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Financial Openness and Business Cycle Volatility

by

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Abstract

This paper discusses whether the integration of international financial markets affects business cycle fluctuations. 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.

Keywords: Open Economy Macroeconomics; Monetary union; Business cycles; Financial markets

JEL classification: F33, F36, F41

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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 deeper than it used to be.[1] At the same time, business cycle fluctuations in OECD countries have declined,[2] and changes in business cycle characteristics seem to be related to changes in the degree of capital mobility (Basu and Taylor 1999).

Economic theory indeed implies that the integration of international financial markets can have important implications for the response of the real economy to policy shocks. For example, work by Fleming (1962), Mundell (1963), or Dornbusch (1976) suggests that, under flexible exchange rates, the impact of a given monetary policy shock on real output is stronger the higher the degree of international capital mobility. The impact of government spending shocks on real output, in contrast, declines with the degree of international capital mobility.

Recent theoretical work likewise supports that the openness of the financial system can have implications for the impact of monetary and fiscal policies on business cycles. For example, Sutherland (1996) and Senay (1998) use a variant of the dynamic sticky-price general equilibrium model developed by Obstfeld and Rogoff (1995) to demonstrate that the amplitude of real output fluctuations in the aftermath of monetary policy shocks should increase with the degree of international financial market

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1 Bordo et al. (1998), for instance, argue that the degree of securitization is much higher now than during earlier episodes of financial integration.

2 See Blanchard and Simon (2000), Romer (1999), and Stock and Watson (2002) for evidence for the United States. Basu and Taylor (1999), Bergman et al. (1998), and Dalsgaard et al. (2002) provide comparative long-term evidence for OECD countries. Kouparitsas (1998) reaches a different conclusion and argues that business cycle volatility has increased in the post-Bretton Woods period, but his dataset does not cover the 1990s.

integration. In contrast, increasing the degree of financial market integration diminishes the short-run output effects of fiscal policy and of labor productivity shocks.

So far, the empirical literature has not been able to establish a statistically significant link between financial openness and business cycle volatility though (Easterly et al. 2000, Razin and Rose 1994).[3] Razin and Rose (1994) argue that this could be due to the fact that idiosyncratic and global shocks are not distinguished properly. A further reason for the missing link could be structural differences of the underlying economies that empirical panel-studies do not take into account (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 business cycle model to derive empirically testable hypotheses on how financial market integration may influence the impact of macroeconomic shocks on business cycle volatility. The theoretical framework builds on Sutherland (1996). To make the model more realistic, we use a consumption function which incorporates habit formation, a stochastic risk premium shock in financial markets, and a richer specification of the stochastic processes describing monetary and fiscal policy. The main qualitative results from the model are unaffected by these modifications.

Stochastic simulations of this model demonstrate that output volatility tends to be higher in economies with more open financial systems if monetary policy shocks hit the model economy. Financial openness magnifies output volatility in the presence of risk premium shocks and productivity shocks. The proportion of output volatility explained by risk premium shocks, however, is small. The implications of financial openness for output volatility in the presence of fiscal policy shocks are less clear-cut. We show that financial openness does affect the impact of fiscal policy on output volatility but the sign of this effect depends upon the specification of the model. Thus, only empirical tests can shed light on the sign of the link between financial openness and output volatility.

Using a panel dataset for OECD countries for the past 40 years, we test the implications of this model for the link between financial openness and output volatility empirically. We confirm the earlier literature in that we find no consistent link between openness and output

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[3] Evidence on the impact of the structure of financial markets on business cycle volatility seems to be more robust. Easterly et al. (2000), Denizer et al. (2000), and Da Silva (2001) find a more sophisticated financial system to be associated with lower volatility.

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 financial openness does not help to explain business cycle fluctuations.[4] In the 1980s and 1990s, monetary, fiscal, and terms of trade volatility help to explain output volatility. Furthermore, financial openness affected the impact of these variables in the 1990s, but not in the 1970s and the 1980s. More specifically, we find that financial openness magnified the impact of monetary policy on output volatility and that it cushioned the impact of fiscal policy on output volatility in the 1990s. These effects are consistent with the results of the theoretical model.

Combining these results, we conclude that (i) the implications of financial openness for output volatility depend upon the nature of the underlying shocks, and, (ii) the link between macroeconomic policy, financial openness, and output 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.

The remainder of the paper is organized as follows. In Section 2, we describe the theoretical model we use to illustrate how financial market integration may influence the impact of various macroeconomic shocks on business cycle volatility. Section 3 describes the dataset we use in our empirical analysis. Section 4 presents our empirical estimates, and Section 5 concludes.

2 Financial Openness and Business Cycle Volatility: A NOEM-View

Since the pioneering work of Obstfeld and Rogoff (1995, 1996), new open economy macro-models (NOEM) have become a standard tool for studying international macro issues.[5] One major advantage of NOEM models is that they provide explicit micro-foundations of dynamic general equilibrium open economy macro-models. Recently, NOEM models have also been

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[4] The finding that business cycle characteristics in the 1970s differ from those in later periods is consistent with work by Blanchard and Simon (2000) or Stock and Watson (2002).

[5] For recent surveys see Lane (2001) and Sarno (2001).

used to study the implications of the integration of international financial market for macroeconomic fluctuations. Sutherland (1996) has shown how the standard NOEM-model can be extended to analyze the implications of global financial market integration for the impact of monetary, fiscal, and productivity shocks on macroeconomic fluctuations. The main difference between the model advanced by Obstfeld and Rogoff and Sutherland's model is that the latter assumes that domestic and foreign bonds are imperfect substitutes. This makes Sutherland's model a natural candidate for illustrating the theoretical foundations of our empirical analysis.

To make this model more realistic, we modify it 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 follow McCallum and Nelson (1999) and incorporate a risk premium shock which allows analyzing the implications of autonomous financial market shocks on business cycle fluctuations under alternative assumption regarding the degree of financial openness. Third, we draw on the work by Ireland (1997) and Taylor (2001) and build into the model 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 (1996) model, the world is made up of two countries, Home and Foreign. Home and Foreign are of equal size. Each country is inhabited by infinitely-lived identical households. The households form rational expectations and maximize their expected lifetime utility. Home and Foreign households have identical preferences. 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 some monopoly power on the goods market, it treats the price it charges for its product as a choice variable. When changing the price of their product, firms incur menu costs, implying that 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.

2.1.1 Households

The expected lifetime utility of a domestic (Home) household is defined as $U_t = E_t \sum_{s=t}^{\infty} \beta^{s-t} u_s$, with $0 < \beta < 1$ being the households' subjective discount factor. 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:

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

where $\mu > 1$, $\sigma > 0$, $\varepsilon > 0$, and $\chi > 0$. The habit formation parameter lies in the interval $h \in [0,1)$. κ_t is a stochastic productivity shock which (measured in terms of deviations from the steady state) follows a first-order autoregressive process with AR(1) parameter ρ_k and standard deviation σ_k. In Eq. (1), C_t denotes a real consumption index, N_t is the households' labor supply, and M_t / P_t denotes the end-of-period real money holdings, where M_t are domestic nominal money balances (there is no currency substitution), and P_t is 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. These goods are sold by Home and Foreign firms in a monopolistically competitive goods market and are indexed by z on the unit interval, such that the aggregate consumption index can be expressed as:

$$C_t = \left[\int_0^1 c_t(z)^{(\theta-1)/\theta} dz \right]^{(\theta-1)/\theta}, \tag{2}$$

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, P_t, defined as the minimum expenditure required to buy one unit of the consumption good C_t, is given by $P_t = \left[\int_0^1 p(z)^{1-\theta} dz \right]^{1/(1-\theta)}$, 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 defined as the amount of domestic currency units required to buy one unit of the foreign currency and P_t^* denotes the aggregate foreign price level. Here, and in the following, an asterisk denotes a foreign variable.

2.1.2 Financial Markets

Households can hold internationally traded domestic and foreign nominal bonds. Whereas the standard NOEM model is based on the assumption that capital markets are perfectly integrated, Sutherland (1996) introduces real transaction costs of trading these bonds internationally. These transaction costs drive a wedge between domestic and foreign interest rates. In this paper, we add an additional cost component which ensures that the foreign asset position is stationary. We assume that the real transactions costs, Z_t, incurred by domestic households consist of two components:

$$Z_t = 0.5\psi_1 I_t^2 + 0.5\psi_2[(F_t - \overline{F})/P_t^*]^2, \tag{3}$$

where $\psi_1 > 0$ and $\psi_2 > 0$ are positive constants, F_t denotes the stock of foreign currency denominated assets held by Home households, and I_t denotes the level of real funds transferred by Home households from the domestic to the foreign bond market. Both Z_t and I_t are denominated in terms of the consumption aggregator, C_t. The first term on the right-hand side of Eq. (3) is identical to the transaction cost function used by Sutherland (1996). The second term on the right-hand side of Eq. (3) is new and represents a quadratic cost of holding a quantity of foreign bonds different from its long-run steady state level. This term implies that the foreign asset position and, thus, the steady state around which the model is log-linearized is stationary.[6] This property of the model will serve useful in the stochastic simulations of the model described in Section 2.1.5 below.

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 domestic and foreign bond holdings, and hold money. In

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[6] See also Neumeyer and Perri (2001) and the discussion in Schmitt-Grohe and Uribe (2001).

addition, they pay taxes. Consequently, the domestic bond holdings of domestic consumer can be described by the following difference equation:

$$D_t = (1+i_{t-1})D_{t-1} + M_{t-1} - M_t + w_t N_t - P_t C_t - P_t I_t - P_t Z_t + \Pi_t - P_t T_t, \tag{4}$$

where D_t denotes the quantity of domestic currency denominated bonds, i_t denotes the nominal interest rate on these bonds between period t and $t+1$, w_t is the nominal wage rate paid in a perfectly competitive labor market, Π_t denotes the profit income, and T_t stands for real taxes (denominated in terms of the consumption aggregator, C_t).

The dynamics of the domestic households' foreign bond holdings are given by:

$$F_t = (1+R^*_{t-1})(1+rp_t)F_{t-1} + P^*_t I_t, \tag{5}$$

where R^*_t denotes the nominal foreign interest rate paid for holding a foreign bond between period t and $t+1$, and rp_t is a stochastic risk premium shock. We assume that the stochastic risk premium shock (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}.

2.1.3 Firms

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

$$y_t(z) = (p_t(z)/P_t)^{-\theta}[C_t + C^*_t + G_t + G^*_t + Z_t + Z^*_t]/2. \tag{6}$$

The 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. Following Calvo (1983), firms therefore maximize the expected present value, $V_t(z)$, of current and future profits, where period-*s*, *s*>*t*, profits are weighted by the probability that the current period price, $p_t(z)$, will still be in force in period *s*. Firms, thus, maximize:

$$V_t(z) = E_t\{\textstyle\sum_{s=t}^{\infty} \gamma^{s-t} R_{t,s}\Pi_s(z)/P_s\}, \tag{7}$$

where $R_{t,s}$ is the discount factor. This Calvo-style price adjustment mechanism introduces price-stickyness and, thereby, dynamics into each firm's price setting decision.

Given the price of the differentiated good *z*, the quantity produced by the firm can be derived from the demand function for this good.

2.1.4 The Government

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

$$G_t = T_t + (M_t - M_{t-1}) / P_t, \tag{9}$$

where real government purchases are denominated in terms of the consumption aggregator, C_t.

In order to specify a stochastic process describing the dynamics of G_t, we assume that fiscal policy behaves according to a simple fiscal policy rule. In the case of the Home economy, this fiscal policy rule is given by (Taylor 2001):

$$\hat{G}_t = f\,\hat{y}_t + \rho_G \hat{G}_{t-1} + \varepsilon_{G.t}, \tag{10}$$

where $\rho_G \in [0,1]$, $\varepsilon_{G.t}$ 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 of Eq. (10) captures the influence of automatic stabilizers on the conduct of fiscal policy. In the stochastic simulations of the model presented in Section 2.1.5, we follow Taylor (2001) and set $f = -0.5$.

In conducting its monetary operations, the government may respond contemporaneously to the technology and the risk premium shocks. We adopt a money supply rule similar to the one suggested by Ireland (1997):

$$\hat{M}_t = \rho_M \hat{M}_{t-1} + \beta_1 \varepsilon_{k.t} + \beta_2 \varepsilon_{rp.t} + \varepsilon_{M.t}, \tag{11}$$

where $\varepsilon_{M.t}$ is a serially uncorrelated stochastic disturbance term with standard deviation σ_M.[7]

[7] Because the risk premium shock affects devaluation expectations and, thus, the international interest rate differential, the third term on the right-hand side of Eq. (11) can

2.1.5 Model Properties

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 Home and Foreign foreign asset positions are zero. In a second step, we use the algorithm developed by Klein (2000) to simulate the model numerically. In the numerical simulations, we assume that the innovation terms, $\varepsilon_{j,t}$, $j \in \{M, G, k, rp\}$ in the stochastic processes driving the Home and Foreign economies are perfectly negatively correlated, i.e., shocks are asymmetric. The calibration of the model is given in Table 1 and closely follows Sutherland (1996).

To analyze the dynamics properties of the model, we present in Graph 1 impulse response functions depicting the response of real domestic output to an unanticipated one unit money supply, government spending, productivity, and risk premium shock, respectively. To shed light on the impact of financial openness on the dynamics of the model, we compare in Graph 1 impulse responses that obtain when the degree of international capital mobility is high ($\psi_1 = 0$) and low ($\psi_1 = 5$).

Money supply shocks lead to a reduction of the domestic interest rate and a rise of the foreign interest rate.[8] Because of the negative international interest rate differential, households seek to accumulate foreign assets. Given the assumed sluggish price adjustment, the resulting nominal depreciation of the domestic currency also leads to a real depreciation, which, in turn, triggers an expenditure switching effect. This expenditure switching effect fosters the demand for domestic products. Domestic consumption demand is further stimulated by the decline in the domestic real interest rate. This leads to 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.

be thought of to capture the response of the money supply to deviations of the interest rate from its steady state value caused by financial market shocks. With such an interpretation, the monetary policy rule given in Eq. (11) would be a special case of the money supply rule used in, among others, McCallum (1983).

[8] Similar effects are at work in the case of fiscal policy, productivity, and risk premium shocks.

With international financial markets being imperfectly integrated ($\psi_1 > 0$), the evolution of the foreign bond holdings of Home and Foreign agents are also reflected directly in the uncovered interest rate parity condition. Neglecting the influence of the second component of the intermediation cost function, Z_t (Eq. 3), this direct effect of the foreign asset position on the international nominal yield differential is absent in a world of high capital mobility ($\psi_1 = 0$). The international nominal yield differential (corrected for the expected transaction costs of taking positions in international financial markets) is directly related to the expected rate of depreciation of the nominal exchange rate. With domestic agents accumulating foreign bonds, the expected rate of depreciation of the domestic currency is smaller in segmented international financial markets. Since the real exchange rate effect is smaller as well, it follows that the output effect of the monetary policy shock is smaller in the case of low capital mobility as compared to the case of high capital mobility. As a result, the output effects of monetary policy shocks are increasing in the degree of international capital mobility.

The impulse responses illustrated in Graph 1 show 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 of government spending and productivity shocks on output. Moreover, moving from imperfect to perfect financial market integration, the effect on output dynamics tends to be most significant in the case of money supply shocks and fiscal policy shocks.

Varying the degree of financial openness also changes the persistence of shock-induced deviations of output from its steady state value. Because this makes it difficult to infer from Graph 1 the implications of financial openness for output volatility, we report in Table 2 the results of stochastic simulations of the model. The table reports mean values of the standard deviation of output averaged over 100 simulation runs, with each simulation run pertaining to a sample consisting of 500 observations.

To get an impression of the robustness of the simulation results, we report results for three alternative policy regimes. In the first policy regime, monetary and fiscal policy follow the policy rules given in Eqs. (10) and (11). In the second policy regime, fiscal policy follows the policy rule given in Eq. (11) but the monetary policy rule in Eq. (11) is simplified by setting $\beta_1 = \beta_2 = 0$. In the third policy regime, the monetary policy rule is as given in Eq. (11) but the fiscal policy rule is assumed to be a simple AR(1) process ($f = 0$).

The simulation results show how financial openness affects output volatility (Table 2). In all three policy regimes considered, output volatility tends to be higher in a world of high capital mobility if monetary policy shocks hit the system. However, a higher degree of international capital mobility tends to cushion the output effects of fiscal policy shocks. Yet, we also note that this results drops only out of the benchmark simulations and out of the simulation in which the monetary policy rule is simplified by $\beta_1 = \beta_2 = 0$. In the third policy regime, in contrast, higher international capital mobility tends to magnify output volatility in the presence of fiscal policy shocks. Thus, financial openness alters the effectiveness of fiscal policy but the sign depends upon the specification of the model. This suggests that only empirical tests can throw light on the sign of the link between financial openness and output volatility caused by fiscal policy shocks. The simulation results further demonstrate that financial openness magnifies output volatility in the wake of risk premium shocks and productivity shocks. In this model, however, the proportion of output volatility explained by risk premium shocks is very small relative to that part explained by monetary and fiscal policy shocks. This small effect of risk premium shocks is also one explanation for the negative response of output volatility to greater financial openness if all types of shocks are considered simultaneously.

3 Stylized Facts

The empirical evidence presented in this paper is based on annual data for 24 OECD countries for the years 1960-2000. A full list of the countries included and the variables used is given in Table 3.

In contrast to earlier work, which makes use of the standard deviation of the growth rate of real GDP to approximate output volatility (Ramey and Ramey 1995, Martin and Rogers 2000, see also Table 4), we follow Mills (2000) who argues that this concept is not probate to measure business cycles. He argues that the underlying first difference filter removes frequencies from the data which are normally not attributed to business cycles. Rather, Mills (2000) suggests to implement time series filters like the one advocated by Hodrick and Prescott (1997) or Baxter and King (1999) to calculate the cyclical component of the time series under investigation. The volatility of a time series at business cycle frequencies is then measured as the standard deviation of the cyclical component of the time series.

One advantage of the bandpass filter proposed by Baxter and King (1999) is that it more closely tracks the true spectrum of a time series. It decomposes the underlying 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). A crucial question is what time span can be considered as the typical length of a business cycle. Baxter and King (1999) recommend to interpret fluctuations shorter than 2 years and longer than 8 years as the cyclical part of the time series. We follow this argumentation and use a bandpass (2,8) filter.

Table 5 shows the development of business cycle volatility, as measured in terms of real GDP volatility, and financial openness over time. Business cycle volatility has peaked in the second half of 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 business cycle 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.

Financial openness is proxied through gross and net foreign assets of commercial banks as measures of the openness of banking systems. 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 the openness of other financial market segments. Because the theoretical model presented above shows that gross and net foreign assets change in response to exogenous shocks hitting an economy, information on capital controls is used as an additional, 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.

Graph 3 plots our measures of financial openness in relation to business cycles volatility for the full sample. To take into account the possibility that correlation coefficients are influenced by some large outliers, we follow Mills (2000) and use robust statistics. In particular, the regression lines depicted in Graph 3 are calculated using a variant of the weighted least squares technique giving less weight to outlier observations. There is 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 output volatility and have even a small positive correlation to consumption volatility.

One reason for a global decline in business cycle volatility could be that volatility declines in some important countries and that, at the same time, the synchronization of business cycles increases. In this case, a global decline in volatility would be reflecting mainly increased correlations. In the literature, a relatively clear picture in fact emerges that the synchronization of business cycles in Europe has increased. However, studies disagree whether this trend has also been visible in the OECD region as a whole. Dalsgaard et al. (2002) look at data for 13 OECD countries for the years 1960-2000 and argue that the recent decline in business cycle volatility has not been due to an increased synchronization of cycles. This result would be in contrast to Bergman et al. (1998) who find that correlations across OECD countries have tended to increase over time. Graph 2 plots the standard deviations of the output gaps in our sample. This graph confirms Dalsgaard et al. (2002) in that the increased synchronization of business cycles seems to be a European rather than an OECD-wide phenomenon.[9] Hence, the decline in output volatility that many OECD countries have experienced does not seem to be driven by a greater co-movement of business cycles.

4 Determinants of Business Cycle Volatility

This section presents more systematic evidence on the determinants of business cycle volatility. Our analysis is based on a panel of non-overlapping averages of the time series under investigation, which allows both the cross-section and the time-series dimension of the data to be exploited. Since the volatility measure is defined over a certain span of time, an assumption on aggregation over time is warranted. Though the concrete choice of the time

[9] This conclusion also holds if emerging market countries such as Mexico, South Korea, and Turkey are excluded from the non-EU sample.

span is clearly arbitrary, it seems appropriate to choose a time span of five years, since five years are often seen as the typical length of a business cycle.[10]

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, since earlier work has suggested that the determinants of business cycle volatility in the 1970s and 1980s may have differed from those in the 1960s and 1990s (Stock and Watson 2002, Blanchard and Simon 2001), we finally use cross-section regressions for individual decades.

4.1 Granger Non-Causality Tests

In this section, we test for the presence of 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 check whether financial openness might be endogenous by testing 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 variables that we will use below to capture monetary and government spending volatility.

As in time series applications, tests for Granger non-causality based on a panel investigate whether a series x Granger-causes a series y. This is the case if the knowledge of x up to t-1 helps to predict the value of y in t. In panel applications, however, estimation problems which arise with dynamic empirical models must be addressed. The idea of Granger-non-causality in panels traces back to Holtz-Eakin, Newey, and Rosen (1988). These authors introduce the concept of panel-vector-autoregressions and consider models of the form:

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[10] For papers using similar non-overlapping five-year averages see Madsen (2002), Kneller

$$y_{t,i} = \alpha_0 + \sum_{j=1}^{m} \alpha_j y_{i,t-j} + \sum_{j=1}^{m} \delta_j x_{i,t-j} + f_i + u_{it}, \tag{12}$$

where i = 1,…,N denotes the number of cross-sections of the panel. By calculating first differences of the data, fixed effects can be eliminated:

$$y_{i,t} - y_{i,t-1} = \sum_{j=1}^{m} \alpha_j (y_{i,t-1} - y_{i,t-j-1}) + \sum_{j=1}^{m} \delta_j (x_{i,t-1} - x_{i,t-j-1}) + (u_{it} - u_{i,t-1}) \tag{13}$$

Within this model, x Granger-causes y if the joint hypothesis $\sum_{j=1}^{m} \delta_j = 0$ cannot be rejected. This assumes that the coefficients are equal across all cross-sections, i.e., that a stable causality pattern exists for the entire panel. Moreover, there are estimation problems since the residuals are by definition correlated with the endogenous variables. Hence, an instrumental variable estimator is warranted.

More generally, the problem of endogenous lagged right hand side variables has to be addressed. Whereas in a lot of applications the resulting so-called Nickell-bias can be neglected since 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).[11]

Results of these tests are reported in Table 6. 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 output volatility. Also, the Anderson/Hsiao estimator suggests that the volatility of short-term interest rates might be driven by changes in business cycle volatility.

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and Young (2001), and Beck and Levine (2001).

11 For a similar empirical strategy see Campos and Nugent (2000).

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 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 (proxied by the standard deviation of the growth in the terms of trade)[12] (see also Karras and Song 1996). We then estimate an equation of the form:

$$\sigma_{i,t}^{cycle} = \alpha_{0,i} + \alpha_{1,t} + \beta\sigma_{i,t}^{control} + u_{i,t} \tag{14}$$

In equation (14), business cycle volatility depends on country ($\alpha_{0,i}$) and time-fixed effects ($\alpha_{1,t}$), and on the volatility of the control variables.

Table 7 summarizes the results of three sets of regressions. First, we estimate a baseline specification with the control variables 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 Table 4). We use principal components to combine two measures of gross financial flows (banks' foreign assets and gross capital flows) into one measure of financial openness. There are, however, some problems with this approach. In particular, the first principal component is not necessarily the component with the highest correlation with the exogenous variable. Thus, if one chooses the first principle component to serve as a regressor, one runs the risk of missing a possible significant influence. To avoid this, the principle component with the highest correlation to the exogenous time series can be chosen. An additional problem is that the principal components are not invariant to the units of measurement of the time series under investigation. Thus, a minor change of the underlying time series can change, for example, the first principle component. Therefore, the time series are standardized before deriving the

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[12] In the panel regressions, we use the price of oil measured in US-dollar. 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.

principal components of financial openness. Using this approach, the first principal component that we obtain is highly correlated with the two variables and explains a high fraction of their variance.

Third, we interact financial openness with the standard deviation of short-term interest rates and government spending.[13] 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.

For the panel regressions, hardly any of the control variables is statistically significant (Table 7). Likewise, financial openness or the interaction between openness and our policy variables enters significantly only in one of the cases.

The poor explanatory power of our control variables and the missing evidence for a stable link between openness and business cycle volatility could be due to structural breaks in the data. Recent empirical evidence on the U.S. business cycle indeed suggests that there have been gradual changes in the determinants of 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 5 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 (Table 7). 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

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[13] Rose (1996) uses a similar approach to explain the volatility of exchange rates.

terms of trade has contributed to output volatility. For the 1990s, we are able to explain almost three-fourth of the cross-country variation in output volatility.[14]

There is some evidence for the hypothesis that the sign of the link between openness and business cycle volatility has changed over time. Whereas 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. Generally, the chaning 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 have been 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 provides significant coefficients , and the signs of these coefficients are in line with economic theory: in financially more open economies, the impact of monetary policy volatility increases, and the impact of the volatility of government expenditure declines. For interest rate volatility, a qualitatively similar result is obtained when including openness *and* the interaction terms in the panel regressions (not reported).

4.3 Robustness Checks[15]

A number of alternative specifications are considered to test the robustness of our results. Since the Granger tests for non-causality presented in Table 6 suggest that the volatility of short-term interest rates might be driven by business cycle volatility, we re-estimate the regressions reported in Table 7 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 estimate the cross-

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

[15] Not all of the following results are reported but are available from the authors upon request.

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

Pooling data for the 1980s and 1990s also provides results that are generally consistent with those obtained when looking at the two decades individually. 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 constant country sample.

Also, 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 8). Although the specifications differ because we cannot use the same methodology as before to measure volatility[16] and because we lack fully comparable data for all countries (see Table 7 for details), some of the qualitative results are similar. Although the coefficient estimates differ because of the 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.[17]

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 confirm our earlier results that the impact of financial openness depends on the type of 'shock' considered. We again find that the interaction between openness and the volatility of interest rates (government spending) has positive (negative) implications for output volatility. However, the latter effect

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16 More specifically, having only 10 or less years of data, we cannot use the bandpass filter but simply compute the standard deviation of the growth rates.

17 This share has been computed as the standard deviation of interest rates, multiplied by the coefficient estimate, and divided by the standard deviation of output volatility.

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 these results, output is less volatile in more developed countries.

5 Conclusions

Although conventional wisdom suggests that increased financial openness might amplify business cycle fluctuations, the literature has been relatively unsuccessful to 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 openness and volatility. Our results suggest that the link between macroeconomic policy, financial openness, and business cycle volatility may have changed over time. Using a panel dataset for OECD countries for the past 40 years, we find no consistent link between openness and output 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. Hence, parameter instability may explain why empirical studies have not been able to establish a statistically significant link between financial openness and business cycle volatility.

Second, the link between openness and 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 output 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 openness more rigorously, it would be necessary to identify monetary and fiscal shocks more precisely by looking at evidence from individual countries. Second, in this paper we focused on the question how financial openness may affect business cycle volatility as measured in terms of the volatility of real GDP. Because the implications of financial openness for real economic 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, it would be possible to extend the theoretical model studied in this paper by modeling the process of capital accumulation.

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Table 1 — The Calibrated Parameters

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 und Nelson (1999). The technology and the money supply processes are in line with the empirical estimates of Ireland (1997), who also provides an estimate of the contemporaneous response of monetary policy to productivity shocks. 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).

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 (0.0)	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 productivity process
σ_k	0.006	Standard deviation of the productivity shock
β_1	0.09	Contemporaneous response of monetary policy to productivity shocks
β_2	0.02	Contemporaneous response of monetary policy to risk premium shocks
f	–0.5	Impact of automatic stabilizers on government spending

Table 2 — Simulation Results

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 $\psi_1 = 5.0$ ($\psi_1 = 0.0$).

Regime / Shock	Low capital mobility		High capital mobility	
	σ_y	σ_c	σ_y	σ_c
	Benchmark simulation			
Money supply	0.0749	0.0438	0.2358	0.0435
Fiscal policy	2.0873	0.07193	1.7666	0.7436
Risk premium	0.0435	0.0301	0.1226	0.0289
Productivity	0.3550	1.1751	0.4045	1.2213
All	2.1402	1.4314	1.7937	1.4384
	Benchmark simulation + $\beta_1 = \beta_2 = 0$			
Money supply	0.0750	0.0438	0.2362	0.0437
Fiscal policy	2.0785	0.7156	1.7693	0.7447
Risk premium	0.0000	0.0000	0.0000	0.0000
Productivity	0.3769	1.1981	0.4423	1.2366
All	2.1555	1.4443	1.8123	1.4148
	Benchmark simulation + $f = 0$			
Money supply	0.0978	0.0323	0.2471	0.0253
Fiscal policy	5.0397	2.0924	5.1964	2.2236
Risk premium	0.0574	0.0213	0.1302	0.0167
Productivity	1.0696	0.8812	1.0583	0.8542
All	5.2588	2.2610	5.1347	2.3066

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).	IMF (2002), OECD (2001)
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: IMF (1998)
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)	IMF (2002)
	2) Banks' net foreign assets in percent of banks' total assets (absolute value)	IMF (2002)
	3) Deposit money banks' foreign assets in percent of banks' total assets (structural break for Ireland in 1980 adjusted manually)	IMF (2002)
interest rate	standard deviation of short-term interest rates (money market rate or alternative short-term lending rate)	IMF (2002)
government expenditure	standard deviation of rate of change of real non-wage government consumption	OECD (2001)
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	OECD (2001)
terms of trade	standard deviation of the change of the terms of trade	OECD (2001)
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, United States	

Table 4 — Selected Previous Empirical Results

Authors	Methodology	Results
Ceccetti and Krause (2001)	OLS regressions 23 OECD countries 2 periods (1982-89, 1990-97) dependent variable: macroeconomic performance as a weighted sum of output and inflation variability	Reductions in inflation and in output volatility are due to reduced state-ownership of banks and the introduction of explicit deposit insurance systems.
Denizer, Iyigun, and Owen (2000)	Quasi-panel, OLS regression 70 countries annual data 1956-1998 (divided into 4 time periods) dependent variable: standard deviation of real per capita income, investment, or income growth	Countries with more developed financial systems experience smaller fluctuations in real per capita output, consumption, and investment growth.
Easterly, Islam, and Stiglitz (2000)	Panel OLS regression 60-74 countries 2 periods (1960-78, 1979-97) dependent variable: growth volatility (standard deviation of the per capita growth rate)	Developing country dummy, trade share over GDP, and standard deviation of M1 growth have a positive impact on volatility. Non-linear effect of the ratio of private sector credit over GDP: level enters with a negative, squared term with a positive coefficient.
Karras and Song (1996)	Cross-Country regression 21 OECD countries dependent variable: growth volatility (standard deviation of the growth rate)	Volatility of money supply represents the monetarist interpretation of the cycle, volatility of investment demand and government spending is attributed to Keynesian explanations of business cycles. Variation of total factor productivity as a measure of supply side shocks to represent the real business cycle school. Business cycles are shown to be a combination of monetary, spending, and real shocks.
Razin and Rose (1994)	OLS and IV regressions 138 countries 1950-1988 dependent variable: standard deviations of de-trended consumption, investment, and output data	Distinguish between transitory and persistent and common versus idiosyncratic shocks. Different measures for the openness of the current and the capital account of the balance of payments have no impact on volatility. Inability to distinguish idiosyncratic from common shocks as a possible explanation.

Table 5 — Descriptive Statistics

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

	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 (%)	2.53	2.23	5.15	5.56

Table 6 — Tests on Granger Non-Causality

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) Arellano / Bond Estimator

Endogenous variable	Coefficient of lagged endogenous variable	Coefficient of lagged exogenous variable	Test of second order autocorrelation (Pr > z)	Sargan test of overidentifying restrictions ($\chi^2(20)$)	Result
		Volatility of GDP			
Capital controls	0.62***	0.03	0.68	39.10***	
	(6.66)	(0.60)			No causality
σ_Y	0.24*	0.11	0.57	26.31	
	(1.78)	(0.85)			
Openness	0.48***	0.09	0.46	24.66	
	(3.68)	(0.37)			No causality
σ_Y	0.26*	–0.09	0.43	23.95	
	(1.93)	(–1.29)			
σ_i	0.80***	0.04	0.17	52.90***	
	(–3.15)	(0.15)			No causality
σ_Y	0.29**	–0.02	0.26	29.58*	
	(2.15)	(–0.95)			
σ_G	0.01	–1.84	0.62	44.79***	
	(0.32)	(–1.27)			σ_X causes σ_Y
σ_Y	0.18	–0.01***	0.76	29.10*	(—)
	(1.41)	(–3.72)			

b) Anderson-Hsiao-Estimator

Endogenous variable	Coefficient of lagged endogenous variable	Coefficient of lagged exogenous variable	Test result
		Volatility of real GDP	
Capital controls	0.98	0.02	
	(0.34)	(0.20)	No causality
σ_Y	0.42	0.06	
	(1.01)	(0.29)	
Openness	2.72	–0.35	
	(0.43)	(–0.20)	No causality
σ_Y	1.25	–0.20	
	(1.59)	(–1.12)	
σ_i	–0.56***	0.29*	σ_Y causes σ_i
	(–5.75)	(1.66)	(+)
σ_Y	0.96	–0.04	
	(1.53)	(–1.02)	
σ_G	–1.29	0.12	
	(–0.98)	(0.17)	No causality
σ_Y	0.49	–0.01	
	(1.10)	(–1.10)	

Table 7 — Multivariate Regressions: OECD Countries

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 five-year non-overlapping averages (1960-1964, 1995-1970, …. 1995-2000). Country and time fixed effects are included. *** (**, *) = significant at the 1 (5, 10) percent level. Robust t-statistics in brackets.

	Panel Regressions	Cross–Section Regressions 1970s	1980s	1990s
		Baseline regression		
constant	0.93***	1.53***	0.91***	0.57***
	(5.97)	(3.84)	(6.73)	(4.79)
σ_i	0.08	0.06	0.06***	0.12***
	(1.18)	(0.37)	(7.30)	(7.72)
σ_G	–0.01	–0.002	0.004***	–0.004
	(–0.51)	(–0.14)	(3.22)	(–0.15)
σ_{TOT}	0.04	–2.84	1.43***	0.03***
	(0.99)	(–0.78)	(3.18)	(15.48)
R^2	0.27	0.04	0.34	0.78
N	77	19	21	24
		Including gross financial openness		
Constant	1.13***	1.66***	0.89***	0.64***
	(5.25)	(4.17)	(5.70)	(4.46)
σ_i	0.05	0.04	0.06***	0.09***
	(0.70)	(0.29)	(6.65)	(4.86)
σ_G	–0.01	–0.01	0.004*	0.05
	(–0.46)	(–0.86)	(1.97)	(1.29)
σ_{TOT}	0.003	–0.84	1.49**	0.03***
	(0.84)	(–0.23)	(2.76)	(8.17)
Openness	0.07	0.17	0.07	–0.14
	(1.19)	(1.36)	(0.41)	(–1.39)
R^2	0.28	0.08	0.35	0.81
N	77	19	21	23
		Including interaction terms		
Constant	1.00***	1.47***	0.90***	0.47***
	(6.96)	(3.61)	(5.63)	(3.65)
σ_i	0.03	0.09	0.05*	0.06**
	(0.48)	(0.60)	(1.85)	(2.64)
σ_G	0.004	–0.01	0.01	0.16**
	(0.25)	(–0.43)	(0.43)	(2.51)
σ_{TOT}	0.006	–0.81	1.54*	0.02***
	(1.36)	(–0.20)	(2.00)	(5.83)
Openness * σ_i	0.03	0.05	0.04	0.03*
	(0.96)	(0.50)	(0.38)	(2.05)
Openness * σ_G	0.01	0.01	–0.00	–0.07**
	(1.24)	(1.08)	(–0.02)	(–2.81)
R^2	0.31	0.09	0.36	0.83
N	77	19	21	23

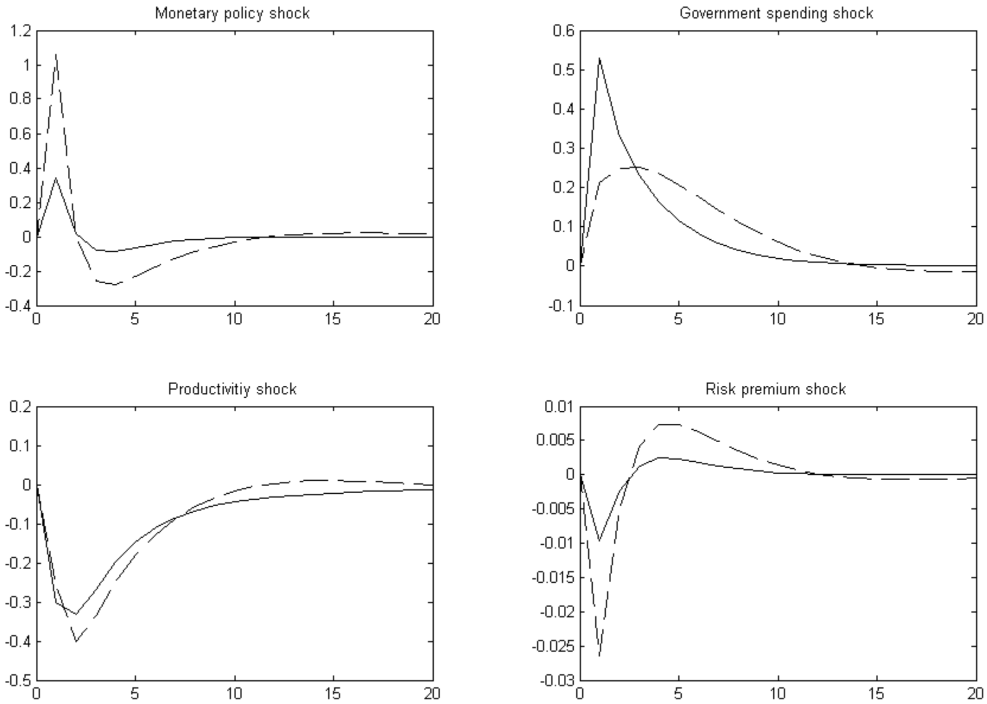
Table 8 — Multivariate Regressions: Extended Country Sample

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. *** (**, *) = significant at the 1 (5, 10) percent level. Robust t-statistics in brackets.

	(1) Baseline	(2) Including capital controls	(3) Including interaction with terms
constant	4.50** (2.14)	4.58* (1.85)	6.08*** (2.89)
σ_i	1.91** (2.15)	1.88** (2.09)	6.62*** (2.98)
σ_G	0.16** (1.94)	0.16* (1.79)	0.01 (0.07)
log (GDP per capita)	–0.28 (–1.42)	–0.29 (–1.24)	–0.52** (–2.48)
Openness		0.09 (0.12)	
Openness * σ_i			–4.38** (–2.03)
Openness * σ_G			0.18 (1.54)
R^2	0.29	0.28	0.32
N	77	74	74

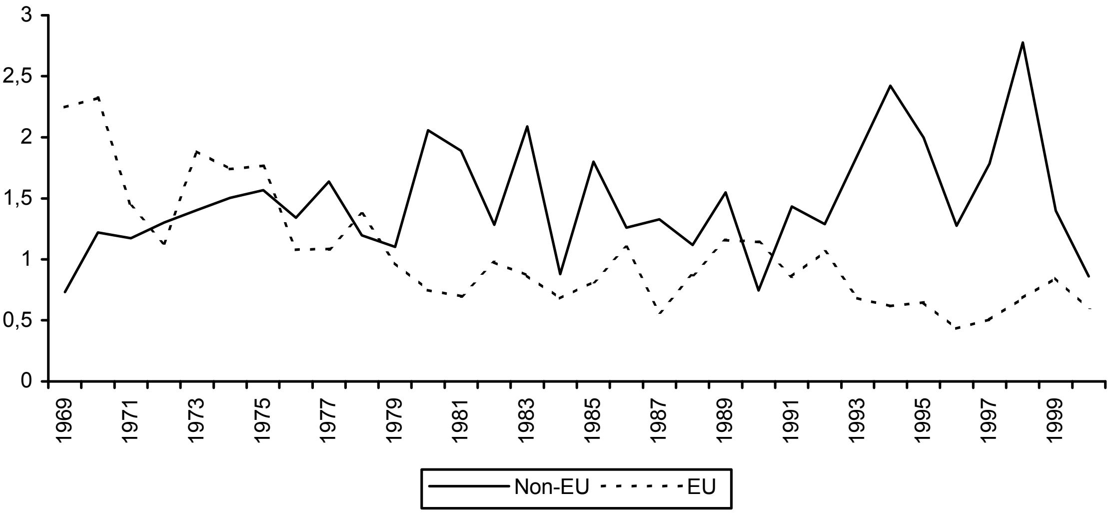
Graph 1 — Impulse Response Functions

The graph depicts the response of output to a one-unit monetary policy, government spending, productivity, and risk premium shock, respectively. Output is measured in terms of percentage deviations from its steady state value. Dashed lines obtain when setting $\psi_1 = 0.0$ and solid lines obtain when setting $\psi_1 = 5.0$.



Graph 2 — Standard Deviation of Output Gaps in OECD Countries

This graph plots the standard deviation of output gaps in OECD countries. The EU sample includes Austria, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Portugal, Spain, Sweden, the U.K. The non-EU sample includes Australia, Canada, Japan, Mexico, New Zealand, Norway, South Korea, Switzerland, and the U.S.



Graph 3 — Correlation Between Business Cycles and Openness