% of ROP fluctuations, while the demand shock accounts for just 10% after 12 months. On the other hand, the oil supply shock accounts for the lowest share of ROP fluctuations (5%), which could be seen as an indication of its low explanatory power. These results are broadly similar to the findings for the impulse responses analysis presented above.

The fluctuations in the KCFSI are mostly caused by its own shock (i.e., financial shock) at all time horizons. The combined contribution of the other shocks to KCFSI fluctuations accounts for approximately 18% after 12 months. Interestingly, we find that the share of KCFSI fluctuations caused by other shocks varies considerably over time. For example, the aggregate demand shock accounts for a larger share of KCFSI fluctuations than the oil supply shock before 8 months. However, the oil supply shock accounts for a larger share of KCFSI fluctuations than the aggregate demand shock thereafter. We also find that the oil-specific demand shock accounts for approximately 3% of KCFSI fluctuations throughout the study period, which is constantly lower than the demand shock.

Table 3.1 Variance Decomposition

Variance Decomposition of ROP				
Period	Oil Supply Shock	Aggregate Demand Shock	Oil-Specific Demand Shock	Financial Shock
1	0.416744 (1.01472)	4.253206 (2.69691)	95.33005 (2.81577)	0.000000 (0.00000)
2	1.160561 (1.79303)	7.290005 (3.77865)	88.94507 (4.51010)	2.604366 (2.47888)
3	1.052779 (1.77647)	9.280690 (4.26891)	78.60043 (5.82840)	11.06610 (4.78827)
4	1.052130 (1.95673)	9.338240 (4.32213)	78.42302 (5.89882)	11.18661 (4.81225)
5	1.958448 (2.50535)	9.809393 (4.41874)	76.30111 (6.03069)	11.93105 (4.87145)
6	2.016218 (2.64949)	9.732950 (4.39945)	75.73076 (5.95444)	12.52008 (4.84863)
7	2.754364 (2.98655)	9.594950 (4.32580)	74.51508 (5.94196)	13.13561 (4.91509)
8	3.426793 (3.31214)	9.535156 (4.32541)	73.58442 (6.03936)	13.45363 (4.93794)
9	4.907015 (3.82581)	9.199454 (4.32022)	70.75708 (6.16794)	15.13645 (5.27097)
10	5.718270 (4.09624)	9.550469 (4.42895)	69.41746 (6.08238)	15.31380 (5.05673)
11	6.634490 (4.31871)	9.678430 (4.37782)	68.51071 (6.03151)	15.17637 (5.02295)
12	7.022894 (4.39053)	10.20350 (4.45174)	67.63761 (5.99786)	15.13600 (4.90906)

Variance Decomposition of KCFSI				
Period	Oil Supply Shock	Aggregate Demand Shock	Oil-Specific Demand Shock	Financial Shock
1	0.141737 (0.77823)	5.336829 (2.80845)	3.353950 (2.29265)	91.16748 (3.56965)
2	0.530050 (1.46804)	6.656845 (3.80057)	2.761399 (2.55255)	90.05171 (4.65829)
3	1.491088	6.659810	2.868996	88.98011

	(2.46534)	(4.47458)	(3.16113)	(5.72175)
4	3.645055	6.522759	3.183499	86.64869
	(3.76431)	(4.97906)	(3.68485)	(6.65331)
5	4.557677	8.026319	2.662871	84.75313
	(4.59902)	(5.99462)	(3.65944)	(7.66033)
6	5.199939	7.984419	2.348945	84.46670
	(5.39784)	(6.48512)	(3.32788)	(8.30988)
7	6.056162	7.544395	2.409160	83.99028
	(6.17774)	(6.59809)	(3.07808)	(8.81112)
8	6.682029	7.025904	2.692125	83.59994
	(6.87982)	(6.61468)	(3.28161)	(9.30681)
9	7.157214	6.736560	3.113042	82.99318
	(7.48462)	(6.63824)	(3.85629)	(9.82989)
10	7.388232	6.483358	3.116752	83.01166
	(7.89428)	(6.51613)	(4.21229)	(10.1276)
11	8.094858	6.296603	3.052101	82.55644
	(8.45574)	(6.39416)	(4.48203)	(10.4587)
12	8.457265	6.133896	3.129012	82.27983
	(8.80301)	(6.37147)	(4.82835)	(10.7651)

Note: The values in parentheses represent the t-statistics when coefficients' standard errors were generated from a Monte-Carlo simulation with 1000 replications.

3.3 Macroeconomic Impacts of Structural Shocks

3.3.1 OLS Regressions

In this section, we examine the impact of the identified structural shocks on the macroeconomic indicators by estimating OLS regressions.²⁶ The explanatory variables are the four structural shocks identified in Section 3.2, which are standardized by subtracting the mean and dividing by the standard deviation. The dependent variables are the following three macroeconomic indicators: the index of industrial production (IIP), consumer price index (CPI), and stock price index (STP).²⁷ An OLS regression is

²⁶Alternatively, we can use a finite distributed lag model such as an Almon lag model in order to avoid potential multicollinearity among the explanatory variables in an OLS regression. However, the variance inflation factor is less than two in all the estimated OLS regressions, indicating that the degree of multicollinearity is low. We therefore use the results of OLS regressions in the presented analysis.

²⁷All the data for IIP and CPI are seasonally adjusted. STP is converted into the real value by deflating by using CPI. The data source is the IMF's International Financial Statistics.

estimated for each of these macroeconomic indicators by using quarterly data from 1993Q1 to 2012Q4.²⁸ The sample includes data for France, Germany, Japan, the UK, and the US. Following Kilian (2009), measures of quarterly shocks are constructed by averaging monthly shocks for each quarter:

$$\hat{g}_{jt} = \frac{1}{3} \sum_{i=1}^3 \hat{\varepsilon}_{j,t,i} \quad j=1,2,3,4$$

where $\hat{g}_{jt}$ denotes the j th structural shock in the t th quarter and $\hat{\varepsilon}_{j,t,i}$ denotes the estimated j th structural shock in the i th month of the t th quarter.

The effects of the estimated structural shocks on the macroeconomic indicators are examined by estimating the following regressions:

$$\begin{aligned} \Delta IIP_t &= \alpha_j + \sum_{i=0}^{12} \phi_{ji} \hat{g}_{jt-i} + r_{jt} & j=1,2,3,4 \\ \Delta CPI_t &= \beta_j + \sum_{i=0}^{12} \tau_{ji} \hat{g}_{jt-i} + v_{jt} & j=1,2,3,4 \\ \Delta STP_t &= \gamma_j + \sum_{i=0}^{12} \omega_{ji} \hat{g}_{jt-i} + s_{jt} & j=1,2,3,4 \end{aligned}$$

where Δ denotes the percentage change in the relevant variables, α_j , β_j , and γ_j are constant terms, ϕ_{ji} , τ_{ji} , and ω_{ji} are the impulse response coefficients at horizon i , and r_{jt} , v_{jt} , and s_{jt} are error terms. The maximum lag is determined by the maximum horizon of the impulse function, which is set to 12 quarters. Since there is a potential problem of serial correlation in the error terms, the block bootstrap method is used to infer the estimated coefficients. Specifically, we use an overlapping moving block bootstrap method with block size 4 and 20,000 bootstrap replications.²⁹

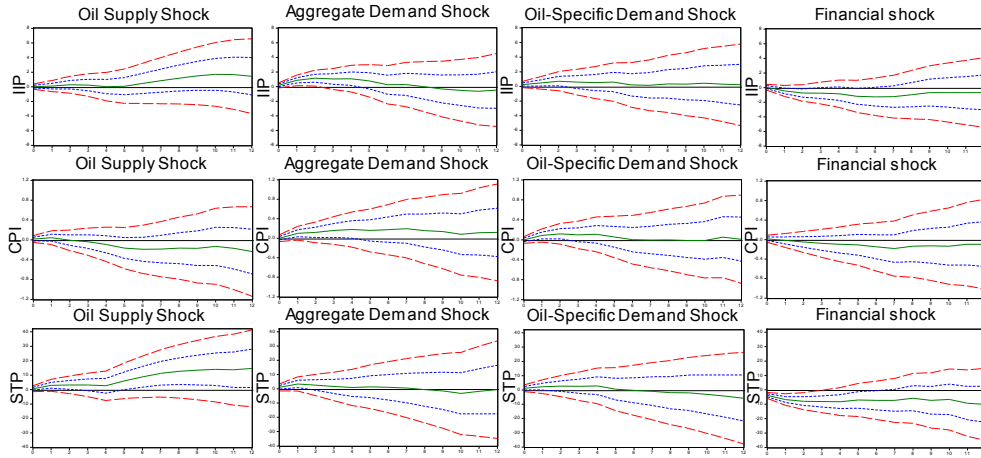
3.3.2 Estimation Results

The cumulative impulse responses of the macroeconomic variables to the structural shocks are shown in Figure 3.2. The dotted and dashed lines represent one-standard error and two-standard error bands, respectively. In the discussion below, the statistical significance is determined based on one-standard error bands.

²⁸ The start date of 1993 reflects the need to accommodate lags in the VAR.

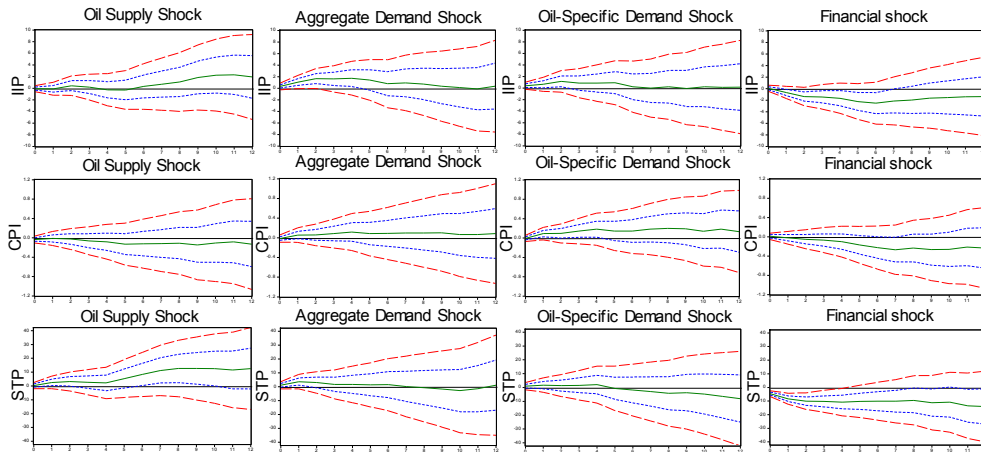
²⁹ See MacKinnon (2006) for a survey of bootstrapping methods.

[France]



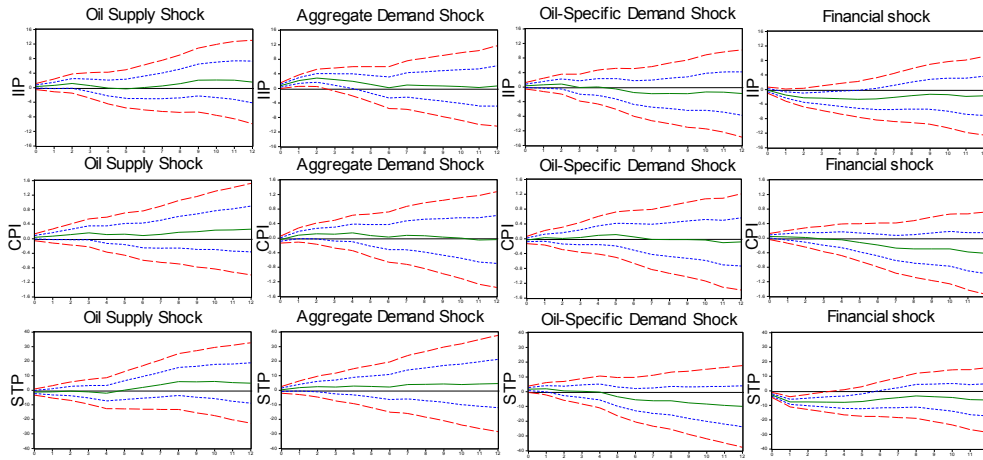
Note: The dotted and dashed lines represent one-standard and two-standard error bands, respectively.

[Germany]



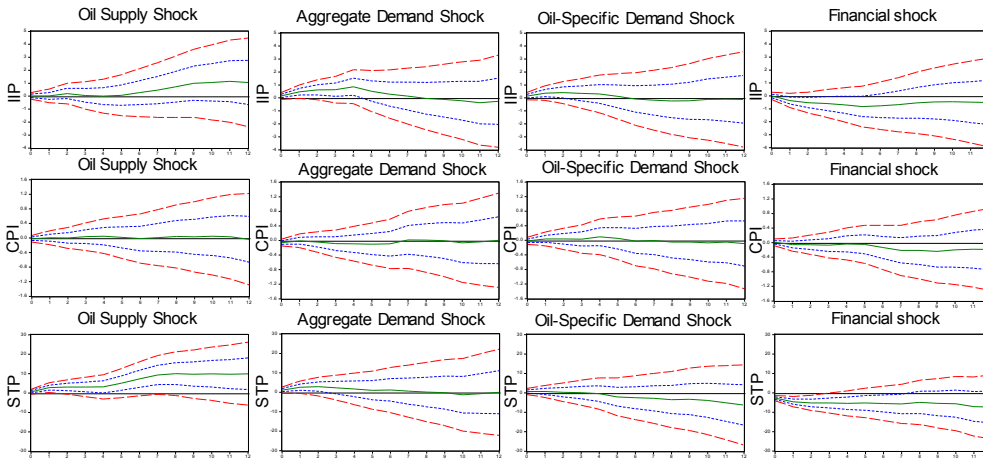
Note: The dotted and dashed lines represent one-standard and two-standard error bands, respectively.

[Japan]

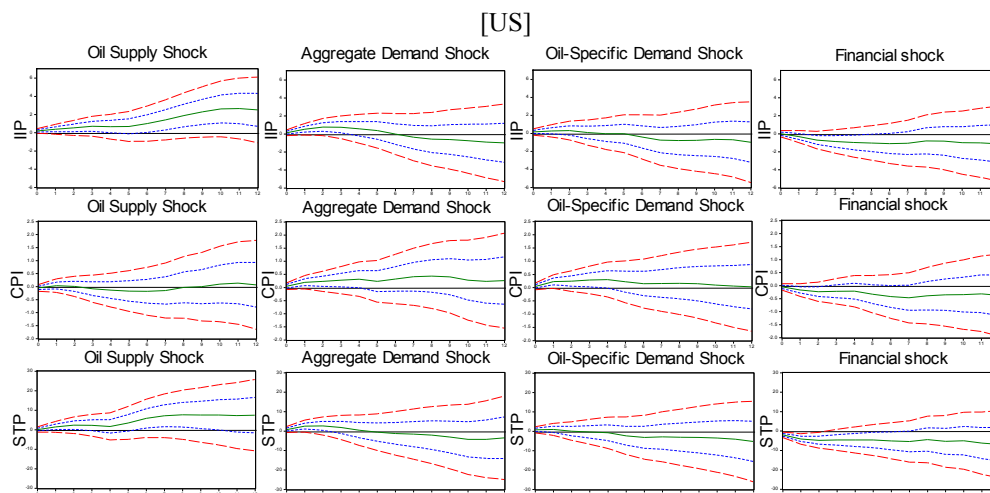


Note: The dotted and dashed lines represent one-standard and two-standard error bands, respectively.

[UK]



Note: The dotted and dashed lines represent one-standard and two-standard error bands, respectively.



Note: The dotted and dashed lines represent one-standard and two-standard error bands, respectively.
Figure 3.2 Cumulative Responses of the Macroeconomic Indicators to the Structural Shocks

A positive oil supply shock (an unexpected increase in oil supply) causes a sustained and statistically significant increase in IIP only in the US. The effects of an oil supply shock on CPI are statistically insignificant in all countries.³⁰ By contrast, a positive oil supply shock leads to a statistically significant increase in STP in all countries except Japan. The corresponding responses in STP are statistically significant for all or almost all horizons in France and the UK. Likewise, the corresponding responses in STP are positive in Germany and the US, although they are statistically significant only after the second year.

A positive aggregate demand shock (an unexpected increase in aggregate demand) causes an increase in IIP in all countries in the first year. In particular, the positive response of IIP is highly statistically significant based on two-standard error bands in France and Japan. However, these cumulative responses peak in the second half of the first year, followed by a statistically insignificant decline towards or below the initial level in the third year. The result is fully consistent with one of the key findings of Kilian (2009). The corresponding response of STP follows a similar pattern, except in Japan. These results indicate that the initial direct effect of the aggregate demand shock on IIP and STP wears off over time, which is offset almost fully or more than fully by the shock's lagged indirect effect through oil price changes. Moreover, a positive aggregate demand shock causes a statistically significant increase in CPI in France and the US in the first year.

³⁰ The effect of the oil supply shock on CPI is similar to that found by Kilian (2009).

A positive oil market-specific demand shock (an unexpected increase in precautionary oil demand) causes a temporary and statistically significant increase in IIP only in European countries. In the US and Japan, the corresponding responses are statistically insignificant. A positive oil market-specific demand shock causes a statistically significant increase in CPI in France, Germany, and the US. By contrast, the shock does not cause a statistically significant increase in STP in all countries. This result is consistent with that of Degiannakis et al. (2013), which find that the precautionary demand oil price shock has no significant impact on the stock returns of European industrial sectors.

The impulse responses of the macroeconomic indicators to financial shocks are of particular interest in this paper. As expected, a positive financial shock (an unexpected increase in the KCFSI, implying increased stress in financial markets) causes a highly statistically significant decline in IIP in all countries. This finding indicates that as financial stress rises, increased financing costs and greater uncertainty depress REA.³¹ These cumulative responses of IIP bottom out after the second year, followed by a statistically insignificant increase. However, the response remains negative in the following period. A similar response pattern is observed for STP. These results indicate that, unlike the aggregate demand shock, the initial impact of the financial shock is sustained presumably due to weaker offsetting indirect effects through oil price changes. There is also a noticeable difference in the associated response pattern between the aggregate demand shock and financial shock, illustrating the importance of identifying the latter shock as an additional source of oil price fluctuations. Finally, a positive financial shock causes a statistically significant decline in CPI in the US.

The findings presented above can be summarized as follows. First, we find evidence that the macroeconomic impact of oil price shocks depends on the underlying causes and that each shock is associated with a distinct response pattern, which is fully consistent with the results in the literature, notably those of Kilian (2009). Second, we advance our understanding of the relation between oil price shocks and stock prices. More specifically, we find that the impact of oil supply shock on stock prices is more persistent, whereas the net effect of the aggregate demand shock on stock price changes over time. This result is broadly in line with the findings of Ciner (2013) and Degiannakis et al. (2013), which show that the relation between oil prices and stock prices depends on the origin or nature of oil price shocks. Finally, we find that the macroeconomic

³¹ The result is consistent with the finding presented in Figure 3.1, which shows that a positive financial shock causes a negative and statistically significant effect on REA.

impact of the financial shock is significant and thus comparable with that of the aggregate demand shock. Moreover, we find a noticeable difference in the associated response pattern between the aggregate demand shock and financial shock, illustrating the importance of identifying the latter shock as an additional source of oil price fluctuations.

3.4 Conclusions

We extend Kilian's (2009) framework to identify an exogenous shock that arises from an unexpected change in financial market conditions and examine the consequent macroeconomic impacts of oil price changes. This extension is meaningful because there is emerging evidence of the financialization of commodity markets, a phenomenon characterized by a high degree of price correlation among a broad set of commodities as well as between commodities and financial assets presumably due to the greater participation of financial investors in commodity markets.

By applying Kilian's (2009) method, we identify a structural financial shock that causes changes in oil prices, assess the relative importance of a financial shock as a source of oil price changes, and examine its macroeconomic impacts. In the first step, we identify four types of structural shocks, including the financial shock, by estimating a VAR. The impulse response analysis shows that a positive financial shock causes a statistically significant decline in oil prices, indicating that the financial shock is a key determinant of oil prices. Moreover, the estimated variance decomposition indicates that the financial shock has a relatively high explanatory power for oil price fluctuations. In the second step, we examine the impact of underlying structural shocks on the macroeconomic indicators in five major industrial countries. The impulse response analysis indicates that the macroeconomic impact of the financial shock is significant and that the importance of financial shocks as sources of macroeconomic fluctuations is comparable with that of the aggregate demand shock. This paper also furthers the understanding of the relation between oil price shocks and stock prices by showing that the impact of the oil supply shock on stock prices is more persistent, whereas the net effect of the aggregate demand shock on stock price changes over time.

The key policy implications derived from our analysis can be summarized as follows. First, policymakers must explicitly take account of changes in global financial market conditions when analyzing the macroeconomic impacts of oil price shocks. Second, the design of a stabilization policy in response to oil price shocks must be tailored in accordance with the underlying causes because different underlying shocks

could have different macroeconomic impacts in different countries. Finally, a stabilization policy is required to be forward-looking because the net effect of underlying shocks, such as aggregate demand shocks and financial shocks, could differ and change significantly over time.

References

- Acemoglu, D., Johnson, S., Robinson, J. A., 2002. Reversal of fortune: Geography and institutions in the making of the modern world income distribution. *Quarterly Journal of Economics* 117(4), 1231-1294.
- Acemoglu, D., Johnson, S., Robinson, J.A., 2001. The colonial origins of comparative development: An empirical investigation. *American Economic Review* 91(5), 1369-1401.
- Aghion, P., Bolton, P., 1997. A theory of trickle-down growth and development. *Review of Economic Studies* 64 (2), 151-172.
- Aloui, C., Jammazy, R., Dhakhlaoui, I., 2008. Crude oil market shocks and stock market returns. *Journal of Energy Markets* 3, 69-96.
- Apergis, N., Miller, S.M., 2009. Do structural oil-market shocks affect stock prices? *Energy Economics* 31, 569-575.
- Arellano, M., Bond, S., 1991. Some tests of specification for panel data: Monte carlo evidence and an application to employment equations. *Review of Economic Studies* 58(2), 277-297.
- Arellano, M., Bover, O., 1995. Another look at the instrumental variable estimation of error-components models. *Journal of Econometrics* 68(1), 29-51.
- Arouri, M., Rault, C., 2012. Oil prices and stock markets in GCC countries: Empirical evidence from panel analysis. *International Journal of Finance and Economics* 17, 242-253.
- Baltagi, B.H., Demetriades, P.O., Law, S., 2009. Financial development and openness: evidence from panel data. *Journal of Development Economics* 89, 285–296.
- Banerjee, A.V., Newman, A., 1993. Occupational choice and the process of development. *Journal of Political Economy* 101 (2), 274-298.
- Basak, S., Pavlova, A., 2013. A Model of Financialization of Commodities.
- Basher, S.A., Haug, A.A., Sadorsky, P., 2012. Oil prices, exchange rates and emerging stock markets. *Energy Economics* 34, 227-240.
- Beck, T.H.L., Demirgüç-Kunt, A., Levine, R., 2007. Finance, Inequality and the Poor. *Journal of Economic Growth* 12(1), 27-49.
- Berger, A.N., Demirgüç-Kunt, A., Levine, R., Haubrich, J.G., 2004. Bank Concentration and Competition: An Evolution in the Making. *Journal of Money, Credit, and Banking*

- 36, 433-451.
- Boubaker, S., Jouini, J., 2014. Linkages between emerging and developed equity markets: Empirical evidence in the PMG framework. *North American Journal of Economics and Finance* 29, 322-335.
- Buyuksahin, B., Robe, M.A., 2012. Speculators, Commodities and Cross-Market Linkages. SSRN: <http://ssrn.com/abstract=1707103>
- Chinn, M.D., Ito, H., 2002. Capital Account Liberalization, Institutions and Financial Development: Cross Country Evidence. NBER Working Papers No. 8967.
- Chinn, M.D., Ito, H., 2006. What matters for financial development? Capital controls, institutions, and interactions. *Journal of Development Economics* 81, 163-192.
- Ciner, C., 2013. Oil and stock returns: frequency domain evidence. *Journal of International Financial Markets, Institutions and Money* 23, 1-11.
- Ciner, C., 2001. Energy shocks and financial markets: Nonlinear linkages. *Studies in Nonlinear Dynamics and Econometrics* 5, 203-212.
- Claessens, S., Laeven, L., 2004. What Drives Bank Competition? Some International Evidence. *Journal of Money, Credit, and Banking* 36, 563-583.
- Clarke, G.R.G., Xu, L.C., Zou, H., 2006. Finance and income inequality: What do the data tell us? *Southern Economic Journal* 72 (3), 578-596.
- Cong, R.G., Wei, Y.M., Jiao, J.L., Fan, Y., 2008. Relationships between oil price shocks and stock market: An empirical analysis from China. *Energy Policy* 36, 3544-3553.
- Cunado, J., Perez de Gracia, F., 2005. Oil prices, economic activity and inflation: evidence for some Asian countries. *The Quarterly Review of Economics and Finance* 45, 65-83.
- Davig, T., Hakkio, C., 2010. What is the effect of financial stress on economic activity. *Economic Review (Q II)*, 35-62.
- Degiannakis, S., Filis, G., Floros, C., 2013. Oil and stock returns: Evidence from European industrial sector indices in a time-varying environment. *Journal of International Financial Markets, Institutions and Money* 26, 175-191.
- Dell'Ariccia, G., Marquez, R., 2004. Information and bank credit allocation. *Journal of Financial Economics* 72, 185-214.
- Demetriades, P.O., Hussein, K.A., 1996. Does financial development cause economic growth? Time-series evidence from 16 countries. *Journal of Development Economics* 51, 387-411.

- Demirgüç-Kunt, A., Detragiache, E., 1998. The determinants of banking crises: Evidence from industrial and developing countries. *IMF Staff Papers* 45(1), 81-109.
- Demirgüç-Kunt, A., Detragiache, E., 2000. Monitoring banking sector fragility: a multivariate logit approach. *World Bank Economic Review* 14 (2), 287-307.
- Demirgüç-Kunt, A., Levine, R., 2008. Finance, Financial Sector Policies, and Long Run Growth. Policy Research Working Paper, WPS4469, Washington DC: World Bank
- Easterly, W., Fischer, S., 2001. Inflation and the poor. *Journal of Money, Credit, and Banking* 33(2), 160-178.
- Edison, H., Klein, M., Ricci, L., Sløk, T., 2004. Capital account liberalization and economic performance: survey and synthesis. *IMF Staff Papers* 51 (2), 220-256.
- Erosa, A., Ventura, G., 2002. On Inflation as a Regressive Consumption Tax. *Journal of Monetary Economics* 49(4), 651-875.
- Fischer, R., Valenzuela, P., 2013. Financial, openness, market structure and private credit: An empirical investigation. *Economics Letters* 121, 478-481.
- Galor, O., Moav, O., 2004. From Physical to Human Capital Accumulation: Inequality and the Process of Development. *Review of Economic Studies* 71(4), 1001-1026.
- Galor, O., Zeira, J., 1993. Income Distribution and Macroeconomics. *Review of Economic Studies* 60(1), 35-52.
- Gimet, C., Lagoarde-Segot, T., 2011. A Closer Look at Financial Development and Income Distribution. *Journal of Banking and Finance* 35(7), 1698-1713.
- Greenwood, J., Jovanovic, B., 1990. Financial Development, Growth, and the Distribution of Income. *Journal of Political Economy* 98 (5), 1076-1107.
- Hakkio, C.S., Keeton, W.R., 2009. What is it, how can it be measured, and why does it matter? Federal Reserve Bank of Kansas City, Economic Review, Second Quarter.
- Hamilton, J.D., 1983. Oil and the Macroeconomy Since World War II. *Journal of Political Economy* 91, 228-248.
- Hamori, S., Hashiguchi, Y., 2012. The effect of financial deepening on inequality: Some international evidence. *Journal of Asian Economics* 23(4), 353-359.
- Hauner, D., Prati, A., Bircan, C., 2013. The interest group theory of financial development: evidence from regulation. *Journal of Banking and Finance* 37, 895–906.
- Henderson, B.J., Pearson, N.D., Wang, L., 2012. New evidence on the financialization of commodity markets. Working Paper, George Town University of Illinois.

- Henry, P.B., 2000. Stock market liberalization, economic reform and emerging market equity prices. *Journal of Finance* 55 (2), 529-564.
- Hermes, N., 2014. Does microfinance affect income inequality? *Applied Economics* 46(9), 1021-1034
- Hooker, M.A., 1996. What Happened to the Oil Price-Macroeconomy Relationship? *Journal of Monetary Economics* 38, 195-213.
- Huang, R.D., Masulis, R.W., Stoll, H.R., 1996. Energy shocks and financial markets. *Journal of Futures Markets* 16, 1-27.
- IMF, 2004. Are credit booms in emerging markets a concern? In *World Economic Outlook*, April 2004, Washington DC: International Monetary Fund.
- Inoue, T., Hamori, S., 2013. How has financial deepening affected poverty reduction in India? Empirical analysis using state-level panel data. *Applied Financial Economics* 22(5), 395-408.
- Inoue, T., Hamori, S., 2013. How has financial deepening affected poverty reduction in India? Empirical analysis using state-level panel data. *Applied Financial Economics* 22(5), 395-408.
- Jammazi, R., Aloui, C., 2010. Wavelet decomposition and regime shifts: Assessing the effects of crude oil shocks on stock market returns. *Energy Policy* 38, 1415-1435.
- Jayarathne, J., Strahan, P., 1996. The Finance-Growth Nexus: Evidence from Bank Branch Deregulation. *Quarterly Journal of Economics* 111, 639-670.
- Kaminsky, G.L., Reinhart, C.M., 1999. The Twin Crises: The Causes of Banking and Balance of Payments Problems. *American Economic Review* 89(3), 473-500.
- Kang, W., Ratti, R.A., 2013. Oil shocks, policy uncertainty and stock market return. *Journal of International Financial Markets, Institutions and Money* 26, 305-318
- 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., Park, C., 2009. The impact of oil price shocks on the U.S. stock market. *International Economic Review* 50, 1267-1287.
- Kilian, L., Vega, C., 2011. Do energy prices respond to U.S. macroeconomic news? A test of the hypothesis of predetermined energy prices. *Review of Economics and Statistics* 93, 660-671.
- Kim, D., Lin, S., 2011. Nonlinearity in the financial development-income inequality nexus.

- Journal of Comparative Economics 39(3), 310-325.
- Kim, D.H., Lin, S.C., Suen, Y.B., 2010. Dynamic effects of trade openness on financial development. *Economic Modelling* 27, 254-261.
- Kim, D.H., Lin, S.C., Suen, Y.B., 2012. Dynamic Effects of Financial Openness on Economic Growth and Macroeconomic Uncertainty. *Emerging Markets Finance and Trade* 48(1), 25-54.
- King, R.G., Levine, R., 1993. Finance and Growth: Schumpeter Might Be Right. *Quarterly Journal of Economics* 108(3), 717-737.
- King, R.G., Levine, R., 1993. Finance, Entrepreneurship, and Growth: Theory and Evidence. *Journal of Monetary Economics* 32(3), 513-542.
- King, R.G., Levine, R., 1993. Financial Intermediation and Economic Development, in Mayer, C., Vives, X.(ed.), *Capital Markets and Financial Intermediation*. London: Centre for Economic Policy Research, 156-189.
- Klein, M.W., Olivei, G.P., 2008. Capital account liberalization, financial depth, and economic growth. *Journal of International Money and Finance* 27, 861-875.
- Laeven, L., Valencia, F., 2012. Systemic Banking Crises Database: An Update. IMF Working Papers 12/163, Washington DC: International Monetary Fund.
- Lane, P. R., Milesi-Ferretti, G.M., 2007. The external wealth of nations mark II: Revised and extended estimates of foreign assets and liabilities.1970-2004. *Journal of International Economics* 73(2), 223-250.
- Law, S.H., Azman-Saini, W.N.W., 2012. Institutional Quality, Governance and Financial Development. *Economics of Governance* 13(3), 217-236.
- Levine, R., 2005. Finance and Growth: Theory and Evidence. In *Handbook of Economic Growth*, Aghion, P., Durlauf, S. (eds.), Amsterdam: North-Holland Elsevier Publishers.
- Levine, R., Loayza, N., Beck, T., 2000. Financial intermediation and growth: causality and causes. *Journal of Monetary Economics* 46, 31–77.
- Loayza, N., Rancière, R., 2006. Financial development, financial fragility, and growth. *Journal of Money Credit and Banking* 38(4), 1051-1076.
- MacKinnon, J.G., 2006. Bootstrap methods in econometrics. *Economic Record* 82, 2-18.
- Morana, C., 2013. Oil price dynamics, macro-finance interactions and the role of financial speculations. *Journal of Banking & Finance* 37, 206-226.
- Nissanke, M., 2012. Commodity market linkages in the global financial crisis: excess

- volatility and development impacts. *Journal of Development Studies* 48, 732-750.
- Park, J., Ratti, R.A., 2008. Oil price shocks and stock markets in the U.S. and 13 European countries. *Energy Economics* 30, 2587-2608.
- Pesaran, M. H., Shin, Y., Smith, R., 1999. Pooled mean group estimator of dynamic heterogeneous panels. *Journal of the American Statistical Association* 94, 621-634.
- Pesaran, M. H., Smith, R., 1995. Estimating Long-Run Relationships from Dynamic Heterogeneous Panels. *Journal of Econometrics* 68(1), 79-113.
- Prasad, E., Rogoff, K., Wei, S.J., Kose, M.A., 2003. Effects of financial globalization on developing countries: some empirical evidence. IMF Occasional Paper No.220, Washington, DC.
- Quinn, D., Schindler, M., Toyoda, A.M., 2011. Assessing measures of financial openness and Integration. *IMF Economic Review* 59(3), 488-522.
- Rajan, R. G., Zingales, L., 2003. The Great Reversals: The Politics of Financial Development in the Twentieth Century. *Journal of Financial Economics* 69, 5-50.
- Ravallion, M., 2001. Growth, inequality, and poverty: looking beyond averages. *World Development* 29, 1803-1815.
- Rodrik, D., 2008. Second-Best Institutions. *American Economic Review* 98(2), 100-104.
- Rodrik, D., Subramanian, A., 2004. From 'Hindu Growth' to Productivity Surge: The Mystery of the Indian Growth Transition. NBER Working Paper No. 10376.
- Roine, J., Vlachos, J., Waldenström, D., 2009. The long-run determinants of inequality: What can we learn from top income data? *Journal of Public Economics* 93(7-8), 974-988.
- Schaeck, K., Cihák, M., 2010. Competition, Efficiency, and Soundness in Banking: An Industrial Organization Perspective. Discussion Paper 2010-68S, Tilburg University, Center for Economic Research.
- Singelton, K.J., 2012. Investor flows and the 2008 boom/bust in oil prices. Working Paper, Stanford Graduate School of Business.
- SSRN: <http://ssrn.com/abstract=2201600>
- Stulz, R., 1999. Globalization of Capital Markets and the Cost of Capital. *Journal of Applied Corporate Finance* 12(3), 8-25.
- Tang, K., Xiong, W., 2012. Index investing and the financialization of commodities. *Financial Analysts Journal* 68, 54-74.

- Wagle, U.R., 2007. Are Economic Liberalization and Equality Compatible? Evidence from South Asia. *World Development* 35(11), 1836-1857.
- Welch, B.L., 1947. The generalization of “Student's” problem when several different population variances are involved. *Biometrika* 34(1-2), 28-35.
- Yoshizaki, Y., Hamori, S., 2013. On the influence of oil price shocks on economic activity, inflation, and exchange rates. *International Journal of Financial Research* 4, 33-41.

List of first appearances

Chapter 1

Chen, W., Hamori, S. and Kinkyo, T. (2014) Macroeconomic Impacts of Oil Prices and Underlying Financial Shocks, *Journal of International Financial Markets, Institutions and Money*, Vol.29, pp.1-12.

Chapter 2

Chen, W. and Kinkyo, T. (2016) Financial Development and Income Inequality: Long-Run Relationship and Short-Run Heterogeneity, *Emerging Markets Finance and Trade*, Vol. 52, Issue 3, pp.733-742.

Chapter 3

Chen, W., Hamori, S. and Kinkyo, T. (2016) Financial Development and Financial Openness Nexus: The Precondition of Banking Competition, *Applied Economics*, Vol.48, pp. 1130-1139.