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The Dynamic Relationship between Stock Market Development and Economic Activity: Evidence from Peru, 1965-2011.

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Abstract

We use real GDP per capita and three standard indicators of stock market development: value traded/GDP, market capitalization/GDP and turnover to study the short-run link between the stock market and economic activity in Peru. Based on annual time series data for the period 1965-2011, we estimate vector autoregressions (VARs) and identify approximate measures of stock market shocks using long-run restrictions. The results can be summarized as follows: (i) stock market indicators contribute to predict real GDP per capita growth only since the early 1990's; (ii) a stock market shock has significant short-run effects on real GDP per capita; however, its contribution to output dynamics has been small.The results imply that policy actions aimed at further developing the Peruvian stock market do have a significant positive impact on the dynamics of economic growth.

JEL Classification: E23, G1
Keywords: Output growth, stock market, VAR, long-run restrictions.

Resumen

Se estudia el rol del mercado bursátil peruano en la dinámica de la actividad económica. Se utiliza el PBI real per cápita y tres indicadores de desarrollo del mercado bursátil: el ratio de valor transado sobre el PBI, la capitalización bursátil sobre el PBI y el ratio de rotación. Se estima un modelo VAR sobre la base de series anuales para el periodo 1965-2011 y se identifica los choques del mercado bursátil por medio de restricciones de largo plazo. Los resultados se resumen como sigue: (i) los indicadores bursátiles contribuyen en la predicción del crecimiento del PBI per cápita sólo desde inicios de los 90's; (ii) un choque en el mercado bursátil tiene efectos significativos de corto plazo sobre el PBI, sin embargo su contribución sobre la volatilidad del producto ha sido pequeña. Los resultados implican que las acciones de política que busquen desarrollar más el mercado bursátil peruano tendrían un efecto significativo y positivo sobre la dinámica de la actividad económica.

Claves: Dinámica del producto, mercado bursátil, VAR, restricciones de largo plazo.

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1 Introduction

How important are stock markets in the dynamics of economic growth? Do they have any causal effect or do they just help to predict economic growth? In this paper we intend to provide answers to these questions by exploring the empirical relationship between stock markets and real economic activity using time series data. We analyse the case of Peru, a small open economy with a stock market in an early stage of development.

Since Smith (1776), the relationship between the financial system and real activity has been widely studied from a theoretical and empirical point of view, as documented by Gertler (1988), King and Levine (1993), Becsi and Wang (1997), Levine (2005), among others. Although there is still no consensus on the causal relationship, it is well accepted in the literature that stock markets and banking systems allow societies to optimally channel resources from savings towards consumption and productive activities. Huge adverse financial events such as the Great Depression, the Asian and Russian crisis at the end of the last century, the recent global financial crisis and the subsequent Great Recession, have all shown that the link between the financial system and real activity is very important in practice. In particular, a malfunctioning financial system can have deep deleterious effects on the level of output and economic growth.

Although the main emphasis has been on banks (e.g. Bagehot (1873) and Schumpeter (1912), among the most important), stock markets have been acknowledged as an important force in the economy especially since the Great Depression,[1] and, recently, after the global financial crisis. On the theoretical side, Levine (1991), Devereux and Smith (1994), and Obstfeld (1994) provide models in which liquid and internationally integrated stock markets may contribute to economic growth. Empirically, the relationship between stock markets and real activity has been studied using cross-sectional data (Levine and Zervos, 1998; Cooray, 2010, and references therein), panel data (Henry, 2000; Beck and Levine, 2004; Gupta and Yuan, 2009, and references therein), and time series data (Arestis et al., 2001; Caporale at al., 2005; Enisan and Olufisayo, 2009; Marques et al., 2013, and references therein). In the latter case, the literature has focused on the empirical causality, i.e. time precedence, based on Granger causality tests and VAR analysis. However, efforts to identify the possible causal effect of stock market indicators on real activity have been scant in time series studies.

In this paper, we go beyond the study of empirical causality and identify the possible causal effect of stock markets on real economic activity. This is important from a policy perspective because, as stated by Cochrane (1994), only responses to an exogenous variable can measure the effects of policy-induced changes in that variable. Based on a VAR approach with stationary series (first-differences), we propose the identification of an approximate measure of stock market shocks using long-run restrictions, as proposed by Blanchard and Quah (1989) and applied by Quah and Vahey (1995) in the monetary-macro literature. In particular, we identify one structural shock that has no long-run effect on real GDP per capita and interpret it as a stock market shock based on

[1] According to Fisher (1933), the economic contraction experienced after 1929 resulted from a high level of borrower's leverage, caused by malfunctioning financial markets

the variance decomposition analysis and the particular features of the Peruvian economy.

We employ annual data for Peru through the period 1965-2011. Specifically, we use real GDP per capita[2] and three conventional indicators of stock market development: volume traded to GDP ratio, stock market capitalization to GDP ratio, and the turnover ratio. The empirical analysis shows that the dynamic relationship between real GDP per capita and stock market in Peru has changed over time and that the stock market has become more important since 1991. We find two main results: (i) stock market indicators contribute to the prediction of real GDP per capita growth only since the early 1990's; and (ii) stock market shocks have had significant short-run effects on real GDP per capita during the period 1991-2011. In particular, a one-standard deviation shock to value traded/GDP, turnover and capitalization/GDP increases real GDP per capita after one year by 1%, 1.4% and 1.0%, respectively; however, its contribution to output dynamics has been small. Therefore, the results imply that policy actions aimed at further developing the Peruvian stock market (e.g. promoting a higher participation of both lenders and borrowers or reducing transaction costs) may have a significant positive impact on the dynamics of economic growth.

The paper is divided into five sections. Section 2 describes the data and the key stylised facts related to the Peruvian stock market. Section 3 describes the empirical methodology. Section 4 provides and discusses the econometric results. Section 5 concludes.

2 Data and stylized facts

The data frequency is annual and covers the period 1965-2011. We use three conventional indicators of stock market development: (i) value traded to GDP ratio, (ii) turnover ratio, and (iii) stock market capitalization to GDP ratio. These data were collected from the National Commission of Securities (CONASEV)'s printed Annual Reports and its online database.[3] We use real GDP per capita to measure real economic activity, so that the growth rate is an approximate measure of welfare. Additionally, we include the bank credit to GDP ratio and the terms of trade. All these macroeconomic data were obtained from the Central Reserve Bank of Peru's online database.

Value traded/GDP and turnover are indicators of stock market liquidity. Value traded is defined as the value of shares traded in the stock market, whereas turnover is defined as the percentage of traded shares relative to total shares valued in the stock market. In particular, a higher value traded/GDP or turnover implies a more liquid stock market which provides potential benefits for real economic activity and growth. Theoretically, illiquid markets could prevent long-run investments because it is difficult to sell shares whenever an investor needs liquidity. However, a lower turnover could also indicate that investors have a long-term investment horizon (e.g. insurance companies and private pension funds), which could contribute to economic growth. Stock mar-

[2]This indicator can be considered as an average measure of welfare.

[3]Since 2012 CONASEV changed its name to Superintendency of Security Markets (SMV).

ket capitalization is defined as the value of shares that are available in the Lima Stock Market. Thus, a higher capitalization/GDP is usually interpreted as an indicator of a bigger and more developed stock market; however, a higher number shares listed does not necessarily affect real activity and growth.

Table 1 shows the evolution of the Peruvian stock market[4] measured by the average value traded/GDP, capitalization/GDP and turnover for different samples. In average terms, value traded shows an important improvement since 1991, in contrast to the reduction observed in turnover during the same period. Stock market capitalization also shows an important improvement between 1991-2000 and 2001-2011.

Table 1. *Average levels of stock market indicators (% of GDP)*

	Value traded	**Turnover**	**Capitalization**
1965-1970	0.16	42.51	n.a.
1971-1980	0.43	82.80	n.a.
1981-1990	0.49	69.57	n.a.
1991-2000	8.39	33.97	19.46
2001-2011	5.47	6.82	54.93

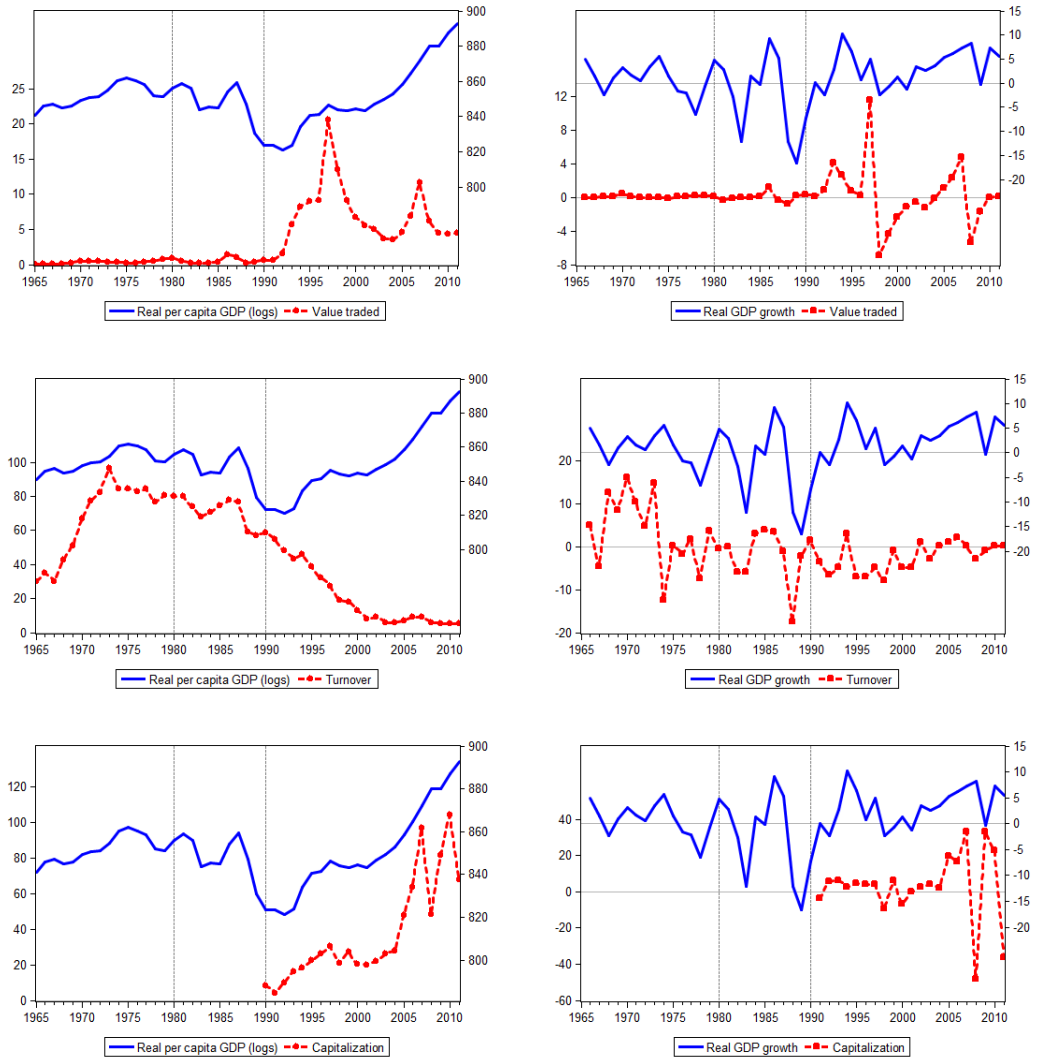
However, compared to developed economies like the United Kingdom (UK) and the United States of America (US), it is evident that the Peruvian stock market is at an early stage of development. For instance, according to the Global Financial Development Database, during the period 2001-2011: (i) capitalization/GDP in the UK and US were 127% and 123%, respectively, whereas in Peru it was 44%, (ii) value traded/GDP in UK and US were 171% and 255%, respectively, whereas in Peru it was 2.8%, and (iii) turnover in UK and US were 138% and 208%, respectively, whereas in Peru it was 6.75%.

Figure 1 shows the evolution of stock market indicators along with real GDP per capita (in logs), both in levels and in first differences. It is possible to distinguish two periods in the evolution of real GDP per capita: (i) a period of stagnation between 1965 and 1990, featuring important drops in 1983 and 1988-1989 as a consequence of a weather phenomenon called "El Niño" and the Peruvian hyperinflation period, respectively; and (ii) a period of economic recovery since 1991, interrupted between 1998 and 2001 (which coincides with international crises in Asia, Russia, Brazil, and a period of domestic political instability between 2000 and 2001) and in 2009 (as a result of the international financial crisis).

The three stock market indicators under consideration show different performances. The establishment of the National Commission for Companies and Securities (CONA-

[4]The Peruvian stock market was established on December 31st, 1860 (during the government of Ramon Castilla), under the name of Commerce Stock in Lima, and began to operate on January 7th, 1861. After the Great Depression and the Second War World, the New Commerce Stock of Lima was created in 1951. The current Stock Exchange Market of Lima was established in 1970, the same year that the National Commission of Securities (CONASEV) began to operate.

Figure 1. *Stock market indicators and output per capita: levels and first-differences.*



NOTE: *Figures in the first column illustrate the levels of the series, whereas the ones in the second column illustrate the first difference of the series. Real GDP per capita is measured on the right axis and is expressed in logarithms and multiplied by 100, so the first difference is expressed in percentages. Stock market indicators are measured on the left axis and are expressed in percentages, so its first difference is measured in percentage points.*

SEV) and the Stock Market Regulation Law in 1970 contributed to increasing the number of shares negotiated in the Lima Stock Exchange during the 1970's and 1980's. The stock market liquidity, as measured by value traded/GDP, showed a stable performance up to 1990 although, featuring an important drop in 1988 that coincides with the beginning of the hyperinflation period. Even though stock market liquidity improved, its average level was 0.49% of GDP between 1980 and 1990, a very low level when compared with other Latin American countries (by 1990, Chile registered 2.5%, Brazil 1.2%, and Mexico 4.6%). After the macroeconomic stabilization programme implemented in 1990, which included the financial liberalization of the economy, stock market liquidity

started to increase, reaching levels of around 10% of GDP. However, it has followed an irregular path since 1997, featuring an important increase in 1993 when the Private Pensions Fund Administrators (AFPs) started to operate, and two important drops between 1997-1998 (Asian crisis) and 2007-2008 (global financial crisis), respectively. Turnover suggests a different evolution of liquidity. It shows a downward trend since the beginning of 1970's, which is consistent with real GDP per capita performance up to 1991; however, according to this indicator, stock market liquidity has kept the same downward path despite the recovery in output. Given that this period coincides with the advent of the AFPs, which collectively have turned out to be the most important investor in the stock market, low stock market liquidity can be explained by the fact that AFPs tend to hold more long-run positions. Also, the important limits to short selling that exist in the Lima stock market (Diaz-Martinez and Fragniere, 2012), might be associated with its low level of liquidity.[5]

Finally, the size of the market, as measured by stock market capitalization to GDP ratio, has shown an upward trend for the available sample 1990-2011. This performance is consistent with the financial liberalization of the economy that began in 1990 and the evolution of real GDP per capita, which led to a peak in capitalization/GDP in 2007 and its subsequent recovery after a major drop in 2008 during the peak of the global financial crisis.

3 Empirical model

We use a vector autoregression (VAR) to analyse the dynamic relationship between real economic activity and stock markets. Specifically, we consider the following bivariate ($n = 2$) VAR model of order p expressed in first-differences:

$$\Delta y_t = a_{10} + a_{11}^{(1)}\Delta y_{t-1} + a_{12}^{(1)}\Delta f_{t-1} + \cdots + a_{11}^{(p)}\Delta y_{t-p} + a_{12}^{(p)}\Delta f_{t-p} + d_t'\alpha^y + \varepsilon_t^y \quad (1)$$

$$\Delta f_t = a_{20} + a_{21}^{(1)}\Delta y_{t-1} + a_{22}^{(1)}\Delta f_{t-1} + \cdots + a_{21}^{(p)}\Delta y_{t-p} + a_{22}^{(p)}\Delta f_{t-p} + d_t'\alpha^f + \varepsilon_t^f \quad (2)$$

where f_t is a real indicator of stock market development, y_t is the log of real per capita output, and the errors are normal white noise processes, that is, $\varepsilon_t^j \sim (0, \sigma_{\varepsilon^j}^2)$ and $Cov(\varepsilon_t^j, \varepsilon_s^j) = 0$ for $j = y, f$. However, the errors between equations are assumed to be correlated, that is, $Cov(\varepsilon_t^y, \varepsilon_t^f) \neq 0$. The row vector d_t' contains dummy variables that account for specific events such as "El Niño" phenomenon and others described below. Based on Hamilton (1994, p. 651-654), this specification is econometrically adequate given that y_t and each f_t are unit root processes[6] and do not cointegrate.[7]

[5]Evidence provided for example by Daouk and Charoenrook (2005) suggests that absence of short selling is linked to illiquid markets.

[6]The unit root hypothesis cannot be rejected using the Augmented-Dickey Fuller (ADF) test and the efficient unit root tests proposed by Elliot, Rothenberg and Stock (1996). The results are available upon request.

[7]The null hypothesis of no cointegration cannot be rejected using Johansen's rank test

3.1 Identification of shocks

Under the VAR framework, it is possible to analyze the dynamic relationship between f_t and y_t in terms of impulse response functions (IRFs) and variance decomposition (VD) analysis. However, IRFs and VD require the identification of orthogonalized errors which represent "innovations" in a particular variable, i.e. movements that cannot be anticipated given a particular information set.

The Cholesky decomposition of the variance-covariance matrix of the residuals is a typical procedure to obtain orthogonalized errors, u_t^y and u_t^f. This method implies restrictions on the contemporaneous response of the variables to each orthogonal shock, which are known as "Cholesky ordering". For instance, a Cholesky ordering $(\Delta y_t, \Delta f_t)$ means that Δy_t responds contemporaneously to surprises in Δy_t only, whereas Δf_t responds contemporaneously to surprises in both Δy_t and Δf_t.

Under the Cholesky decomposition, the IRFs describe the dynamic response of Δy_t and Δf_t to an orthogonalized shock in either Δy_t or Δf_t. In particular, the response of Δy_t to an orthogonalized shock in Δf_t can be interpreted as the change in the forecast of Δy_t due to "new information" about Δf_t. On the other hand, VD analysis provides the contribution of each orthogonalized shock, u_t^y and u_t^f, to the fluctuations in Δy_t and Δf_t. Furthermore, if Δf_t (Δy_t) were mostly explained by their own shocks u_t^f (u_t^y), then Δf_t (Δy_t) can be considered as relatively exogenous.

One important limitation of the Cholesky method is that results from IRFs and VD depend on the Cholesky ordering. However, if one particular ordering is "reasonable", then at least one of the orthogonalized shocks can be interpreted as a structural or primitive shock, i.e. a shock whose true origin could be known conditional on the VAR specification.[8] However, in our case there does not seem to exist a unique reasonable Cholesky ordering: real activity might respond to a stock market shock within the same year the shock occurs, whereas it is almost certain that the stock market will react immediately to shocks in the real sector.

3.2 Long-run restrictions

As discussed in Cochrane (1994), we recognize the identification of the source of shocks is not an easy task. However, we believe that it is still possible to get a reasonable approximation of a particular shock using the information available in a VAR. In particular, we propose to combine the information of the effects of a shock identified using long-run restrictions and the information provided by the variance decomposition analysis. First, consider the moving-average representation of the stationary sequences $\{\Delta y_t\}$ and $\{\Delta f_t\}$ (omitting intercepts) written as follows:

$$\begin{bmatrix} \Delta y_t \\ \Delta f_t \end{bmatrix} = \begin{bmatrix} C_{11}(L) & C_{12}(L) \\ C_{21}(L) & C_{22}(L) \end{bmatrix} \begin{bmatrix} u_t^1 \\ u_t^2 \end{bmatrix}$$

[8] This is the case, for example, in monetary policy analysis, where interest rate is ordered last. See for example Bernanke et al. (2005)

where, for example, $C_{12}(L) = c_{21}(0)L^0 + c_{21}(1)L^1 + c_{21}(2)L^2 + \cdots$. For simplicity, the covariance matrix of structural errors is assumed to be represented by the identity matrix, $\Sigma_s = I$. Following Blanchard and Quah (1989) and Quah and Vahey (1995), we consider an identification scheme based on long-run restrictions. Specifically, we assume that u_t^2 does not have long run effects on real output per capita, which means that coefficients $c_{12}(i)$ should satisfy:[9]

$$C_{12}(1) \equiv \sum_{i=0}^{\infty} c_{12}(i) = 0 \tag{3}$$

The remaining terms are left unrestricted. This means that u_t^2 may have long-run effects on any stock market indicator (i.e., $C_{22}(1) \neq 0$), and that u_t^1 may have permanent effects on both stock market indicators and real output per capita (i.e., $C_{11}(1) \neq 0$ and $C_{21}(1) \neq 0$). Furthermore, all short-run dynamics are left unrestricted.

Under this identification scheme, u_t^2 can be interpreted as a structural stock market shock under two assumptions. First, shocks originating in the stock market have no long-run effect on real GDP per capita. Second, it must be true that u_t^2 actually represents shocks originating in the stock market. We argue that these two assumptions are satisfied for the Peruvian case.

In a developed economy like the US, where stock markets have shown persistent effects on real economic activity, especially during the Great Depression and the recent global financial crisis, the first assumption does not seem plausible. However, an economy in which the participation of the stock market has been relatively small, as is the case in Peru, the first assumption is reasonable. As described in section 2, data from Peru show that the stock market has a relatively small participation in the economy compared to the banking sector. Furthermore, major shifts in the level of real GDP per capita have not been associated with particular events originating in the stock market, but have been closely related to adverse weather shocks (e.g. "El Niño" weather phenomenon) and external events (both real and financial), among the most important ones. Therefore, we include exogenous additive dummy variables to the VAR in order to control for these major events,[10] so that u_t^2 represents shocks that have no long-run effect on real output per capita.

However, it is difficult to argue that u_t^2 represents only shocks coming from the stock market. Instead, u_t^2 can be viewed as an "average" of all possible shocks that have no long-run effect on real economic activity but may have long-run effects on the stock market, one of which is a "true" exogenous stock market shock.[11] Therefore, u_t^2 is a

[9]The model satisfies the required assumptions for the validity of this identification approach, as in Quah and Vahey (1995): log output is integrated of order 1, $log(Y_t) \sim I(1)$, and the change in financial indicators is stationary, $\Delta f_t \sim I(0)$.

[10]The dummy variables are defined for the following events: "El Niño" phenomenon in 1983 and 1998, hyperinflation 1988-1990, the beginning of the private pension system in 1993, the 1997-1998 financial crisis, and the global financial crisis that erupted in 2008.

[11]Given the orthogonality assumption between u_t^1 and u_t^2, all shocks whose possible long-run effect on output cannot be discarded are represented by u_t^1. Thus, u_t^1 can also be viewed as an average

noisy measure of stock market shocks.

Given this, one possible strategy to get a better approximation of stock market shocks is to extend the VAR including a real indicator of the banking system, closely related to stock markets. We choose the ratio credit/GDP, b_t, which is a common banking indicator used in the literature. Thus, the moving average representation of the dynamic relationship between Δy_t, Δb_t, and Δf_t can be written as:

$$\begin{bmatrix} \Delta y_t \\ \Delta b_t \\ \Delta f_t \end{bmatrix} = \begin{bmatrix} C_{11}(L) & C_{12}(L) & C_{13}(L) \\ C_{21}(L) & C_{22}(L) & C_{23}(L) \\ C_{31}(L) & C_{32}(L) & C_{33}(L) \end{bmatrix} \begin{bmatrix} u_t^1 \\ u_t^2 \\ u_t^3 \end{bmatrix}$$

In this case, we assume that u_t^3 has no long-run effects on either Δy_t or Δb_t, i.e. $C_{13}(1) \equiv \sum_{i=0}^{\infty} c_{13}(i) = 0$ and $C_{23}(1) \equiv \sum_{i=0}^{\infty} c_{23}(i) = 0$. The restriction $C_{23}(1) \equiv 0$ is consistent with the Peruvian experience, given the small participation of the stock market in the economy compared to the banking sector, whereas $C_{13}(1) \equiv 0$ is similar to $C_{12}(1) \equiv 0$ in the two-variable VAR. In order to achieve identification, we impose the additional restriction $C_{12}(1) \equiv \sum_{i=0}^{\infty} c_{12}(i) = 0$ which states that u_t^2 represents shocks that do not have long-run effects on real output per capita. Thus, u_t^3 can be considered as a better approximation of stock market shocks.

This conjecture can be supported by the variance decomposition analysis of Δf_t. With high frequency data (daily, monthly or even quarterly), stock markets react immediately to new information, real or financial. However, with annual data it is plausible to believe that shocks that are the main drivers of stock market fluctuations must come mainly from the stock market. Therefore, if most of the variance in Δf_t is explained by innovations in u_t^3, then it is reasonable to state that u_t^3 represents stock markets shocks.

Once stock market shocks are identified, it is possible to use IRFs and variance decomposition analysis in order to quantify the importance of the Peruvian stock market in the dynamics of economic growth. We do this in the next section.

4 Results

We estimate (1) and (2) with one lag ($p = 1$) for each stock market indicator using maximum likelihood. The lag specification for each model was determined using a sequential Likelihood Ratio (LR) test, which provides VARs with normally distributed errors, homoskedastic and serially uncorrelated.

Based on the historical evolution of the Peruvian stock market described in section 2, one would expect this dynamic relationship to have become stronger over time and thus that the response of output to stock market shocks might have changed as well. Given this, we estimate the models for three sub samples: (i) 1965-2011 (full sample), (ii) 1965-1990, which considers the early development of stock markets in Peru and the

of technological, preferences, government, factor prices, and consumption shocks, among the most important. For this reason, we do not impose any "label" on u_t^1.

military government (1968-1980), the decline of state-owned development banks, the attempted nationalization of the private banking system, and the hyperinflation period that ended in 1990, and (iii) 1991-2011, which covers the period of structural macroeconomic reforms, stabilization and low inflation.

As a preliminary step, we perform bivariate Granger causality tests between the growth of real GDP per capita and the change in each stock market indicator in order to determine the direction of empirical causality. The results shown in Table 2 indicate that Granger causality has changed over time and varies with each stock market indicator. Based on all the information available for each indicator, value traded/GDP and capitalization/GDP help to predict real GDP growth per capita, whereas real GDP per capita helps to predict turnover. In the case of value traded/GDP, the direction of causality reverses: real GDP growth contributes to the prediction of value traded/GDP in the period 1965-1990, whereas value traded/GDP causes real GDP growth in the period 1991-2011. Although turnover causes real GDP growth in the period 1991-2011, there is no evidence of any causal relationship between them in the sub-sample 1965-1990.

Table 2. *Bivariate Granger causality tests.*

Indicator	**1965-2011**	**1965-1990**	**1991-2011**
$\Delta VT \nRightarrow \Delta y$	0.09	0.37	0.01
$\Delta y \nRightarrow \Delta VT$	0.43	0.03	0.99
$\Delta T \nRightarrow \Delta y$	0.48	0.27	0.06
$\Delta y \nRightarrow \Delta T$	0.04	0.28	0.36
$\Delta C \nRightarrow \Delta y$	n.a	n.a	0.01
$\Delta y \nRightarrow \Delta C$	n.a	n.a.	0.65

Note: ΔVT, ΔT, ΔC, and Δy denote the first difference of value traded/GDP, turnover, capitalization/GDP, and the growth rate of real GDP per capita. The null hypothesis is that the left-hand side variable does not Granger cause the right-hand side variable.

As discussed in section 3, no Cholesky ordering between real GDP per capita growth and the variation in stock market indicators seems to be plausible with annual data. Thus, we propose to identify a stock market shock by imposing long-run restrictions on the impulse-response functions of the VAR. In particular, we assume that u_t^2 does not have long-run effects on real GDP per capita.[12]

The first block of Table 3 (Baseline VAR) shows the contribution of u_t^2 to the variance of real GDP per capita growth, Δy, and the first difference of each stock market indicator, Δf. On the one hand, u_t^2 explains 87.7%, 87.8%, and 54.7% of the fluctuations in Δf when value traded, turnover, and capitalization are used, respectively. Therefore, u_t^2 can be interpreted as a noisy measure of stock market shocks for the whole sample. By sub-samples, the contribution of u_t^2 is not constant. For the case of value traded, the contribution of u_t^2 remains at high levels (53.4% in 1965-1990 and 69.7% in 1991-2011);

[12]The impulse-response functions based on Cholesky decomposition are shown in Figure A-1, Appendix A.

however, with turnover its contribution drops to 30.6% during the period 1991-2011.

The overall contribution of u_t^2 to the variance of Δy is small and is not constant for all the samples considered, but shows an increasing importance in recent years. u_t^2 explains less that 15% of the fluctuations in Δy when value traded is used: 1.7% for the full sample, 5.5% for the 1965-1990 sample, and 12.9% for the 1991-2011 sample. In the case of turnover, the contribution of u_t^2 to the dynamics of economic growth reaches a maximum level of 21.7% for the sample 1991-2011, whereas for capitalization it is 15.7%.

Figure 2 shows the cumulative response of real GDP per capita to a stock market shock (one-standard deviation), based on the long-run identification procedure. Rows 1, 2 and 3 of Figure 2 contain the IRFs for samples 1965-2011, 1980-2011 and 1991-2011, respectively.[13] Unlike the VAR with orthogonalized Cholesky errors, the structural IRFs show that the response of real GDP per capita to a stock market shock has changed over time and has become more important and significant in recent years. In particular, the maximum effect of a stock market shock is achieved after one year and is equal to 1.0% for value traded, 1.4% for turnover, and 1.0% for capitalization.

In order to get a better approximation of stock market shocks, we also estimate VARs including banking credit/GDP, a real indicator of banking sector development. Based on the second block of Table 3 (VAR with bank credit), and following the same reasoning as before, we conclude that u_t^3 is a reasonable approximation of stock market shocks except when turnover is used. In particular, 52.8% of the fluctuations in capitalization are explained by shocks to u_t^3, whereas the contribution of u_t^3 to the fluctuations in value traded/GDP is 87.3% for the whole sample and 54.1% for the period 1991-2011. In contrast, the contribution of u_t^3 to the fluctuations in turnover falls from 63.7% (full-sample) to 36.6% (1965-1990) and 10.5% (1991-2011). Compared to the results obtained from the two-variable VAR, the contribution of stock market shocks to fluctuations in Δy has also increased in recent years, reaching 10.6%, 13.7% and 15.9% for value traded, turnover and capitalization, respectively.

Figure 3 shows the dynamic response of real GDP per capita to a one-standard deviation shock to the stock market. As in the previous case, a stock market shock is significant only in the period 1991-2011. In all cases, the maximum effect is achieved one year after the shock occurs and is equal to 1.0% for value traded/GDP, 1.1% for turnover, and 1.0% for capitalization.

Finally, as a robustness exercise, we extend the original VAR with the log of terms of trade in order to control for other possible macroeconomic sources of shocks. The variance decomposition analysis for this alternative three-variable model is shown in the third block of Table 3 (VAR with terms of trade). Although fluctuations in value traded/GDP and turnover are mainly explained by u_t^3, this contribution is reduced by sub-samples. In the case of capitalization, only 22.8% of the fluctuations in real GDP growth is explained by u_t^3. Therefore, u_t^3 can still represent stock market shocks when either value traded/GDP (for the sample 1991-2011) or turnover (for the sample 1965-

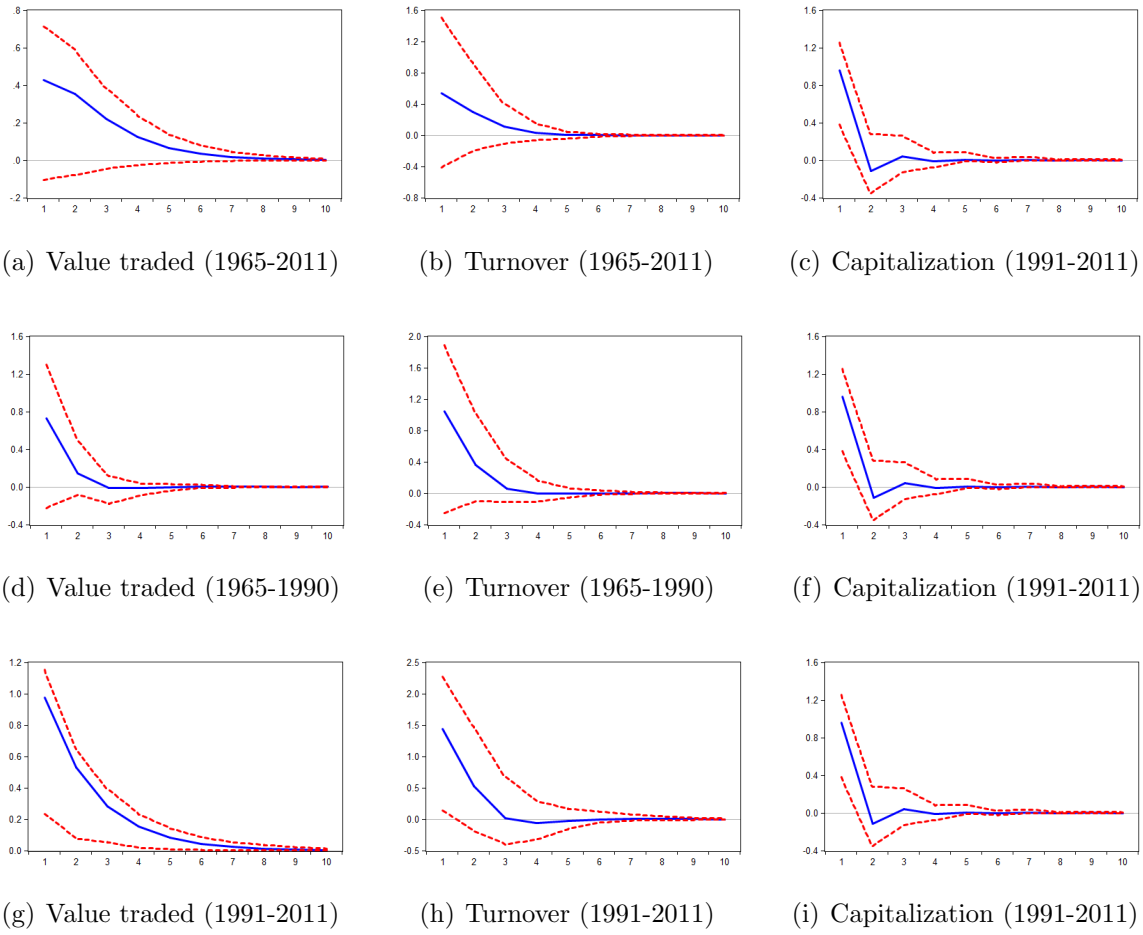
[13]Data for market capitalization are only available from 1991, after considering one lag in the VAR.

Table 3. *Contribution of stock-market shocks (u_t^3) to the variance of output growth and stock markets indicators.*

Sample	Forecast horizon	Value traded		Turnover		Capitalization	
		Δy	Δf	Δy	Δf	Δy	Δf
Baseline VAR							
Sample	1 year	1.5	85.2	2.3	89.7	n.a.	n.a.
1965-2011	2 years	1.4	87.0	2.4	88.8	n.a.	n.a.
	10 years	1.7	87.7	2.7	87.8	n.a.	n.a.
Sample	1 year	4.0	59.4	8.4	75.9	n.a.	n.a.
1965-1990	2 years	5.4	53.8	10.0	75.7	n.a.	n.a.
	10 years	5.5	53.4	10.5	75.4	n.a.	n.a.
Sample	1 year	10.5	71.2	18.7	25.2	10.3	58.1
1991-2011	2 years	12.2	70.0	20.3	30.6	15.6	54.8
	10 years	12.9	69.7	21.7	30.6	15.7	54.7
VAR with bank credit							
Sample	1 year	1.0	88.1	8.1	63.1	n.a.	n.a.
1965-2011	2 years	1.1	88.3	6.5	65.4	n.a.	n.a.
	10 years	1.3	87.3	7.6	63.7	n.a.	n.a.
Sample	1 year	0.3	66.7	21.4	33.6	n.a.	n.a.
1965-1990	2 years	0.9	27.2	16.2	37.2	n.a.	n.a.
	10 years	0.9	27.0	17.5	36.6	n.a.	n.a.
Sample	1 year	8.8	60.2	10.0	10.0	10.7	56.4
1991-2011	2 years	10.1	55.6	13.8	10.7	15.8	52.9
	10 years	10.6	54.1	13.7	10.5	15.9	52.8
VAR with terms of trade							
Sample	1 year	0.5	83.0	0.5	86.9	n.a.	n.a.
1965-2011	2 years	0.6	84.2	0.7	85.3	n.a.	n.a.
	10 years	0.6	85.0	0.8	83.3	n.a.	n.a.
Sample	1 year	0.0	40.0	1.2	69.0	n.a.	n.a.
1965-1990	2 years	0.1	36.7	2.4	66.4	n.a.	n.a.
	10 years	0.1	36.1	2.4	65.7	n.a.	n.a.
Sample	1 year	0.4	49.2	15.2	16.0	13.8	28.1
1991-2011	2 years	0.4	53.0	20.0	20.1	18.1	22.9
	10 years	0.5	54.8	20.6	20.1	16.9	22.8

Note: The results are based on the estimation of three variable VARs that includes the growth rate of real output per capita and stock market indicators. The model with bank credit considers bank credit to GDP ratio. The model with terms of trade considers the growth rate of terms of trade.

Figure 2. *Cumulative output response to a positive stock market shock: long-run restrictions.*



(a) Value traded (1965-2011) (b) Turnover (1965-2011) (c) Capitalization (1991-2011)

(d) Value traded (1965-1990) (e) Turnover (1965-1990) (f) Capitalization (1991-2011)

(g) Value traded (1991-2011) (h) Turnover (1991-2011) (i) Capitalization (1991-2011)

NOTE: *The shock size is one standard deviation of the structural residual. The vertical axis is measured in percent. Market capitalization data are available only from 1991. The impulse response functions display bootstrapped 90% confidence intervals.*

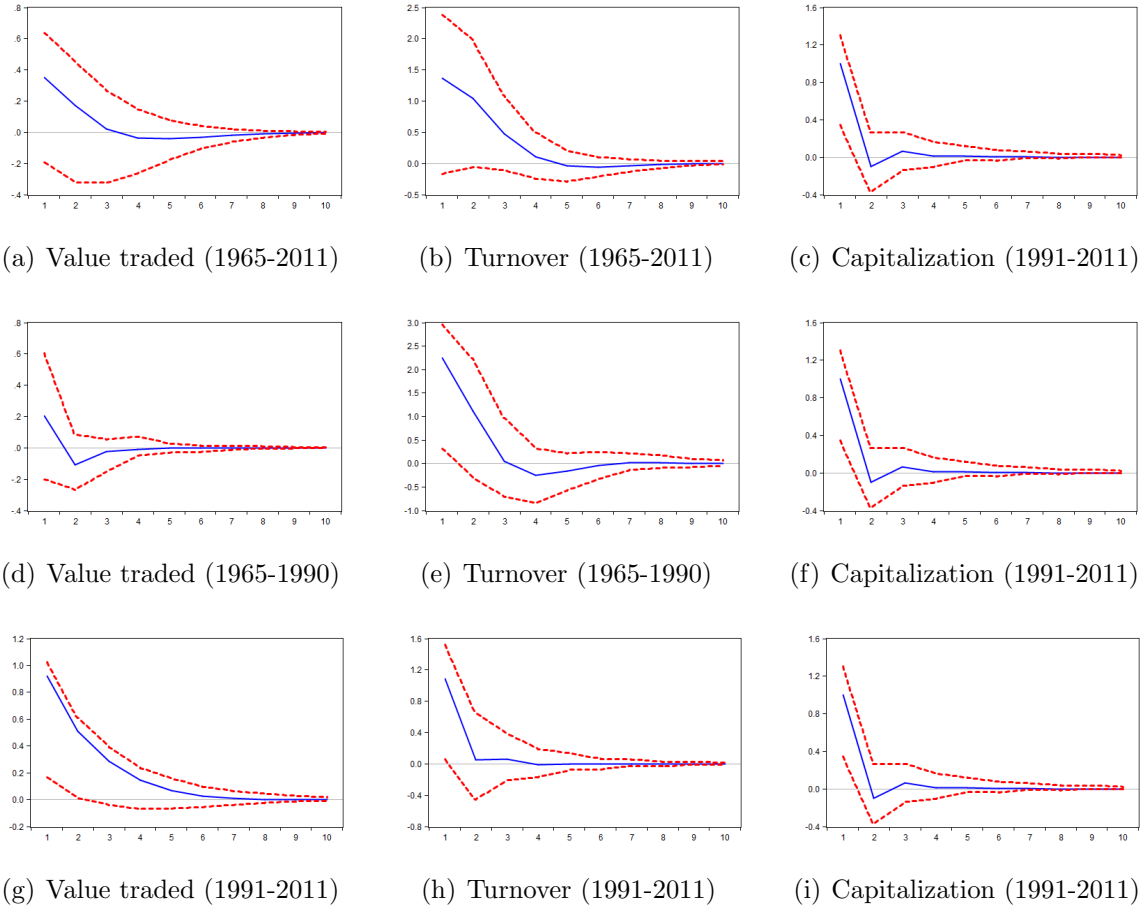
2011) are used. However, the contribution of stock market shocks to explain fluctuations in real GDP per capita growth is small (less that 1.0% in all the cases where u_t^3 can be interpreted as a stock market shock). Figure B-1 in Appendix B shows the dynamic response of real GDP per capita to shocks in u_t^3.

5 Conclusions

Using annual time series data for 1965-2011, we estimate a vector autoregression (VAR) and identify stock market shocks using long-run restrictions. We use GDP per capita and three financial indicators associated with stock markets: value traded/GDP, stock market capitalization/GDP and turnover ratio.

The empirical analysis shows that the dynamic relationship between real GDP per capita and the stock market in Peru has changed over time and that the role of the

Figure 3. *Cumulative output response to a positive stock market shock: long-run restrictions and banking credit.*



(a) Value traded (1965-2011) (b) Turnover (1965-2011) (c) Capitalization (1991-2011)

(d) Value traded (1965-1990) (e) Turnover (1965-1990) (f) Capitalization (1991-2011)

(g) Value traded (1991-2011) (h) Turnover (1991-2011) (i) Capitalization (1991-2011)

NOTE: *The shock size is one standard deviation of the structural residual. The vertical axis is measured in percent. Market capitalization data are available only from 1991. The impulse response functions display bootstrapped 90% confidence intervals.*

Peruvian stock market has become more significant since 1991. Thus, for the sample 1991-2011, we find two main results: (i) all stock market indicators contribute to the prediction of the growth rate of real GDP per capita, and (ii) a stock market shock has significant short-run effects on the growth rate of real GDP per capita. In particular, a one-standard deviation shock to value traded/GDP, turnover and capitalization/GDP increases real GDP per capita after one year by 1%, 1.4% and 1.0%, respectively; however, its contribution to output dynamics has been small.

Overall, the results imply that policy actions aimed at further developing the Peruvian stock market (e.g. promoting a higher participation of both lenders and borrowers) will have a significant positive impact on the dynamics of economic growth. However, it is possible that above a given threshold of development, further improvements of stock markets might have adverse effects on output and growth volatility, a topic that merits additional research.

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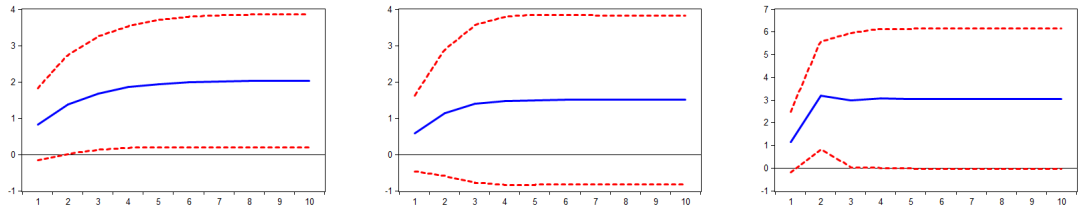
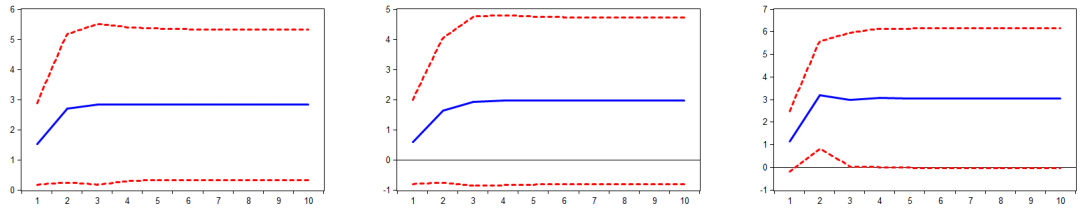
Appendix

A Impulse-response functions: Cholesky decomposition.

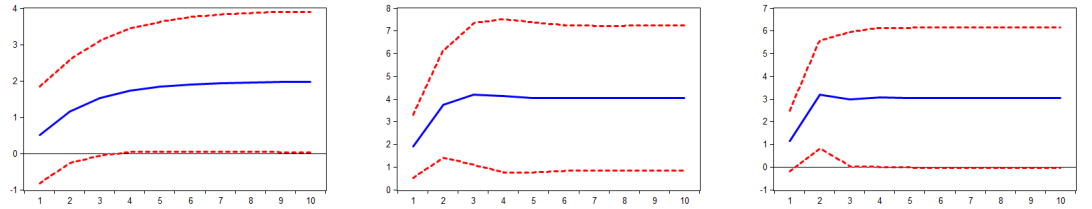
Figure A-1. *Cumulative output response to a positive stock market shock: Cholesky approach.*



(a) Value traded (1965-2011) (b) Turnover (1965-2011) (c) Capitalization (1991-2011)



(d) Value traded (1965-1990) (e) Turnover (1965-1990) (f) Capitalization (1991-2011)

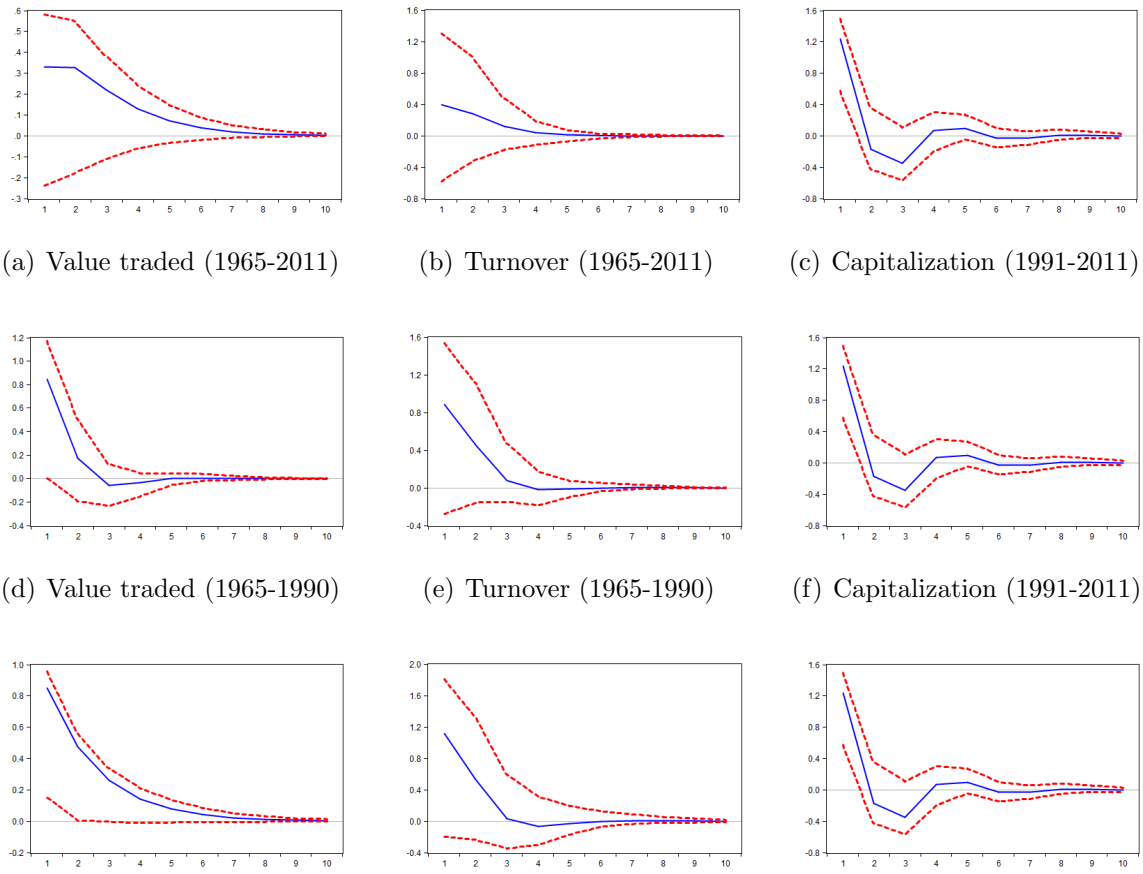


(g) Value traded (1991-2011) (h) Turnover (1991-2011) (i) Capitalization (1991-2011)

NOTE: *The shock size is one standard deviation of the orthogonalized residual. The vertical axis is measured in percent. Market capitalization data are available from 1991.*

B Impulse-response functions: terms of trade.

Figure B-1. *Cumulative output response to a positive stock market shock: long-run restrictions and terms of trade.*



(a) Value traded (1965-2011) (b) Turnover (1965-2011) (c) Capitalization (1991-2011)

(d) Value traded (1965-1990) (e) Turnover (1965-1990) (f) Capitalization (1991-2011)

(g) Value traded (1991-2011) (h) Turnover (1991-2011) (i) Capitalization (1991-2011)

NOTE: *The shock size is one standard deviation of the structural residual. The vertical axis is measured in percent. Market capitalization data are available only from 1991. The impulse response functions display bootstrapped 90% confidence intervals.*

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