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Financial openness and business cycle volatility

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Abstract

This paper discusses whether the integration of international financial markets affects business cycle volatility. In the framework of a new open economy macro-model, we show that the link between financial openness and business cycle volatility depends on the nature of the underlying shock. Empirical evidence supports this conclusion. Our results also show that the link between business cycle volatility and financial openness has not been stable over time.
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1. Motivation

The past two decades have witnessed a rather unprecedented process of deregulation of financial markets and of liberalization of cross-border capital flows. In quantitative terms, capital market integration has at least reached the levels observed during the Gold Standard. In qualitative terms, integration is now probably much

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deeper than it used to be as financial markets have increasingly become securitized. At the same time, business cycle fluctuations in OECD countries have declined.[1] These observations suggest that changes in business cycle characteristics may be related to changes in the degree of financial market integration (Basu and Taylor, 1999).

Economic theory indeed implies that financial integration can have implications for the response of the economy to policy shocks and, thus, for business cycle volatility. For example, the seminal work by Fleming (1962), Mundell (1963), and Dornbusch (1976) suggests that, under flexible exchange rates, the impact of a monetary policy shock on output is stronger the higher the degree of international capital mobility is. The impact of government spending shocks on output, in contrast, declines with the degree of international capital mobility. More recent theoretical work essentially supports these findings. For example, Sutherland (1996) (see also Senay, 1998) uses a variant of the dynamic sticky-price general equilibrium model developed by Obstfeld and Rogoff (1995) to demonstrate that the output fluctuations in the aftermath of monetary policy shocks increase with the degree of financial integration. However, increasing the degree of financial integration diminishes the short-run output effects of fiscal policy shocks.

So far, however, empirical literature has not been able to establish a statistically significant link between financial openness and business cycle volatility (Easterly et al., 2001; Razin and Rose, 1994).[2] Razin and Rose (1994) argue that studies might not distinguish idiosyncratic and global shocks properly. A further reason for the missing link could be structural differences of the underlying economies (Mendoza 1994).

In this paper, we revisit the link between financial openness and business cycle volatility. To this end, we lay out a stochastic dynamic general equilibrium model to derive empirically testable hypotheses on how financial market integration may influence the impact of macroeconomic shocks on business cycle volatility. We extend a model by Sutherland (1996) to incorporate a consumption function which features habit formation, a stochastic risk premium shock in financial markets, and a richer specification of the stochastic processes describing monetary and fiscal policy. Stochastic simulations of the model demonstrate that output volatility tends to be higher in economies with more open financial systems following monetary policy shocks. Financial openness also magnifies output volatility in the presence of risk premium shocks and tends to bring about only moderate changes in output volatility caused by labor supply shocks. Output volatility due to fiscal policy shocks tends to be an inverse function of financial openness.

Using a panel dataset for OECD countries for the past 40 years, we test the implications of this model empirically. We confirm earlier literature and find no

[1] See Blanchard and Simon (2001), Romer (1999), and Stock and Watson (2002) for evidence for the US. Basu and Taylor (1999), Bergman et al. (1998), Dalsgaard et al. (2002), and Kose et al. (2002) provide comparative long-term evidence for OECD countries. Kouparitsas (1998) argues that business cycle volatility has increased in the post-Bretton Woods period, but his dataset does not cover the 1990s.

[2] Evidence on the impact of the structure of financial markets on business cycle volatility seems to be more robust. Easterly et al. (2001), Denizer et al. (2002), and DaSilva (2002) find a more sophisticated financial system to be associated with lower volatility.

consistent link between financial openness and business cycle volatility for the entire sample period. Our results rather suggest that the sources of business cycle fluctuations in OECD countries have changed over time. For the 1970s, we find that measures of financial openness do not help to explain volatility. In the 1980s and 1990s, monetary, fiscal, and terms of trade volatility helped to explain business cycle volatility. Furthermore, financial openness had an impact on volatility in the 1990s, but not in the 1970s and the 1980s. We find that financial openness magnified the impact of monetary policy on volatility and that it cushioned the impact of fiscal policy on volatility in the 1990s. These results are consistent with our theoretical model. Hence, we conclude that: (i) the implications of financial openness for business cycle volatility depend upon the nature of the underlying shocks, and that, (ii) the link between macroeconomic policy, financial openness, and business cycle volatility has undergone changes over time. Parameter instability may thus be one reason why previous empirical studies have not been able to detect a statistically significant link between financial openness and business cycle volatility.

We organize the remainder of the paper as follows. In Section 2, we describe the theoretical model we use to illustrate how financial market integration may influence the impact of macroeconomic shocks on business cycle volatility. In Section 3, we describe the dataset we use in our empirical analysis. We present our empirical results in Section 4. We conclude in Section 5.

2. Financial openness and business cycle volatility: theoretical framework

Since the pioneering paper of Obstfeld and Rogoff (1995), new open economy macro (NOEM) models have become a standard tool for studying international macro issues (for recent surveys see Lane, 2001; Sarno, 2001). Compared to open economy macro-models of earlier vintage, one major advantage of NOEM models is that they provide explicit micro-foundations. Recently, Sutherland (1996) has shown how the standard NOEM model can be extended to study the implications of financial openness for macroeconomic fluctuations. This makes Sutherland's model a natural candidate for illustrating the theoretical foundations of our empirical analysis. Sutherland in particular departs from the standard model by assuming that domestic and foreign bonds are imperfect substitutes. Here, we modify Sutherland's model in three respects. First, as suggested by the results of recent empirical studies (see, e.g., Fuhrer, 2002), we assume that household consumption choices reflect habit formation. Second, we incorporate a risk premium shock that allows analyzing the implications of autonomous financial market shocks on business cycle fluctuations under alternative assumptions regarding the degree of financial openness. Third, we draw on the work by Ireland (1997) and Taylor (2001) and build into the model more complex policy reaction functions. This allows the robustness of the link between financial openness and business cycle volatility with respect to the specification of the policy regime to be analyzed.

In Sutherland's model, the world is made up of two countries which are of equal size. Each country is inhabited by infinitely lived identical households. The households form rational expectations and maximize their expected lifetime utility. In addition, each country is populated by a continuum of firms. Each country's households own the respective domestic firms. The firms sell differentiated products in a monopolistically competitive goods market. Because each firm has monopoly power on the goods market, it treats the price it charges for its product as a choice variable. When changing the prices of their products, firms have to take into consideration that there is a positive probability that they cannot change their prices again in future periods. Hence, prices are sticky. The capital stock is fixed, and the only production factor used by the firms is labor. Firms hire labor in a perfectly competitive labor market. Labor is immobile internationally.

2.1. Households

The expected lifetime utility of a domestic household is defined as $U_t = E_t \sum_{s=t}^{\infty} \beta^{s-t} u_s$, where $0 < \beta < 1$. The operator E_t denotes expectations conditional on the information set available to the household in period t. The period-utility function, u_t, is given by:

$$u_t = \left(\frac{\sigma}{\sigma-1}\right)\left(C_t/C_{t-1}^h\right)^{(\sigma-1)/\sigma} + \chi(M_t/P_t)^{1-\varepsilon}/(1-\varepsilon) - \kappa_t N_t^{\mu}/\mu, \tag{1}$$

where $\mu > 1$, $\sigma > 0$, $\varepsilon > 0$, $\chi > 0$, and $h \in [0,1)$. κ_t is a stochastic labor supply shock which (measured in terms of deviations from the steady state) follows a first-order autoregressive process with AR(1) parameter ρ_k, standard deviation σ_k, and shock term ε_{kt}. In Eq. (1), C_t denotes a real aggregate consumption index, N_t denotes the household's labor supply, M_t denotes domestic nominal money balances, and P_t denotes the aggregate domestic price index defined below. The aggregate consumption index, C_t, is defined as a CES aggregate over a continuum of differentiated, perishable domestic and foreign consumption goods of total measure unity. The aggregate consumption index is defined as $C_t = [\int_0^1 c_t(z)^{(\theta-1)/\theta}\,\mathrm{d}z]^{(\theta-1)/\theta}$, where $\theta > 1$ and $c(z)$ denotes consumption of good z. Assuming that the law-of-one-price holds for all differentiated goods, the aggregate price index is given by $P_t = [\int_0^1 p(z)^{1-\theta}\,\mathrm{d}z]^{1/(1-\theta)}$, where $p_t(z)$ is the domestic currency price of good z. With identical preferences at home and abroad and the law-of-one-price holding, purchasing power parity holds: $P_t = S_t P_t^*$, where S_t denotes the nominal exchange rate (amount of domestic currency units required to buy one unit of the foreign currency), and P_t^* denotes the aggregate foreign price level. An asterisk denotes a foreign variable.

2.2. Financial markets

Households can hold domestic and foreign nominal one-period bonds, which are traded internationally. Whereas the standard NOEM model is based on the assumption that capital markets are perfectly integrated, Sutherland (1996) introduces real transaction costs (Z_t) that drive a wedge between domestic and foreign interest rates. Here, we add an additional cost component which ensures that the foreign asset position is stationary:

$$Z_t = 0.5\psi_1 I_t^2 + 0.5\psi_2 \left[(F_t - \bar{F})/P_t^* \right]^2, \tag{2}$$

where $\psi_1 > 0$ and $\psi_2 > 0$ are positive constants, F_t denotes the stock of foreign currency denominated assets held by domestic households, and I_t denotes the level of real funds transferred from the domestic to the foreign bond market. Both Z_t and I_t are denominated in terms of C_t. The first component of transaction costs is identical to the transaction cost function used by Sutherland (1996). The second component of transaction costs implies that the foreign asset position and, thus, the steady state around which the model is log-linearized is stationary (see also Schmitt-Grohe and Uribe, 2003.) This property of the model will serve useful in the stochastic simulations of the model described in Section 2.5.

The total income received by households consists of the yield on domestic and foreign bonds, wage income, and profit income. Using this total income, households determine their consumption level, decide on their bond holdings, and hold money. In addition, they pay taxes. Consequently, the Home bond holdings of domestic consumers can be described by the following difference equation:

$$D_t = (1 + i_{t-1})D_{t-1} + M_{t-1} - M_t + w_t N_t - P_t C_t - P_t I_t - P_t Z_t + \Pi_t - P_t T_t, \tag{3}$$

where D_t = quantity of domestic currency denominated bonds, i_t = nominal interest rate on these bonds between period t and $t + 1$, w_t = nominal wage rate paid in a perfectly competitive labor market, Π_t = profit income, and T_t = real taxes (denominated in terms of C_t). The dynamics of the domestic households' foreign bond holdings are given by:

$$F_t = \left(1 + R_{t-1}^*\right)(1 + \mathrm{rp}_{t-1})F_{t-1} + P_t^* I_t, \tag{4}$$

where R_t^* = nominal foreign interest rate paid for holding a foreign bond between period t and $t + 1$, and rp_t = stochastic risk premium shock. We assume that rp_t (measured in terms of deviations from the steady state) follows a first-order autoregressive process with AR(1) parameter ρ_{rp} and standard deviation σ_{rp}.

2.3. Firms

Each profit-maximizing firm hires labor to produce a differentiated good indexed by z according to the production function $y_t(z) = N_t(z)$. The firm faces the following demand curve for its good in the monopolistically competitive goods market:

$$y_t(z) = (p_t(z)/P_t)^{-\theta} \left[C_t + C_t^* + G_t + G_t^* + Z_t + Z_t^* \right]/2. \tag{5}$$

Firm's profits are given by $\Pi_t(z) = p_t(z)y_t(z) - w_t y_t(z)$. When maximizing these profits, each firm has to take into account that there is a positive probability $0 < \gamma < 1$ that it cannot revise its price setting decision made in period $s < t$ in period t. Thus, prices are sticky. Following Calvo (1983), firms maximize the expected present value, $V_t(z)$, of current and future profits, where profits in period s ($s > t$) are weighted by the probability that the current period price, $p_t(z)$, will still be in force in period s. Firms, thus, maximize:

$$V_t(z) = E_t\left(\sum_{s=t}^{\infty} \gamma^{s-t} R_{t,s} \Pi_s(z)/P_s\right), \tag{6}$$

where $R_{t,s}$ is the market real discount factor. Given the price of the differentiated good z, the quantity produced by the firm can be derived from the demand function for this good.

2.4. The government

The domestic government collects lump-sum taxes and uses them together with seignorage revenues to finance real government purchases, G_t:

$$G_t = T_t + (M_t - M_{t-1})/P_t, \tag{7}$$

where real government purchases are denominated in terms of C_t. We assume that fiscal policy behaves according to a simple fiscal policy rule (Taylor, 2001): $\hat{G}_t = f\hat{y}_t + \rho_G \hat{G}_{t-1} + \varepsilon_{Gt}$, where $\rho_G \in [0,1]$, ε_{Gt} is a serially uncorrelated stochastic disturbance term with standard deviation σ_G, and a variable with a hat denotes percentage deviations from the steady state. The first-term on the right-hand side captures the influence of automatic stabilizers. To model monetary policy, we assume that the government may respond contemporaneously to the technology shocks. We adopt a money supply rule similar to the one suggested by Ireland (1997): $\hat{M}_t = \rho_M \hat{M}_{t-1} + \beta_{Mk}\varepsilon_{kt} + \varepsilon_{Mt}$, where ε_{Mt} is a serially uncorrelated stochastic disturbance term with standard deviation σ_M.

2.5. Simulation results

To derive testable implications for the impact of financial openness on business cycle volatility, we solve the model numerically. In a first step, we follow Obstfeld and Rogoff (1995) and Sutherland (1996) and log-linearize the model around a symmetric flexible-price steady state in which the countries foreign asset positions are zero. In a second step, we simulate the model numerically.[3] In the numerical simulations, we assume that the innovation terms, $\varepsilon_{j,t}$, $j \in \{M, G, k, \text{rp}\}$ are perfectly negatively correlated (shocks are asymmetric across countries). The calibration of the model is given in Table 1 and closely follows Sutherland (1996).

Fig. 1 depicts impulse response functions that give the response of real domestic output to an unanticipated one unit money supply, government spending, labor

[3] We use Paul Klein's algorithm in Matlab to solve the model numerically. See Klein (2000) for a discussion of the technical details.

Table 1
The calibrated parameters

Parameter	Value	Description
β	1/1.05	Subjective discount factor
σ	0.75	Intertemporal elasticity of substitution
θ	6.0	Intratemporal elasticity of substitution
ε	9.0	Elasticity of utility from real balances
μ	1.4	Labor supply elasticity
h	0.8	Habit persistence parameter
ψ_1	5 (0)	First component of costs for undertaking positions in international financial market in the case of low (high) capital mobility
ψ_2	0.05	Second component of costs for undertaking positions in international financial market
n	0.5	Country size
ρ_G	0.9	Autoregressive coefficient of the fiscal policy process
σ_G	0.03	Standard deviation of fiscal policy shock
ρ_M	0.7	Autoregressive coefficient of the money supply process
σ_M	0.002	Standard deviation of monetary policy shock
ρ_{rp}	0.5	Autoregressive coefficient of the risk premium process
σ_{rp}	0.04	Standard deviation of risk premium shock
ρ_k	0.95	Autoregressive coefficient of the labor supply shock process
σ_k	0.006	Standard deviation of the labor supply shock
β_{Mk}	0.09	Contemporaneous response of monetary policy to labor supply shocks
f	−0.5	Impact of automatic stabilizers on government spending

The habit persistence parameter is taken from Fuhrer (2002). The autoregressive coefficient of the fiscal policy process is in line with the specification given in Chari et al. (1995). The risk premium process is calibrated as in McCallum and Nelson (1999). The technology and the money supply processes are in line with the empirical estimates of Ireland (1997), which also provide an estimate of the contemporaneous response of monetary policy to productivity shocks. We use this result to calibrate our labor supply shock process. The numerical values assigned to the autoregressive parameter and the standard deviation of the fiscal policy shock are set equal to the values used in Chari et al. (1995). The parameter capturing the impact of automatic stabilizers on government spending is taken from Taylor (2001). The other parameters are as in Sutherland (1996).

supply, and risk premium shock. We compare impulse responses that we obtain when the degree of international capital mobility is high ($\psi_1 = 0$) and low ($\psi_1 = 5$). Money supply shocks, for instance, lead to a reduction of the domestic nominal interest rate. Because prices are sticky, the domestic real interest rate declines and domestic households start increasing consumption. Because the foreign nominal interest rate increases the international nominal interest rate differential becomes negative. Domestic households seek to accumulate foreign assets and the nominal exchange rate depreciates. The latter translates into a depreciation of the terms of trade, which, in turn, triggers an expenditure switching effect. The result is an increase of the demand-determined output at home. Abroad, opposite effects are at work and, thus, a real contraction ensues in the short run.

If international bond markets are segmented, the condition of uncovered interest rate parity implies that the international nominal interest rate differential is equal to

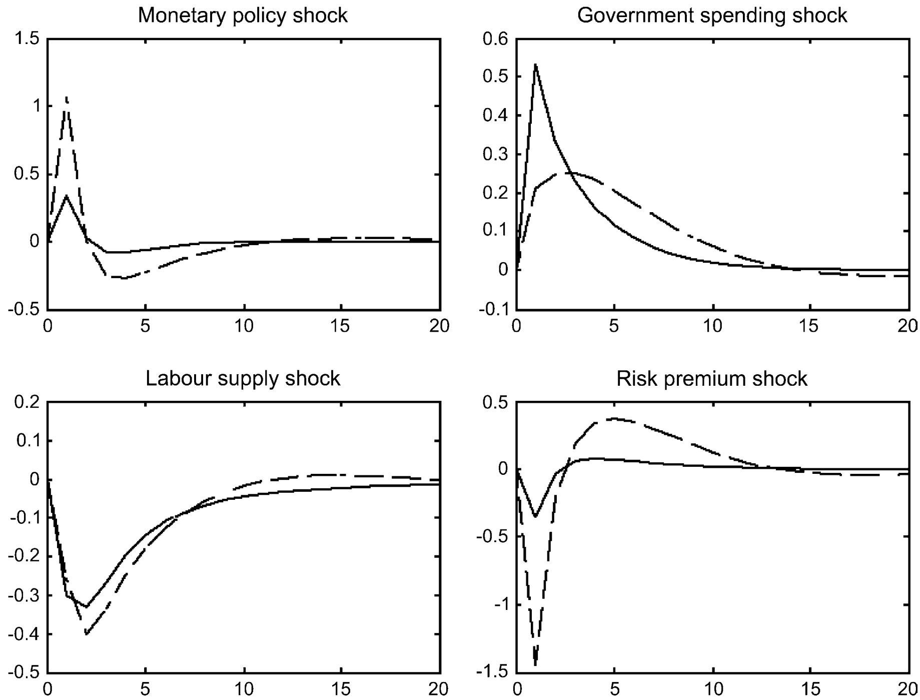

Fig. 1. Impulse response functions. The figure depicts the response of output to a one-unit monetary policy, government spending, labor supply, and risk premium shock. Output is measured in terms of percentage deviations from its steady state value. Dashed lines obtained when setting $\psi_1 = 0$ (high financial integration) and solid lines obtained when setting $\psi_1 = 5$ (low financial integration).

the sum of the expected rate of change of the nominal exchange rate and the transaction costs for investing in international bond markets. Neglecting the second component of the transaction cost function (2), the transaction costs are proportional to the expected rate of change of the cross-border flow of funds. Because domestic households accumulate foreign bonds in the short run but the foreign asset position is zero in the long run, the expected rate of change of the cross-border flow of funds is negative. It follows that, for a given negative international nominal interest rate differential, the appreciation expectations are smaller in the case of segmented bond markets than in the case of perfectly integrated bond markets. In consequence, the actual exchange rate depreciation triggered by an expansionary monetary policy shock is smaller when bond markets are segmented. It follows that the output effect of a monetary policy shock is smaller in the case of low capital mobility as compared to the case of high capital mobility.

Fig. 1 reveals that the impact of money supply and risk premium shocks on output tends to be stronger when international bond markets are integrated perfectly. In contrast, financial market integration tends to dampen the impact effects of a government spending shock on output. In the case of labor supply shocks, the impact effect of the shock on output is stronger if financial markets are imperfectly integrated, but the maximum deviation of output from steady state is larger if financial markets are perfectly integrated.

To explore further the quantitative implications of financial openness for output volatility, we report in Table 2 the results of stochastic simulations of the model. We report results for three policy regimes. In the first policy regime, monetary and fiscal policies follow the policy rules given above. In the second policy regime, fiscal policy follows the policy rule given above but the monetary policy rule is simplified by setting $\beta_{Mk} = 0$. In the third policy regime, the monetary policy rule is as given above but the fiscal policy rule is assumed to be a simple AR(1) process ($f = 0$).

The simulation results show that in the three policy regimes, output volatility tends to be higher in a world of high capital mobility if monetary policy shocks hit the system. In contrast, a higher degree of international capital mobility tends to cushion the output effects of fiscal policy shocks. These results are in line with the implications of the textbook version of the Mundell–Fleming model, which implies that financial market integration enhances (cushions) the output effects of monetary policy (fiscal policy) shocks. The simulation results further demonstrate that financial openness magnifies output volatility in the wake of risk premium shocks.

Table 2
Simulation results

Shock	Regime	
	Low capital mobility	High capital mobility
	Benchmark simulation	
Money supply	0.0749	0.2358
Fiscal policy	2.0873	1.9479
Risk premium	1.5550	6.7166
Labor supply	0.3550	0.4045
	Benchmark simulation + $\beta_{Mk} = 0$	
Money supply	0.0750	0.2362
Fiscal policy	2.0785	1.9535
Risk premium	1.5449	6.66892
Labor supply	0.3769	0.4423
	Benchmark simulation + $f = 0$	
Money supply	0.0978	0.2471
Fiscal policy	7.4391	7.1255
Risk premium	2.1403	7.2154
Labor supply	1.0696	1.0583

The table reports standard deviations for Home output for alternative Taylor-rule specifications. To compute the standard deviations, 100 time series of the endogenous variables of the model were generated, each time series consisting of 500 observations. In the simulations, it was assumed that Home and Foreign monetary policy shocks are perfectly asymmetric. To simulate the models with a low (high) degree of international capital mobility, it was assumed that $\psi_1 = 5$ ($\psi_1 = 0$).

Finally, financial market integration brings about only moderate changes in output volatility caused by labor supply shocks.

3. Stylized facts

The empirical evidence we use to test the implications of the above model is based on annual data for 24 OECD countries for the years 1960–2000. We give a full list of the countries and the variables used in Table 3. We use the band-pass filter advocated by Baxter and King (1999) to calculate the cyclical component of the time series under investigation (Christiano and Fitzgerald, 1999; Mills, 2000). We then measure the volatility of a time series at business cycle frequencies as the standard deviation of the cyclical component of the time series. One advantage of the band-pass filter is that it decomposes time series into trend, cycle, and irregular components that correspond to the low frequencies, the business cycle, and the high frequencies of the spectrum (Stock and Watson, 1998). Baxter and King (1999) recommend interpreting fluctuations longer than two years and shorter than eight years as the cyclical part of the time series. We follow this line of argumentation and use a band-pass (2,8) filter.

Table 4 shows the development of business cycle volatility, and of financial openness over time. Volatility peaked in the 1970s and has been on a decline in the 1980s and the 1990s. However, the aggregated data cloud that some countries have also witnessed an increase in volatility in the 1990s compared to earlier decades. This group of countries includes Finland, Japan, Norway, Mexico, and Turkey. The fact that some of these countries have also experienced quite severe financial crises might be seen as a first hint that the financial sector has an impact on output fluctuations.

We approximate financial openness through various variables. Gross and net foreign assets of commercial banks measure the openness of the banking system. Gross capital flows (the sum of foreign direct investment, portfolio flows, and bank lending) relative to GDP provide a broader assessment of financial openness since they also capture other financial market segments. Because gross and net foreign assets may be endogenous, we also use information on capital controls as an exogenous measure of financial openness. Generally, the countries in our sample have shown increasing degrees of financial openness over the sample period, irrespective of the measure used. Gross and net foreign assets of commercial banks have increased continuously. Likewise, capital flows over GDP have shown an upward trend throughout, particularly in the second half of the 1990s. Japan, Korea, and Mexico are the only countries for which capital flows have declined recently as a response to the financial crises of the 1990s.

Fig. 2 plots our measures of financial openness against business cycle volatility for the full sample. Correcting for outliers (Mills, 2000), we find some evidence (correlation coefficients of around −0.3) that business cycles in more open financial systems are less pronounced than elsewhere. This negative link, however, holds only for our gross measures of financial openness. Banks' net foreign assets are virtually uncorrelated to volatility.

Table 3
Data and definitions

Variable	Definition	Source
Business cycle volatility	Standard deviation of the band/pass (2,8) filtered real GDP (volume index, 1995 = 100).	Organization for Economic Cooperation and Development (OECD), 2001. International Direct Investments (CD-Rom), Paris. International Monetary Fund (IMF), 2002. International Financial Statistics on CD-Rom, Washington, DC
Capital controls	Index of restrictions on capital account transactions (after 1996: controls on financial or commercial credits)	Before 1996: kindly provided by Gian-Maria Milesi-Ferretti. After 1996: International Monetary Fund (IMF), 1998. Exchange Arrangements and Exchange Restrictions – Annual Report 1998, Washington, DC
Financial openness	(1) Gross capital flows (average of capital in- and outflows) in percent of GDP where gross capital inflows (outflows) = sum of foreign direct investment in reporting country (FDI abroad), portfolio liabilities (assets), other investment liabilities (assets) (2) Banks' net foreign assets in percent of banks' total assets (absolute value) (3) Deposit money banks' foreign assets in percent of banks' total assets (structural break for Ireland in 1980 adjusted manually)	International Monetary Fund (IMF), 2002. International Financial Statistics on CD-Rom, Washington, DC
Interest rate	Standard deviation of short-term interest rates (money market rate or alternative short-term lending rate)	International Monetary Fund (IMF), 2002. International Financial Statistics on CD-Rom, Washington, DC
Government expenditure	Standard deviation of rate of change of real non-wage government consumption	Organization for Economic Cooperation and Development (OECD), 2001. International Direct Investments (CD-Rom), Paris
Oil price shock	Standard deviation of the rate of change of the price of oil measured in US dollar deflated by the US-deflator of real GDP	Organization for Economic Cooperation and Development (OECD), 2001. International Direct Investments (CD-Rom), Paris
Terms of trade	Standard deviation of the change of the terms of trade	Organization for Economic Cooperation and Development (OECD), 2001. International Direct Investments (CD-Rom), Paris
Country sample	Australia, Austria, Belgium and Luxembourg, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, Korea, Mexico, The Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, Turkey, United Kingdom, US	

Table 4
Descriptive statistics

	1960s	1970s	1980s	1990s
Output volatility (percentage points)	1.29	1.58	1.20	1.17
Capital controls on financial credits (1 = capital controls, 0 = no controls)	0.78	0.65	0.48	0.17
Gross capital flows/GDP (%)	6.36	9.06	12.76	21.53
Banks' gross foreign assets/total assets (%)	8.99	14.81	19.32	21.43
Banks' net foreign assets/total assets (%)	2.93	2.74	5.5	6.42

The table reports mean values for the variables used in the panel regressions. For definitions of the variable used see Table 3.

4. Determinants of business cycle volatility

In this section, we present more systematic evidence on the determinants of business cycle volatility. Our analysis is based on a panel of non-overlapping averages of the data. Though the concrete choice of the time span is clearly arbitrary, we average over a time span of five years, which is often seen as the typical length of a business cycle. Our empirical analysis proceeds in three steps. First, we use Granger non-causality tests for the entire panel to test whether financial openness and business cycle volatility are linked. Second, we use multivariate panel regressions to account for additional factors that might influence business cycle volatility. Third, because business cycle volatility may have changed over time (Stock and Watson, 2002; Blanchard and Simon, 2001), we use cross-section regressions for individual decades.

4.1. Granger non-causality tests

In this section, we test for the presence of non-causality between financial openness and business cycle volatility. Although the theoretical model discussed in Section 2 suggests that greater financial openness might influence the volatility of business cycles, the presence of reverse causality is possible as well. Agents might decide to hold more international assets and, thus, to diversify risk to a greater degree if business cycle fluctuations are large. Hence, the observed degree of financial openness might be a consequence rather than a cause of business cycle volatility. We thus test for Granger non-causality between financial openness and business cycle volatility. We also include tests for the direction of causality between the volatility of short-term interest rates and real government spending, on the one hand, and business cycle volatility, on the other hand. These are the variables that we will use below to capture monetary and government spending volatility. To test for Granger non-causality, we use the panel model by Holtz-Eakin et al. (1988):

$$y_{i,t}=\alpha_0+\sum_{j=1}^{m}\alpha_j y_{i,t-j}+\sum_{j=1}^{m}\delta_j x_{i,t-j}+f_i+u_{i,t}, \tag{8}$$

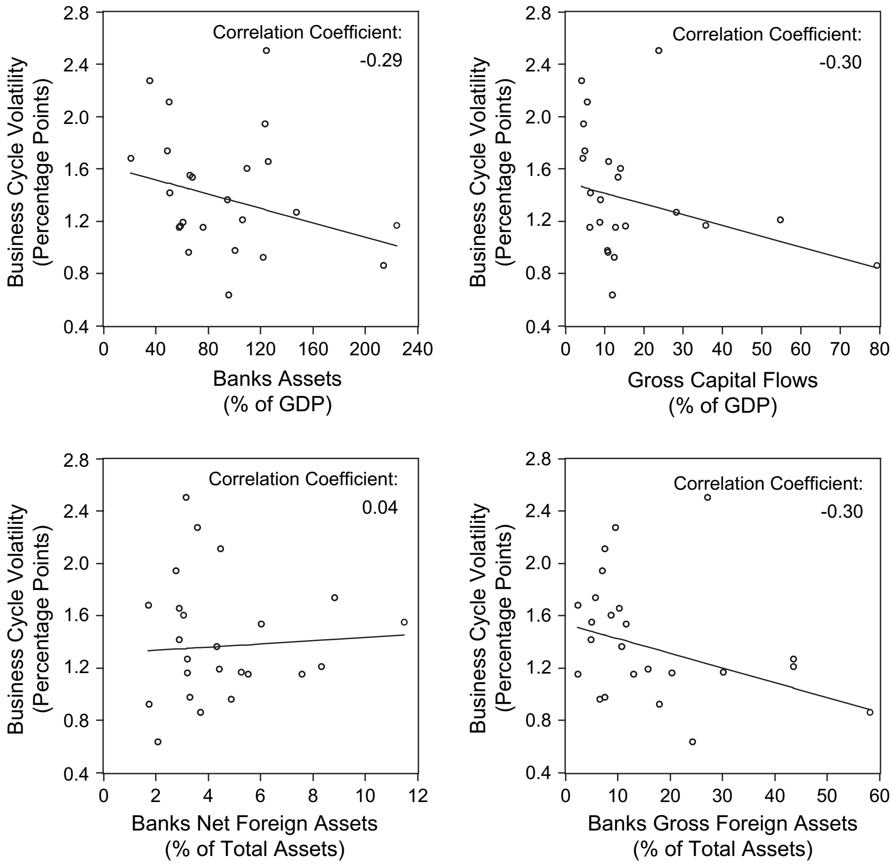

Fig. 2. Correlation between business cycle volatility and openness. Business cycle volatility is measured through the volatility of real GDP. See Table 3 for details.

where $i = 1, \ldots, N$ denotes the number of cross-sections of the panel, f_i is a country-specific fixed effect, and $u_{i,t}$ is on error terms. By calculating first differences of the data, fixed effects can be eliminated:

$$y_{i,t} - y_{i,t-1} = \sum_{j=1}^{m} \alpha_j \left(y_{i,t-j} - y_{i,t-j-1} \right) + \sum_{j=1}^{m} \delta_j \left(x_{i,t-j} - x_{i,t-j-1} \right) + \left(u_{i,t} - u_{i,t-1} \right). \tag{9}$$

Within this model, x Granger-causes y if the joint hypothesis $\delta_j = 0 \ \forall \ j$ cannot be rejected. Since the residuals are, by definition, correlated with the endogenous variables, an instrumental variable estimator is warranted. Also, the problem of lagged endogenous regressors has to be addressed, whereas in a lot of applications the resulting so-called Nickell-bias can be neglected because the time dimension is relatively large as compared to the cross-section dimension, this does not hold in our case. The use of non-overlapping averages implies that our panel is of a ‘‘short and wide"-type, and simply using OLS would lead to inconsistent and biased estimators. Therefore, we follow Judson and Owen (1999) who show that, for an unbalanced panel with $T \leq 10$, Arellano and Bond's one-step GMM-estimator (Arellano and Bond, 1991) outperforms alternative estimators. To check the robustness of our results, we also report the dynamic panel IV-estimator advocated by Anderson and Hsiao (1982).

The results of the causality tests are reported in Table 5. Causality would imply that the coefficient on the lagged exogenous variable (column 2) is exogenous. Generally, we do not find evidence for a link between financial openness and business cycle volatility. This holds irrespective of the way we measure volatility and openness. For the policy variables, we obtain two significant effects. In one specification, there is a negative effect running from government spending volatility to business cycle volatility. One specification also suggests that the volatility of short-term interest rates might be driven by changes in business cycle volatility. We will account for the possible endogeneity of interest rates in one of the robustness checks below.

4.2. Multivariate regressions

A possible shortcoming of bivariate Granger tests of non-causality is that they do not account for the possibility that a third variable influences both time series under investigation. Therefore, we include additional sources of business cycle volatility in a multivariate panel regression. We control for the volatility of government spending (standard deviation of the growth in real non-wage government consumption), the volatility of monetary policy (standard deviation of short-term interest rates), and supply side volatility (the standard deviation of the growth in the terms of trade)[4] (see also Karras and Song, 1996):

$$\sigma_{i,t}^{\text{cycle}} = \alpha_{0,i} + \alpha_{1,t} + \sum_{i=1}^{m} \beta_i \sigma_{i,t}^{\text{control}} + u_{i,t}. \tag{10}$$

In Eq. (10), business cycle volatility ($\sigma_{i,t}^{\text{cycle}}$) depends on country ($\alpha_{0,i}$) and time-fixed effects ($\alpha_{1,t}$), and on the volatility of the control variables ($\sigma_{i,t}^{\text{control}}$). Table 6 summarizes the results of three sets of regressions. First, we estimate a baseline specification with the control variables being included only.

Second, we include a measure of financial sector openness. A large number of financial openness measures have been used in the literature (see, e.g., Easterly et al., 2001; Razin and Rose, 1994). We use principal components to combine two measures of financial openness (banks' foreign assets and gross capital flows, both variables being standardized) into one measure of gross financial openness.

Third, we include interaction terms between financial openness and the standard deviation of short-term interest rates and government spending. (Rose (1996) uses

[4] In the panel regressions, we use the price of oil measured in US-dollars instead. The use of this variable is also motivated by Kneller and Young (2001), who find that oil prices are important when modeling cross-country difference in business cycle volatility.

Table 5
Tests on Granger non-causality

Endogenous variable	Coefficient of lagged endogenous variable (1)	Coefficient of lagged exogenous variable (2)	Test of second order autocorrelation (Pr > z)	Sargan test of overidentifying restrictions ($\chi^2(20)$)	Result
(a) Arellano/Bond estimator					
Capital controls	0.62*** (6.66)	0.03 (0.60)	0.39	39.10***	No causality
σ_Y	0.24* (1.90)	0.11 (0.63)	0.57	26.31	
Openness	0.48*** (3.68)	0.09 (0.37)	0.42	24.66	No causality
σ_Y	0.26* (1.92)	−0.09 (−1.29)	0.56	23.95	
σ_i	−0.80*** (−3.15)	0.04 (0.15)	0.17	52.90***	No causality
σ_Y	0.29** (2.15)	−0.02 (−0.95)	0.26	29.58*	
σ_G	0.01 (0.32)	−1.84 (−1.27)	0.62	44.79***	σ_G causes σ_Y (−)
σ_Y	0.18 (1.41)	−0.01*** (−3.72)	0.76	29.10*	

Endogenous variable	Coefficient of lagged endogenous variable (1)	Coefficient of lagged exogenous variable (2)	Test result
(b) Anderson/Hsiao-estimator			
Capital controls	0.98 (0.34)	0.02 (0.20)	No causality
σ_Y	0.42 (1.01)	0.06 (0.29)	
Openness	2.72 (0.43)	−0.35 (−0.20)	No causality
σ_Y	1.25 (1.59)	−0.20 (−1.12)	
σ_i	−0.56*** (−5.75)	0.29* (1.66)	σ_Y causes σ_i (+)
σ_Y	0.96 (1.53)	−0.04 (−1.02)	
σ_G	−1.29 (−0.98)	0.12 (0.17)	No causality
σ_Y	0.49 (1.10)	−0.01 (−1.10)	

σ_Y = Standard deviation of band-pass filtered real GDP, σ_i = standard deviation of short-term interest rates, σ_G = standard deviation of band-pass filtered real government spending. Tests are based on non-overlapping 5 year averages. Openness = first principal component of gross measures of financial openness defined in Table 3. Capital controls refer to controls on financial credits. The Anderson/Hsiao estimator has been specified using the twice lagged level of the endogenous variable as instruments. In brackets: z-values. ***(**, *) Denotes rejection of the null hypothesis of a zero coefficient at a 1(5, 10) percent level of significance.

a similar approach to explain the volatility of exchange rates.) The motivation behind these interaction terms is the result of the theoretical model presented above that the effectiveness of different policy measures should depend on the degree of financial openness.

In the panel regressions, hardly any of the control variables is statistically significant (Table 6, column 1), and the explanatory power of the regressions is fairly small (adjusted R^2 of 0.2). We find only weak evidence for a positive link between the volatility of oil prices and the volatility of real GDP.[5] The poor explanatory power of our control variables and the missing evidence for a stable

[5] Evidence for a positive link between interest rate volatility and real GDP is somewhat stronger if we use a GLS estimator instead (results not reported). In this specification, we also find a positive effect of the interaction term between the volatility of government spending and financial openness.

Table 6
Multivariate regressions: OECD countries

	Panel regressions (1)	Cross-section regressions		
		1970s (2)	1980s (3)	1990s (4)
	Baseline regression			
Constant	1.15*** (5.91)	1.53*** (3.84)	0.91*** (6.73)	0.57*** (4.79)
σ_i	0.01 (0.61)	0.06 (0.37)	0.06*** (7.30)	0.12*** (7.72)
σ_G	0.003 (0.15)	−0.002 (−0.14)	0.004*** (3.22)	−0.004 (−0.15)
σ_{TOT}	0.01 (1.59)	−2.84 (−0.78)	1.43*** (3.18)	0.03*** (15.48)
R^2	0.15	0.04	0.34	0.78
N	110	19	21	24
	Including gross financial openness			
Constant	1.37*** (5.69)	1.66*** (4.17)	0.89*** (5.70)	0.64*** (4.46)
σ_i	0.01 (0.56)	0.04 (0.29)	0.06*** (6.65)	0.09*** (4.86)
σ_G	0.04 (0.24)	−0.01 (−0.86)	0.004* (1.97)	0.05 (1.29)
σ_{TOT}	0.006 (1.46)	−0.84 (−0.23)	1.49** (2.76)	0.03*** (8.17)
Openness	0.08 (1.46)	0.17 (1.36)	0.07 (0.41)	−0.14 (−1.39)
R^2	0.17	0.08	0.35	0.81
N	110	19	21	23
	Including interaction terms			
Constant	1.21*** (6.37)	1.47*** (3.61)	0.90*** (5.63)	0.47*** (3.65)
σ_i	0.003 (0.16)	0.09 (0.60)	0.05* (1.85)	0.06** (2.64)
σ_G	0.01 (0.64)	−0.01 (−0.43)	0.01 (0.43)	0.16** (2.51)
σ_{TOT}	0.008* (1.80)	−0.81 (−0.20)	1.54* (2.00)	0.02*** (5.83)
Openness $\times\ \sigma_i$	0.03 (1.57)	0.05 (0.50)	0.04 (0.38)	0.03* (2.05)
Openness $\times\ \sigma_G$	0.01 (1.04)	0.01 (1.08)	−0.00 (−0.02)	−0.07** (−2.81)
R^2	0.20	0.09	0.36	0.83
N	110	19	21	23

The dependent variable is the average volatility of business cycles in OECD countries (volatility of real GDP). σ_{TOT} = standard deviation of the change in the terms of trade (for the panel: of oil prices). Openness = first principal component of gross measures of financial openness defined in Table 3. σ_G = standard deviation of band-pass filtered real government spending. Definitions of the explanatory variables are given in Table 3. The panel regressions in the first column are based on 5-year non-overlapping averages (1960–1964, 1995–1970, …, 1995–2000). Country and time-fixed effects are included. *** (**, *) Denotes significant at the 1 (5, 10) percent level. Robust t-statistics are given in brackets.

link between openness and business cycle volatility could be due to structural breaks in the data. Recent empirical evidence on the US business cycle indeed suggests that there have been gradual changes in business cycle volatility, and that the 1970s and the first half of the 1980s have been 'special' (Stock and Watson, 2002). The stylized facts presented in Table 4 likewise suggest that the relationship between financial openness and business cycle volatility might have changed over time. While, in the 1970s, countries have gradually opened up for foreign capital, business cycle volatility has been on the rise. In the 1980s and 1990s, in contrast, financial openness has increased even further, this time being accompanied by a decline in business cycle volatility.

Although we have so far allowed for possible shifts in the intercept over time by including time-fixed effects in the panel regressions, this does not rule out the possibility that individual coefficient estimates have changed. To account for this, we additionally run our baseline regression using cross-section data for the averages of business cycle volatility in the 1970s, 1980s, and 1990s. These cross-section regressions do indeed reveal quite significant changes in the determinants of output volatility (columns 2–4 of Table 6). For the 1970s, we find none of our explanatory variables to be of statistical significance, and the overall R^2 is very low. For the 1980s and 1990s, there is evidence that the volatility of interest rates, government spending, and the terms of trade has contributed to business cycle volatility. For the 1990s, we are able to explain almost three-fourths of the cross-country variation in volatility.[6]

There is some evidence for the hypothesis that the link between financial openness and business cycle volatility has changed over time, whereas higher financial openness seems to have been associated with higher business cycle volatility in the 1970s, this does not seem to be the case anymore in the 1990s. While financial openness, measured through gross financial flows, is insignificant in all cases, the negative link between openness and volatility is significant in the 1990s when using capital controls as a proxy (results not reported). Generally, the changing impact of openness could explain the 'missing link' between openness and volatility reported above for the panel regressions. One possible explanation for these results is that the 1970s were special due to the large oil price shocks and the end of the Bretton Woods System.

Interacting openness with the standard deviation of interest rates and government spending, as is done in the last panel of Table 6, provides significant coefficients for the 1990s and the signs of these coefficients are in line with economic theory: in financially more open economies, the impact of volatility due to monetary policy increases, and the impact of volatility due to government expenditure declines (column 4).

4.3. Robustness checks[7]

We consider a number of alternative specifications to test the robustness of our results. Because the Granger tests for non-causality presented in Table 5 suggest that the volatility of short-term interest rates might be driven by business cycle volatility, we re-estimate the regressions reported in Table 6 using lagged interest rate volatility as an instrument for the current one. The results of these IV-estimators are essentially the same as the OLS-estimators, suggesting that the endogeneity of the interest rates does not bias our results significantly.

For the cross-section regressions, given the small sample, our results might be influenced by some influential outliers. To take this possibility into account, we also

[6] A qualitatively similar result (not reported) is obtained when using capital controls as a proxy for financial openness.

[7] Not all of these results are reported.

estimate the cross-country regressions using the trimmed-least-squares technique. The results do not change qualitatively.

Also, we test whether the changing sample size when moving from the 1970s to the 1990s affects the results. When estimating the regressions for a constant country sample, the different impact of openness between the different decades remains significant. Some of the coefficients of the policy variables, in contrast, remain insignificant when moving from the 1970s to the later decades in the fixed country sample.

Moreover, the results might be influenced by the fact that only OECD countries are included that do not differ sufficiently along certain structural dimensions such as financial sector openness. Therefore, we re-estimate the cross-section equation for the 1990s with a substantially larger cross-section sample of almost 80 developed and developing countries (Table 7). Although the specifications differ because we cannot use the same methodology as before to measure volatility[8] and because we lack fully comparable data for all countries, some of the qualitative results are similar. Although specific coefficient estimates differ because of differences in the computation of the variables, the economic significance of, for instance, the volatility of interest rates is similar, explaining about 40% in the variation of output volatility.[9] Moreover, we now measure financial openness through capital controls (a dummy variable which is equal to one if controls are imposed on cross-border financial credits). Hence, a positive coefficient on this variable would imply that countries with more restrictive capital control regimes have *higher* business cycle volatility. Although the capital control dummy alone is insignificant, we again find that the interaction between openness and the volatility of interest rates (government spending) has positive (negative) implications for output volatility (column 3 of Table 7). However, the latter effect is significant only if the volatility of government spending itself is not included due to a high degree of multicollinearity of this variable with the interaction term.

In the specification using the interaction terms, (log) GDP per capita is also significant and negative. We use this variable to capture characteristics of less developed countries that might contribute to business cycle volatility, such as a low degree of development of domestic financial markets or a high volatility of their terms of trade. According to our results, output is less volatile in more developed countries.

5. Conclusions

Although conventional wisdom suggests that increased financial openness might change business cycle fluctuations, the literature has been relatively unsuccessful to

[8] More specifically, having only 10 or fewer years of data, we cannot use the band-pass filter but simply compute the standard deviation of the respective growth rates.

[9] This figure has been computed as the standard deviation of interest rates, multiplied by the coefficient estimate, divided by the standard deviation of output volatility.

Table 7
Multivariate regressions: extended country sample

	Baseline (1)	Including capital controls (2)	Including interaction with terms (3)
Constant	4.50** (2.14)	4.58* (1.84)	6.08*** (2.89)
σ_i	1.91** (2.15)	1.88** (2.09)	6.26*** (2.98)
σ_G	0.16** (1.94)	0.16* (1.79)	0.01 (0.08)
log (GDP per capita)	−0.28 (−1.43)	−0.29 (−1.24)	−0.52** (−2.48)
Openness		0.09 (0.12)	
Openness × σ_i			−4.38** (−2.04)
Openness × σ_G			0.18 (1.54)
R^2	0.32	0.32	0.36
N	77	74	74

The dependent variable is the standard deviation of real GDP growth in the 1990s. The volatility of government consumption is the standard deviation of the growth in real government consumption. The volatility of interest rates is the coefficient of variation of nominal lending rates. Capital controls = dummy set equal to one if country has capital controls on cross-border financial credits, i.e., greater openness implies a *decline* in the variable. *** (**, *) Denotes significant at the 1 (5, 10) percent level. Robust *t*-statistics are given in brackets.

date to establish this link empirically. Structural differences between countries included in panel studies (Mendoza, 1994) or the inability to distinguish between idiosyncratic and global shocks (Razin and Rose, 1994) have been offered as explanations. This paper confirms that there has not been a stable relationship between financial openness and business cycle volatility over time, and it has offered two additional explanations for this ‘missing link’:

First, parameter instability may be one explanation why empirical studies fail to find a link between financial openness and business cycle volatility. Our results suggest that the link between macroeconomic policy, financial openness, and business cycle volatility has changed over time. Using a panel dataset for OECD countries for the past 40 years, we find no consistent link between financial openness and business cycle volatility for the entire sample period. Estimates for individual decades show that the sources of business cycle fluctuations have changed. Financial openness seems to have been cushioning rather than magnifying business cycle fluctuations in the 1990s but not in earlier decades.

Second, the link between financial openness and business cycle volatility seems to depend upon the nature of the underlying shock. In line with the theoretical model outlined in Section 2, our empirical estimates indicate that the impact of interest rate volatility on business cycle volatility is enhanced in open financial markets, while the impact of volatility of government spending is diminished. One interpretation of this finding could be that monetary policy is more effective in financially integrated markets while the reverse seems to hold true for fiscal policy. Such an interpretation would be consistent with the results of both the classic Mundell–Fleming model and the NOEM model outlined in this paper.

There are a number of dimensions along which this paper could be extended. First, in order to test theoretical models on the link between business cycle volatility and financial openness more rigorously, it would be necessary to identify monetary and fiscal shocks more precisely by looking at evidence from individual countries. Second, because the implications of financial openness for business cycle volatility may depend upon the macroeconomic aggregate considered, it would be interesting to study the implications of financial openness for, e.g., investment volatility. To derive empirically testable implications in this respect, it would be possible to extend the theoretical model studied in this paper by modeling the process of capital accumulation. Third, because our empirical results have shown that the 1970s were 'special', it would also be interesting to extend the model to incorporate a sector using imported intermediate raw materials (oil) and to incorporate a third oil-producing country. Such extensions would allow for a more detailed analysis of the role played by imported raw materials for the link between financial openness and business cycle volatility.

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