Macroeconomic Fluctuations in Developing Countries: Some Stylized Facts

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Abstract

This paper documents the main stylized features of macroeconomic fluctuations and business cycle regularities for a group of 12 developing countries. Cross-correlations between domestic industrial output and a large group of macroeconomic time series (including government revenue and expenditure, wages, inflation, money, velocity, private sector credit, international trade, and exchange rates) are presented. The effects of industrial country economic conditions on output fluctuations in these countries are also analyzed. The robustness of the results is examined using three detrending procedures: a modified Hodrick-Prescott filter, the Baxter-King band-pass filter, and a nonparametric technique. Overall, the results show some similarities between macroeconomic fluctuations in developing and industrial countries (e.g., procyclical real wages and countercyclical variation in government expenditure), but also some important differences in business cycle characteristics (e.g., countercyclical variation in the velocity of monetary aggregates).

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1 Introduction

Understanding and distinguishing among the various factors affecting the short- and long-run behavior of macroeconomic time series has been one of the main areas of research in quantitative macroeconomic analysis in recent years. Using a variety of econometric techniques, a substantial body of literature has documented a wide range of empirical regularities in macroeconomic fluctuations and business cycles across countries.[1] The documentation of stylized facts has often been used to provide an empirical basis for the formulation of theoretical models of the business cycle, and as a way to discriminate among alternative classes of models.

However, virtually all of the new research in this area has focussed on industrial countries, with relatively little attention paid to developing countries. At least two factors may help account for this. First, limitations on data quality and frequency could be constraining factors in some cases. For instance, quarterly data on national accounts are available for only a handful of developing countries; even when they are available, they are considered to be of significantly lower quality than annual estimates. Second, developing countries tend to be prone to sudden crises and marked gyrations in macroeconomic variables, often making it difficult to discern any type of "cycle" or economic regularities.

At the same time, there are a number of reasons why more attention to the documentation of the stylized facts regarding macroeconomic fluctuations in developing countries could be useful. Such an exercise could have important analytical value for examining whether similar empirical regularities are observed across countries at different levels of income. Differences from the types of reduced-form relationships observed in an industrial country context could provide an empirical basis for constructing analytical models for short-run fluctuations that incorporate features and relationships that are particularly important for developing countries. In addition, as argued by Agénor and Montiel (1996) and others, these findings may have important policy implications because they could be crucial, for instance, in the design of stabilization and adjustment programs.

Accordingly, this paper provides the first systematic attempt to document a wide range of regularities in macroeconomic fluctuations for a large

[1]For an overview of this literature, see, e.g., Backus and Kehoe (1992), Fiorito and Kollintzas (1994), and Van Els (1995). We will discuss below more specific results from this literature, in our attempt to compare business cycle characteristics across countries.

group of developing countries. The choice of countries in our sample was determined by various considerations. The first was the desire to select a group of countries for which data of reasonable quality could be assembled—thereby addressing the view that such exercises have limited validity due to data inaccuracies. The second consideration was the need to cover different geographic areas and a wide range of macroeconomic experiences, but also to select countries without substantial economic turmoil (in the form of, say, sustained episodes of hyperinflation) over the relevant sample period—thereby avoiding crisis-prone countries and the difficulties associated with data interpretation in such cases.[2] Moreover, by looking for a consistent set of relationships among macroeconomic variables in a relatively large group of countries that have undergone diverse experiences with structural change, we provide a set of stylized macroeconomic facts that are unlikely to reflect country-specific episodes.

Specifically, our study of business cycle regularities is based on quarterly data for a group of twelve middle-income countries: Colombia, Chile, India, Korea, Malaysia, Mexico, Morocco, Nigeria, the Philippines, Tunisia, Turkey, and Uruguay. On the one hand, the decision to use quarterly, rather than annual, data imposes an additional constraint on the size of our sample, because relatively few developing countries produce quarterly output indicators. On the other hand, using quarterly data provides us with sufficiently long time series for reliable statistical inference.[3]

The data cover a wide range of macroeconomic variables and include industrial output, prices, wages, various monetary aggregates, domestic private sector credit, fiscal variables, exchange rates, and trade variables. Thus, and in contrast to earlier studies, we are able to examine macroeconomic fluctuations in a number of different dimensions. In addition, we examine the relationship between economic fluctuations in these countries and two key indicators that proxy for economic activity in industrial countries—an index

[2]Of course, the notion of what represents "sustained economic turmoil" involves a value judgment on our part.

[3]There are two additional considerations in choosing quarterly, as opposed to annual data. First, some of the series used in this study have become readily available (and comparable across countries) for only a limited time. For instance, the data on effective exchange rates that we use have been published by the International Monetary Fund since 1978. Second, establishing large enough samples on an annual basis would imply going back to the early 1960s. It is reasonable to argue that the quality of the data, where available, was substantially lower in those earlier years.

of industrial country output and a measure of the world real interest rate.

There are two methodological aspects of the paper that are worth highlighting at the outset. First, in line with the recent literature on business cycles for industrial countries, many of the results discussed in the paper are based on unconditional correlations between different variables. We naturally recognize that such correlations do not imply causal relationships and, in some cases, attempt to complement our correlation results by examining bivariate exogeneity tests. We also recognize that reduced-form relationships between certain variables depend crucially on the sources of macroeconomic shocks. Nevertheless, our results are useful in that they provide an indication of the types of shocks that could be important for different countries and set the stage for more formal structural models of business cycle fluctuations.

Second, many of the macroeconomic series used in this paper have distinct trends over time and, hence, need to be rendered stationary prior to empirical analysis. Empirical results could, of course, be sensitive to the choice of the econometric procedure used to remove long-term trends from the data and derive cyclical components. This paper makes an additional methodological contribution by examining the sensitivity of correlations and other stylized facts to the detrending procedure used. We use three detrending techniques here: a modified version of the Hodrick-Prescott filter developed by McDermott (1997), the band-pass filter proposed by Baxter and King (1995), as well as a nonparametric detrending procedure.

The remainder of the paper is organized as follows. Section II examines the properties of the alternative filters used in the paper. Section III describes the data set, along with a number of economic features of the countries included in the sample, and presents summary statistics for the behavior of output. Section IV provides a more rigorous characterization of macroeconomic fluctuations in these countries, and contrasts the results with available stylized facts of business cycles in industrial and developing countries. Section V summarizes the main results of the paper. Section VI offers some final remarks and suggestions for further empirical and theoretical analysis.

2 Univariate Detrending Techniques

As indicated earlier, the objective of our paper is to examine economic fluctuations at business cycle frequencies rather than to study longer-term growth.[4] To do so, it is necessary to decompose all of our macroeconomic series into nonstationary (trend) and stationary (cyclical) components, because certain empirical characterizations of the data, including cross-correlations, are valid only if the data are stationary. As shown in Figure 1, for instance, the industrial output indices for the countries in our sample appear clearly nonstationary.

For a given series, in finite samples, stationary components obtained using different filters can often display very different time series properties (see, for instance, Canova, 1998). In this paper, we take a rather agnostic approach and report results obtained with three alternative filters: a modified version of the Hodrick-Prescott (HP) filter, the band-pass filter, and a nonparametric detrending method. We also discuss the robustness of our results to the choice of filter. In the remainder of this section, we provide a brief discussion of the main features of these three filters. Details concerning the formulation and implementation of these filters are provided in Appendix I.

The filtering techniques presented here can be discussed within the framework of a classical decomposition of a time series into its trend and cyclical components. Consider a seasonally-adjusted variable x_t that can be written as the sum of an unobserved trend component, x_t^*, and a residual cyclical component, x_t^c:

$$x_t = x_t^* + x_t^c. \tag{1}$$

2.1 The Hodrick-Prescott filter

The standard HP filter (see Hodrick and Prescott, 1997) employs an adjustment rule whereby the trend component moves continuously and adjusts gradually. Formally, the unobserved trend component x_t^* is extracted by solving the following minimization problem:

$$\underset{x_t^*}{Min} \sum_{t=1}^{T}(x_t - x_t^*)^2 + \lambda \sum_{t=2}^{T-1}[(x_{t+1}^* - x_t^*) - (x_t^* - x_{t-1}^*)]^2. \tag{2}$$

[4] It should be noted that, in the real business cycle literature, there is no clear distinction between trend and cycles since both short- and long-term fluctuations are regarded as being driven by the same stochastic process.

Thus, the objective is to select the trend component that minimizes the sum of the squared deviations from the observed series, subject to the constraint that changes in x_t^* vary gradually over time. The Lagrange multiplier λ is a positive number that penalizes changes in the trend component. The larger the value of λ, the smoother is the resulting trend series.

The HP filter has been subject to various criticisms (see Stadler, 1994, pp. 1768-69). In particular, it has been argued that it removes potentially valuable information from time series (King and Rebelo, 1993), and that it may impart spurious cyclical patterns to the data (Cogley and Nason, 1995).[5] More important for our purpose is the choice of the value of λ. The usual practice in the literature is to set λ to 1600 for quarterly time series. However, imposing this specific value (which was derived from an examination of the properties of US. output data) for each series and each country in our sample is somewhat arbitrary, and may reflect an overly stringent implicit assumption about the degree of persistence in x_t.

Our own approach here is to choose a value of λ for each individual series, using a data-dependent method. Specifically, we use a method of generalized cross-validation. The basic principle of cross-validation is to leave the data points out one at a time and to choose the value of the smoothing parameter under which the missing data points are best predicted by the remainder of the data. The virtue of this approach is that a priori assumptions about the appropriate value of the smoothing parameter are not required and the smoothing parameter does not have to be held constant across all series. Appendix I provides further details on the implementation of this filter (also see McDermott, 1997).

The estimates of the smoothing parameter from the application of the modified HP filter showed a wide range of variation across countries and across data series. For instance, for the industrial output series, the values for this parameter ranged from 380 to 5100 for the countries in our sample and was close to the "traditional" value of 1600 for only one country. We interpret these estimates as indicating the importance of choosing the smoothing parameter for the HP filter in a data-dependent manner, as we do in this paper, rather than mechanically using the same parameter across all series and for all countries.[6]

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[5]For a comparison of the HP filter with other detrending methods, also see Blackburn and Ravn (1991) and Canova (1991).

[6]Preliminary experiments indicated that the choice of the smoothing parameter can significantly affect the properties of the detrended series. For instance, over the range of

2.2 The band-pass filter

The band-pass filter, developed by Baxter and King (1995), is essentially a moving average that filters both high frequency "noise" and low frequency "trends", leaving behind fluctuations at the typical business cycle frequencies. The filter is constructed by combining a low-pass filter and a high-pass filter and imposing constraints that eliminate fluctuations at frequencies higher and lower than those corresponding to typical business cycle frequencies. Baxter and King (1995) argue that, for quarterly data, one should choose frequency cut-offs that correspond to 6 quarters and 32 quarters. They also note that the band-pass filter eliminates stochastic trends and deterministic polynomial time trends of order 2 or less. In our application of this filter, we use the same frequency cut-offs as Baxter and King.

2.3 Nonparametric method

This technique uses a univariate nonparametric regression estimation method to estimate the trend and cyclical components of a series without having to specify the functional form of the trend component of the underlying series or the degree of smoothing applied to the actual data. This is a flexible method in that it permits the modeling of trends that involve higher-order polynomials without imposing a particular functional form on the trend component. The method can also be extended to control for discontinuities or isolated change points in the series (that may be interpreted, for instance, as level shifts in the underlying series).

2.4 Trade-offs among filters

As discussed by Baxter and King (1995), there are at least five criteria by which univariate detrending filters are usually evaluated:

- the filter should extract a specified range of periodicities, and otherwise leave the properties of this extracted component unaffected;
- the filter should not introduce phase shifts;
- the filter should result in a stationary time series;

parameter estimates noted above, the implied cyclical volatility of U.S. postwar output differs by about 50 percent.

- the filter should yield business cycle components that are unrelated to the length of the sample period;
- the filter should be operational in the sense that it should be easy to use so that different researchers can easily reproduce the same result when using identical data sets.

There is an important trade-off involved in using alternative filters; while all filters are constructed to generate stationary time series by extracting low frequencies from the data, the band pass filter has the advantage of extracting both low and high frequencies that lie outside the range of typical business cycle frequencies. The ideal band-pass filter can be better approximated with the longer moving averages, but adding more leads and lags also means that observations must be dropped at the beginning and end of the sample thus leaving fewer observations for analysis. Both the HP filter and the nonparametric method do not lose observations at the end of the sample but, instead, generally have other end of sample problems. It is well known that the HP filter provides poor estimates of the trend component towards the end of the data sample. This problem seems to be also true of the nonparametric method. In addition, the method of choosing the order of the moving average in the band pass filter is somewhat arbitrary, as is the method of usually choosing 1600 as the smoothing parameter in the HP filter. The advantage of the nonparametric method and the modified HP method used in this paper is that they employ an automatic selection procedure based on objective criteria to determine the smoothing parameter (or bandwidth). As noted above, the nonparametric method can also be adapted to control for isolated change points in the underlining series.

The magnitude of the differences among the cyclical components obtained using various filtering techniques is, of course, an empirical matter. Using postwar U.S. quarterly industrial production data as a benchmark, we found average bivariate correlations of about 0.7 between the cyclical components obtained using the BP, HP, and NP filters. For the developing countries examined here, these correlations were generally much smaller, suggesting that the macroeconomic relationships studied below could be sensitive to the choice of detrending procedure. Hence, to examine the robustness of the results, in our empirical work we employ all of the filters that were discussed in this section. In addition, following the tradition in the business cycle literature, we also use the first-difference filter. However, we use four-quarter (logarithmic) differences instead of quarterly differences for two reasons. First, this provides a check on whether annual variations are significantly different from higher frequency (that is, quarterly) relationships in the data. Second, this enables comparison with earlier country-specific studies that have generally used annual data.

3 The Data

In this section, we describe a number of important economic features of the developing countries in our sample that are relevant for the analysis in this paper. In addition, we present summary statistics for output and inflation and provide a preliminary characterization of business cycle fluctuations in our group of countries. We also compare the properties of business cycles in these countries with those observed in industrial countries. The sample period for most of the data series used in this study goes from 1978:Q1 to 1995:Q4. The data sources are described in detail in Appendix II.

Figure 2 contains information on a number of key characteristics of the economies in our sample. As the figure suggests, most of the countries in our sample could be reasonably characterized as middle income countries. Although India and Nigeria have relatively low per capita incomes, we have included them in the sample because they are among the largest market economies in their respective continents. The urbanization rates and the proportions of agricultural output as a share of total GDP indicate that agriculture is an important, but not dominant, sector in most of these economies.

Because we were unable to obtain reliable quarterly GDP data for all the countries in our sample, we use indices of industrial output for constructing measures of the aggregate business cycle.[7] As shown in Figure 2, the manufacturing sector accounts for a significant fraction of total GDP in these countries. Except in Nigeria, this share is over 15 percent for all countries in our sample, compared to an average share of about 25 to 30 percent in most industrial economies. In addition, because industrial sector output roughly corresponds to output in the traded goods sector (excluding primary com-

[7]It might also be argued that the use of GDP data for measuring business cycle activity in a developing-country context can be problematic. Agriculture, which still accounts for a large share of aggregate output in many cases (including several countries in our sample), is more influenced by weather conditions than cyclical factors. Poor measurement of services and informal sector activities may also impart significant biases.

modities) and is most closely related to what are traditionally thought of as business cycle shocks, either exogenous or policy-determined, we would argue that this variable is a reasonable proxy for measuring the aggregate cycle.

Figure 2 also provides information about a number of other economic indicators, such as government expenditure and revenue, average growth in imports and exports, and external debt service ratios. For all countries except Nigeria, export growth is an important contributor to overall GDP growth. Standard measures of openness to international trade—as indicated by the average openness ratio, defined as the ratio of the sum of imports and exports over GDP—illustrate the importance of foreign trade transactions for these countries. Hence, an important part of our analysis will focus on the relationship between the domestic business cycle and prices and quantities related to international trade.

As indicated in the introduction, an important consideration in choosing our sample (in addition to data availability) was that we wanted to exclude countries that had undergone sustained episodes of hyperinflation over the period under study. Figure 2 suggests that, although some of the countries in the sample (such as Mexico, Turkey, and Uruguay) have had high levels of inflation over the past two decades, none of these countries had sustained episodes of hyperinflation during this period. This is also apparent from the last two columns of Table 1, which show average annual rates of consumer price inflation and also the volatility of inflation, as measured by the standard deviation of annual inflation rates.

A key issue at this juncture relates to the nature of business cycle fluctuations in developing countries. In particular, are aggregate fluctuations in these economies characterized by basic time-series properties—such as volatility and persistence—that are similar to those observed in industrial countries? A simple way of approaching this issue is to examine summary statistics for the stationary components of industrial output. The second panel of Table 1 reports means and standard deviations of output growth rates as well as standard deviations of the cyclical components of output derived using the three filters described in the previous section.[8] As noted before, growth rates are measured here as four-quarter differences of the log levels of the relevant variables.

[8]These filters, by construction, deliver stationary components that have zero mean. The output series as well as all the other time series used in this paper were deseasonalized using the X-11 procedure.

The first column of Table 1 indicates that mean annual growth rates of industrial output over the last two decades have varied substantially across the countries in our sample, ranging from almost 14 percent for the Philippines to about 2.5 percent for Colombia, Mexico, Morocco, and Tunisia. Uruguay, in fact, recorded a negative mean growth rate of industrial production over this period.[9] The volatility of growth rates also varies markedly across countries. On average, volatility in our group of countries is much higher than the level typically observed in industrial countries. These results are in line with those obtained in various other recent studies of business cycle fluctuations in developing countries, most notably Mendoza (1995).

A similar picture emerges from an examination of the standard deviations of the filtered cyclical components of industrial output.[10] Because the filters used here tend to eliminate more of the low frequency variation than the growth filter, these standard deviations are generally lower, although the ordering of countries in terms of cyclical volatilities is quite similar and, in general, these volatilities are higher than those observed for industrial countries. An interesting point to note is that the volatility of the cyclical components obtained using the BP filter is almost always much lower than those from the HP and NP filters. This is attributable to the fact that the BP filter also eliminates some of the very high frequency variation in the data, unlike the latter two filters that eliminate only low frequency variation.

To examine the persistence of business cycle fluctuations, Table 1 also reports the first four autocorrelations of the filtered series. A striking aspect of these results is that the autocorrelations are generally strongly positive, indicating considerable persistence in the cyclical components. We interpret these results as suggesting that it is appropriate to view these developing countries as having short-term fluctuations that could be reasonably characterized as business cycles.

[9]This is also evident from figure 1, which shows that the level of industrial output in Uruguay is higher at the beginning of the sample than at the end.

[10]These standard deviations are interpretable as quarterly percentage standard deviations. For purposes of comparison, the standard deviation of HP-filtered postwar quarterly industrial production for the U.S. is about 2 percent.

4 Main Features of Macroeconomic Fluctuations

We measure the degree of comovement of a series y_t with industrial output x_t by the magnitude of the correlation coefficient $\rho(j)$, $j \in \{0, \pm 1, \pm 2, ...\}$. These correlations (as reported in the tables discussed below) are between the stationary components of y_t and x_t, with both components derived using the same filter. In line with much of the literature (see, for instance, Fiorito and Kollintzas (1994)), a series y_t is said to procyclical, acyclical, or countercyclical, depending on whether the contemporaneous correlation coefficient $\rho(0)$ is positive, zero, or negative. In addition, we deem the series y_t to be strongly contemporaneously correlated if $0.26 \leq |\rho(0)| < 1$, weakly contemporaneous correlated if $0.13 \leq |\rho(0)| < 0.26$, and contemporaneously uncorrelated with the cycle if $0 \leq |\rho(0)| < 0.13$.[11]

The cross-correlation coefficients $\rho(j)$, $j \in \{\pm 1, \pm 2, ...\}$ indicate the phase-shift of y_t relative to the cycle in industrial output. Again, in line with the existing literature, we say that y_t leads the cycle by j period(s) if $|\rho(j)|$ is maximum for a negative j, is synchronous if $|\rho(j)|$ is maximum for $j = 0$, and lags the cycle if $|\rho(j)|$ is maximum for a positive j.

4.1 Correlations with industrial country business cycles

In this subsection, we examine the relationship between domestic industrial output fluctuations in the countries in our sample and variables that represent economic activity in the main industrial countries, a relationship that could be particularly important for countries that have substantial trade links with industrial economies. As discussed earlier, the magnitude of the links between macroeconomic fluctuations in industrial and developing countries and the channels through which shocks propagate across these two sets of countries are of considerable interest from a number of different perspectives.

Table 2 reports the correlations between domestic industrial production and a composite index of industrial production in the main industrial

[11] The approximate standard error of these correlation coefficients, computed under the null hypothesis that the true correlation coefficient is zero, and given the average number of observations per country in our sample, is about 0.13.

economies.[12] The contemporaneous correlations are positive for a majority of the countries in the sample, indicating that business cycle fluctuations in developing economies tend to be correlated with fluctuations in industrial country business cycles. For many of the countries that have positive contemporaneous correlations, the correlations generally peak at or near lag zero, suggesting that output fluctuations in industrial economies are transmitted fairly quickly to these countries.[13] These results are generally robust across filters, barring a couple of exceptions. For instance, in the case of Mexico, the BP filter alone yields a strong negative contemporaneous correlation, whereas the other filters yield positive correlations. The correlations at lag 4 are, however, all strongly positive, indicating a lagged effect of industrial country output on Mexican output. The contemporaneous correlations are close to zero for Morocco and Nigeria, and marginally negative for Turkey. For these countries, there is some evidence that industrial country output appears to have a positive effect on domestic industrial output with a lag of about four to eight quarters.

Next, we explore one additional channel by which business cycle conditions in industrial economies could influence fluctuations in developing economies. The world real interest rate is regarded as likely to have an important effect on economic activity in the developing world, not only because it affects domestic interest rates but also because it reflects credit conditions in international capital markets. These capital markets could be especially important for those countries (even in the middle-income range) that do not have well-developed domestic capital markets. To examine this issue, we report in Table 3 correlations of industrial output with a weighted index of real interest rates in the major industrial countries.

For a few countries in our sample, the contemporaneous correlations between HP-filtered output and the world real interest rate are in fact positive. This could reflect the fact that the real interest rate in industrial economies tends to be procyclical and changes in industrial country output,

[12] The construction of the industrial country variables used in this section is described in Appendix II. In this and all other tables where cross-correlations are reported, lag j indicates the correlation between the contemporaneous value of domestic industrial production and the j'th lag of the second variable (for instance, the index of industrial country output in Table 2). A negative lag denotes a lead.

[13] Business cycles in the industrial economies are, of course, not perfectly synchronized but Lumsdaine and Prasad (1997), among others, argue that there is a substantial common component in business cycle fluctuations across the main industrial economies.

through trade links, have positive spillover effects on industrial output in these middle-income countries.[14] Morocco and Turkey are the only countries in our sample where this correlation is negative using any filter. For a number of countries, the lagged correlations are negative, indicating a lagged effect of the world real interest rate on domestic industrial output. An interesting case is that of Mexico, where the contemporaneous correlation is positive but most of the correlations at short leads and lags are close to zero, indicating that the effects of changes in the world interest rate are transmitted to Mexican industrial output quite rapidly. This is not surprising given the physical proximity and close trade links between Mexico and the United States, which is the dominant industrial economy and therefore has a large weight in the composite indices of industrial country output and our proxy for the world real interest rate.

Overall, these results suggest that the level of economic activity in industrial countries has a positive but relatively weak influence on industrial output in the middle-income countries in our sample. The procyclical behavior of real interest rates in industrial countries may imply that the relationship between these interest rates and developing country industrial output is muted by the indirect opposite effect of aggregate economic activity in industrial countries. The correlations we have presented indicate the need for further work to separate out the quantitative importance of these different influences on business cycle propagation. An important issue in this context (which we will return to later) is the measurement problem caused by the absence of data on country-specific risk premia in measuring interest rates that individual countries face on world capital markets.

4.2 Cyclical behavior of public sector variables

We now turn our attention to the relationship between the business cycle and various domestic quantity and price variables that could be related to short-term output fluctuations. The relationship between fluctuations in aggregate output and the components of aggregate demand has been well documented for industrial countries. Unfortunately, we were unable to obtain consistent and reliable time series data on consumption and investment for all the countries in our sample. One set of variables for which we were

[14]Over the period 1975-95, the correlation between the cyclical components of the output and real interest rate indices for industrial countries is strongly positive, irrespective of the detrending procedure used.

able to obtain data, although only for a limited set of countries, relates to the public sector. Examining the relationship between aggregate economic activity and public sector expenditure and revenues has analytical value from the perspective of business cycle modeling and is also of importance from a policy perspective, including in the design of macroeconomic stabilization programs.

The top panel of Table 4 shows that there is a robust negative relationship between government expenditure and the domestic business cycle in all four countries for which we have data available—Chile, Korea, Mexico and the Philippines. Thus, there is fairly clear evidence of a countercyclical role for government expenditure in these countries. These results are in contrast to those obtained for industrial countries (for government consumption rather than government spending) by Fiorito and Kollintzas (1994) for instance, which suggest no clear pattern. The negative contemporaneous correlation between government consumption expenditure and industrial output is consistent with the prediction of a variety of models, such as, for instance, the class of intertemporal optimizing models with imperfect capital mobility and flexible prices discussed by Agénor (1997). In such models, an increase in public spending leads to a net increase in domestic absorption (if the degree of intertemporal substitution in consumption is not too large), a real exchange rate appreciation, and a fall in output of tradables on impact.

The contemporaneous correlations in the second panel of Table 4 show that government revenues are significantly countercyclical in Colombia, Korea, the Philippines and Uruguay.[15] This negative correlation may result from the negative effects of increases in tax revenues (possibly induced by increases in effective tax rates) on disposable income and aggregate demand.[16] In Mexico, the relationship appears to be acyclical, although this result is sensitive to the choice of filter. To examine the net effect of government revenue and expenditure on the domestic business cycle, we constructed a measure of the fiscal impulse for the three countries for which both the revenue and expenditure series were available. The fiscal impulse is defined as the ratio of government spending to government revenue. This variable is negatively correlated with the business cycle, either contemporaneously or at

[15]A countercyclical pattern of government revenue was also established by Rodríguez-Mata (1997) for Costa Rica.

[16]A close short-run correlation between current income and expenditure patterns in developing countries has been well documented, and is attributed to the existence of liquidity constraints and finite horizons. See Agénor and Montiel (1996, Chapter 3).

short lags, in Korea, Mexico, and the Philippines, indicating that the fiscal impulse measure is countercyclical and plays a role in short-run macroeconomic stabilization.

To summarize, the correlations examined in this subsection suggest that the government balance does play a significant role in dampening domestic fluctuations in countries such as Korea, Mexico and the Philippines. However, the countercyclical behavior of government revenues in some countries indicates the need to re-examine revenue sources in order to ensure that they do not exacerbate domestic fluctuations. An alternative possibility is that a tightening in government finances could lead to increases in future output growth by, for instance, "crowding in" private investment and by signaling the future stability of domestic macroeconomic policy, thereby stimulating foreign investment. Based on the negative lagged correlations, there is some evidence of this effect in our sample for Korea.

4.3 Correlations with wages and prices

In this section, we examine the cyclical behavior of wages and prices. Establishing stylized facts here has important implications for discriminating among different classes of models based on their predictions concerning the cyclical behavior of these variables. For instance, Keynesian models imply that real wages are countercyclical while equilibrium models of the business cycle imply that real wages are procyclical (Abraham and Haltiwanger, 1995). Similarly, the implications of the cyclical behavior of prices, inflation (and, as discussed subsequently, various monetary aggregates) for discriminating among different classes of business cycle models have been the subject of considerable debate in the business cycle literature recently (Chadha and Prasad, 1994). Hence, it is of interest to extend this set of stylized facts to the countries in our sample.

We begin by examining correlations between average nominal wages in the industrial sector and industrial output. Consistent time series data on wages were available for only five of the twelve countries in our sample. As shown in the upper panel of Table 5, the cyclical behavior of nominal wages varies markedly across these countries. In Chile, nominal wages appear to be procyclical. The results are not robust across filters for the other countries, although there is limited evidence of countercyclical movements in nominal wages in Colombia and Mexico.

In interpreting these results, it is also useful to look at the cyclical behavior of real wages, constructed by deflating nominal wages by the consumer price index. As indicated earlier, the real wage is often the relevant wage variable from the perspective of business cycle analysis. In this regard, alternative theories offer different predictions. For instance, traditional Keynesian models of the business cycle posit short-run movement along a stable labor demand schedule and, therefore, predict that real wages are countercyclical. Real business cycle (RBC) models, as well as new Keynesian macroeconomic models with imperfect competition and countercyclical markups, on the contrary, predict procyclical wages.[17] Finally, efficiency wage models predict no tight contemporaneous relationship between output (employment) and real wages. More generally, as noted for instance by Abraham and Haltiwanger (1995, p. 1230), shocks of different types can have very different implications for the cyclicality of the real wage. Technology shocks will tend to produce procyclical real wage behavior, whereas nominal shocks (such as money supply shocks) will generate countercyclical real wage movements.

The lower panel of Table 5 reports correlations between industrial output and real wages. These results are striking. For all five countries for which data are available, and across all filters, we find strong evidence of procyclical real wage variation, consistent with the implications of RBC models that ascribe a dominant role to technology shocks that shift the labor demand schedule in the short run and in line with the evidence for the United States provided, for instance, by Kydland and Prescott (1994).[18]

Next, we turn to the correlations between prices and output. A substantial literature has documented the countercyclical behavior of prices in industrial economies (see, for instance, Backus and Kehoe (1992), Fiorito and Kollintzas (1994), Kydland and Prescott (1994), and Cooley and Ohanian (1991)). Many of these papers have argued that the countercyclical behavior of (the level of) prices provides support for supply-driven models of the business cycle, including RBC models that attribute a predominant

[17]See Rotemberg and Woodford (1991) for a discussion of this class of models. Rotemberg and Woodford offer several hypotheses that could help explain why markups might vary countercyclically. For instance, the price elasticity of demand may be positively related to the level of sales. Alternatively, firms may be tempted to reduce markups in the expansionary phase of the business cycle in order to expand their customer base.

[18]Our analysis of real wage cyclicality only considers the consumption wage, and not the producer wage, in the manufacturing sector. The two measures could display very different behavior over time, as illustrated for the United States by Abraham and Haltiwanger (1995).

role to technology shocks in driving business cycle fluctuations. However, Chadha and Prasad (1994) have argued that the correlation between *inflation* and cyclical output is the appropriate correlation for discriminating between demand- and supply-driven models of the business cycle. They document that inflation has in fact been procyclical during the postwar period in the G-7 economies. We therefore examine the cyclical behavior of both the price level and the inflation rate.

Table 6 reports correlations between industrial output and the aggregate consumer price index. The contemporaneous correlations are generally negative for Colombia, India, Korea, Mexico, Morocco, Nigeria and Turkey, indicating countercyclical variation of the price level. For many of the other countries including Chile, the Philippines, Tunisia, and Uruguay, the correlations are significantly positive. Thus, unlike in the case of industrial countries, there does not appear to be a consistent negative relationship between the stationary components of the levels of output and prices for the countries in our sample.

Having examined price-output correlations, we turn next to the correlations between the level of inflation and alternative measures of the cyclical component of output.[19] The contemporaneous correlations in Table 7 indicate that it is difficult to find strong evidence of procyclical inflation for the countries in our sample. The correlations at the leads also do not provide a clear indication of a positive relationship between output and lagged inflation. Indeed, for some countries such as Mexico and Turkey, we find negative correlations between inflation and the cyclical component of output, indicating countercyclical variations in inflation.

Our interpretation of the results in this section is that supply rather than demand shocks appear to be the dominant influence on high frequency macroeconomic fluctuations in this group of middle-income countries. For instance, for Mexico and Turkey, the procyclical behavior of real wages and the countercyclical behavior of both the price level and the inflation rate provide strong evidence that supply shocks have been a key determinant

[19]As discussed earlier, our sample of countries was chosen in a manner that excluded countries with sustained hyper-inflationary episodes. However, unit root tests for inflation indicated that, for about half of the countries in the sample, we could not reject the null hypothesis of nonstationarity. Hence, for all countries, we detrended inflation using the same filters as for output. We also examined the correlations using the raw series for inflation and filtered output. For the countries for which we report significant correlations in Table 7, the choice of filtered or unfiltered inflation did not matter.

of domestic macroeconomic fluctuations over the last two decades. It is worth emphasizing here that, for this group of countries, the term "supply shocks" could have a different connotation than it does for large industrialized economies. In particular, these developing countries could be subject to large terms-of-trade shocks rather than prototypical productivity shocks—although it should be noted that terms-of-trade shocks could, in principle, have both supply-side and demand-side effects.

4.4 Money and credit

To further analyze the relative importance of different types of shocks on macroeconomic fluctuations, we now examine the cyclical behavior of a set of monetary variables. In recent years, it has become increasingly evident that equilibrium business cycle models can and often need to incorporate a role for monetary variables to capture important business cycle phenomena. The relationship between monetary variables and the business cycle has, therefore, become a topic of increasing interest (see, for instance, Kydland and Prescott, 1994). This is of particular relevance to middle-income countries where the monetary mechanism could play a potentially important stabilization role.

A large literature has evolved around the question of whether monetary variables influence output in industrial countries or, in more loaded terminology, whether money causes output. From a different perspective, King and Plosser (1984) have argued that positive correlations between money and the business cycle largely reflect the endogenous response of inside money to exogenous shocks that drive business cycle fluctuations rather than indicating a causal relationship from money to output. Given this debate, and because it is unclear what definition of money corresponds precisely to the concept used in theoretical models, we examined money-output correlations using a number of alternative definitions of monetary aggregates.

Table 8 reports correlations between industrial production and an index of broad money. This latter variable roughly corresponds to the definition of M2 for industrialized economies. Although in some cases the sign (and statistical significance) of the correlations is affected by the detrending procedure, the contemporaneous correlations are broadly positive for a majority of the countries, including Chile, Colombia, India, Morocco, the Philippines, Tunisia, Turkey, and Uruguay. Among the remaining countries, the contemporaneous correlations are often close to zero, although in the cases of Korea and Malaysia (and, possibly, Mexico), there is some evidence of countercyclical variation in broad money.

Among the countries that have positive correlations between money and output, the pattern of lead-lag correlations and, in particular, the lag at which the peak positive correlation occurs, could be interpreted as indicating the speed with which innovations in monetary variables are transmitted to real activity. For these countries, as shown in Table 8, the peak positive correlations generally occur at very short lags, suggesting that the transmission of monetary shocks to real activity is fairly rapid in these economies. Of course, as noted earlier, this could simply reflect the endogenous response of money to output fluctuations that are driven by non-monetary shocks. Indeed, when we ran bivariate Granger-causality tests between these two variables, we found little evidence that money "causes" output (in the Granger-causal sense) even in those countries where the correlations between the two variables were strongly positive.

The patterns of correlations were similar when we used two alternative monetary aggregates—reserve (or base) money and narrow money (currency in circulation plus sight deposits in the banking system).[20] The main features of the results in Table 8 were preserved when using the other monetary aggregates. The contemporaneous correlations were positive for about half of the countries in the sample, generally statistically insignificant for many of the others, and, in the case of Nigeria, clearly negative. Overall, therefore, we find limited evidence for the countries in our sample of the type of procyclical behavior of monetary aggregates that has been documented for many industrial countries; see, for instance, Backus and Kehoe (1992).[21] More importantly, we were unable to detect evidence of Granger causality from money to output. These results suggest to us the need for a very different analytical framework for studying the relationship between monetary policy and macroeconomic fluctuations in developing countries (see the discussion below).

We also examined the cyclical behavior of measures of velocity corresponding to the alternative definitions of monetary aggregates discussed above. Again, to conserve space, we present only the results for the measure

[20]To conserve space, these results are not presented here but are available from the authors upon request.

[21]The results of Fiorito and Kollintzas (1994) indicate that the correlation between money and output varies substantially across both countries and definitions of the money stock.

of velocity based on broad money.[22] These correlations, shown in Table 9, are striking. For ten of the twelve countries in our sample, and independent of the filter used, the contemporaneous correlations between the velocity of broad money and industrial output are strongly negative. The only exceptions are Turkey, where the results are sensitive to the choice of filter, and Mexico, where the correlations are close to zero. From a quantity theory perspective, of course, the countercyclical behavior of velocity would be expected given the procyclical behavior of broad money and countercyclical variation in the aggregate price level in a majority of the countries. This result stands in sharp contrast to the weakly procyclical behavior of velocity in the G7 economies that has been documented by Fiorito and Kollintzas (1994).

Finally, we consider another monetary variable that could have a significant influence on economic activity—domestic private sector credit. This is especially relevant for middle-income countries where equity markets tend to be relatively weakly capitalized and private sector credit typically plays an important role in determining investment and the financing of working capital needs and, thus, overall economic activity, especially in the industrial sector.[23] Hence, we now examine correlations between industrial sector output and the level of domestic private sector credit. It should be noted that changes in credit could partly reflect the derived demand for credit that could be affected by exogenous shocks influencing the level of industrial activity. Nevertheless, even in these circumstances, changes in the availability of credit could dampen the effects of these shocks on industrial output. Thus, the pattern of these correlations is still of considerable analytical value.

Table 10 shows that, for a number of countries, including Colombia, India, Mexico, and Turkey, there is a positive contemporaneous association between domestic credit and industrial output, although the strength of this relationship is not always robust to the choice of detrending procedure. In many other countries, most notably in Chile and Uruguay, there is a negative correlation between these two variables. In those countries where the asso-

[22]Measures of velocity corresponding to the reserve money and narrow money aggregates yielded velocity-output correlations that were broadly similar to the results discussed in this paragraph. These results are available upon request.

[23]Although the role played by private sector credit in many developing countries is well documented, few studies have provided a systematic quantitative assessment of the relative importance of money and credit in the transmission of monetary policy in these countries. We intend to pursue this issue in future work.

ciation is positive, the correlations peak at or close to lag zero, indicating that the availability of domestic credit affects activity in the industrial sector fairly rapidly. However, as noted earlier, this could simply reflect cyclical fluctuations in the demand for private sector credit, where the latter variable is determined primarily by other factors.

To test this hypothesis, we ran bivariate Granger-causality tests between the stationary components of private sector credit and industrial output. For some countries where we reported positive correlations between these two variables, we did find that private sector credit had predictive power for industrial output in the Granger-causal sense. However, for a couple of these countries, there was also evidence of reverse causation from output to credit. Thus, we do not find robust evidence of a unidirectional causal relationship from credit to economic activity. Nevertheless, the strong positive association between private sector credit and the domestic business cycle in some of the countries in our sample has important implications for the design of stabilization programs. Ignoring this link may exacerbate the output cost of a restrictive monetary policy aimed at bringing down inflation.

4.5 Foreign trade and the business cycle

In this subsection, we explore the relationship between domestic business cycle fluctuations and fluctuations in price and quantity variables that are relevant for international trade. In particular, we examine correlations of fluctuations in industrial production with fluctuations in merchandise trade and measures of both nominal and real effective exchange rates.

An adequate measure of foreign trade transactions would be the trade balance, constructed as the difference between real exports and real imports, and divided by real GDP in order to control for scale effects. In the absence of reliable data on price deflators for exports and imports, many authors use the ratio of nominal exports and imports to output. Unfortunately, we are even more constrained because we have only real industrial output data available for most of the countries in our sample. Hence, we use the ratio of exports to imports at current prices as a rough measure of the trade balance. Because terms-of-trade changes could be large and important for these countries, we return to that issue later.

Table 11 presents the correlations between our proxy for trade balance movements and industrial output. For Chile, Mexico, Turkey, and Uruguay, the contemporaneous correlations are negative irrespective of the filter used, similar to the findings for industrial countries reported by a number of authors—see Fiorito and Kollintzas (1994), Prasad and Kumar (1996), and references therein. However, for a few countries—including Morocco, Nigeria, and the Philippines—the contemporaneous correlations are strongly positive. The latter result possibly reflects a strong link between changes in industrial output and exports of manufactures, or the fact that merchandise imports are not highly sensitive to domestic demand fluctuations. In addition, where we find significant correlations between the trade ratio and domestic output, these correlations peak at (or near) lag zero. We interpret this as indicative of the close relationship between international trade and industrial sector output in these middle-income economies, with the latter variable being a good proxy for output in the traded goods sector (other than primary commodities).

We were able to obtain unit values of imports and exports and construct a quarterly index of the terms of trade for only three of the countries in our sample—Colombia, Korea, and Mexico. For these three countries, as shown in Table 12, there is a strong positive correlation between the cyclical components of industrial production and the terms-of-trade index. For Colombia and Mexico, the band-pass filtered data yield the strongest correlations, suggesting that the positive relationship between output and the terms-of-trade might be obscured when using other filters because of the large amount of high frequency variation in the terms-of-trade data.

Because these middle-income countries are unlikely to affect the world price of any industrial commodities, the positive correlations may be interpreted as reflecting demand shifts that lead to simultaneous increases in the world price and in the export demand for the industrial sector output of these countries. The strong positive correlations between lagged values of the terms-of-trade index and contemporaneous output provide further evidence in favor of this interpretation. Overall, our results are consistent with those of Rodríguez-Mata (1997) for Costa Rica, those of Kose and Riezman (1998) for sub-Saharan Africa, and those of Mendoza (1995), whose results suggest that almost half of the fluctuations in output in developing countries can be explained by terms-of-trade disturbances, and Deaton and Miller (1995), who find evidence that in sub-Saharan Africa export price shocks have had substantial contemporaneous effects on output.[24]

[24]Mendoza's analysis of the importance of terms-of-trade shocks in driving business cycles in small open developing economies uses stochastic simulation techniques. It should

4.6 Cyclical behavior of exchanges rates

Next, we examine the correlations between industrial output and measures of nominal and real effective exchange rates.[25] The interpretation of these unconditional correlations is complicated by the fact that the short-run relationship between these variables is crucially dependent on the sources of macroeconomic fluctuations. Nonetheless, it is useful to look at unconditional correlations for two reasons. First, the sign and magnitude of unconditional correlations could provide some indication of the types of shocks that have dominated fluctuations over a particular period of time. Second, these correlations could help in interpreting the correlations between output and trade variables examined above.

Table 13 reports correlations between nominal effective exchange rates and industrial output. In a majority of the countries in our sample, the contemporaneous correlations are not significantly different from zero. In Morocco, Tunisia, and Turkey, there is some evidence of a positive relationship between these two variables, while the correlations are generally negative for the Philippines and Uruguay. The correlations between output and real effective exchange rates, presented in Table 14, reveal a similar picture, but with a few important differences. The contemporaneous correlations for Mexico and Uruguay now turn positive while the correlations for Morocco and the Philippines are now close to zero. The lack of a systematic, strong relationship between real exchange rates and the business cycle is consistent with the notion that this relationship is affected by an amalgam of supply, real demand, and nominal shocks—all of which could have very different effects on this correlation.

One interesting aspect of these results is that, for many countries, the correlations are quite similar using either nominal or real measures of effective

be noted that Hoffmaister and Roldós (1997), using structural VARs estimated over the period 1970-93, find that, for a group of Asian and Latin American countries, terms-of-trade shocks play a small role in output fluctuations.

[25] There are, of course, important differences among these countries in terms of exchange rate arrangements. However, since we use trade-weighted measures of both real and nominal effective exchange rates, the fact that certain bilateral exchange rates could be fixed does not, in principle, affect the interpretation of our results. It should also be noted that the effective exchange rates are defined here such that an increase in the exchange rate implies an appreciation of the currency (in real or nominal terms, as the case may be). Thus, a positive correlation indicates that the exchange rate tends to appreciate when the cyclical component of output rises.

exchange rates. This is in line with a substantial body of research that has documented that, for industrial countries, nominal and real exchange rates are strongly positively correlated at business cycle frequencies—see, for instance, Mussa (1986) and Taylor (1995). Indeed, the contemporaneous correlations between real and nominal effective exchange rates were very strongly positive for all countries in our sample, irrespective of the filter used.

5 Summary of the Findings

The main findings of the paper can be summarized as follows:

- Output volatility—as measured by standard deviations of the filtered cyclical components of industrial production—varies substantially across developing countries, but is on average much higher than the level typically observed in industrial countries. There is also considerable persistence in output fluctuations in developing countries.

- Activity in industrial countries has a positive but relatively weak influence on output in developing countries. Real interest rates in industrial countries tend to be positively associated with output fluctuations in our sample of countries.

- Government expenditure is countercyclical. Government revenues are acyclical in some countries, and significantly countercyclical in others—a phenomenon that appears difficult to explain. The fiscal impulse (defined as the ratio of government spending to government revenue) is negatively correlated with the business cycle.

- The cyclical behavior of nominal wages varies markedly across countries and is not robust across filters. By contrast, the evidence strongly supports the assumption of procyclical real wages.

- There is no consistent relationship between the stationary components of the levels of output and prices, or the levels of output and inflation. Variations in the price level and inflation are countercyclical in a number of countries and procyclical in a few.

- Contemporaneous correlations between money (measured through various monetary aggregates) and output are broadly positive, but not very strong—in contrast to the evidence for many industrial countries.

- The contemporaneous correlations between the velocity of broad money and industrial output are strongly negative across all filters for almost all the countries in our sample. This result is in contrast to the weakly procyclical behavior of velocity observed in most advanced industrial countries.

- Domestic credit and industrial output are positively associated for some countries. However, the strength of the relationship between credit and output is not always robust to the choice of detrending procedure. In some countries, there is a negative correlation between these two variables.

- There is no robust correlation between merchandise trade movements (as measured by the ratio of exports to imports) and output. For some countries, the contemporaneous correlations are negative (irrespective of the filter used), whereas for others the contemporaneous correlations are strongly positive—the latter result possibly indicating that industrial output fluctuations are driven by export demand and that merchandise imports are not as sensitive to domestic demand fluctuations as in industrial countries.

- Cyclical movements in the terms of trade are strongly and positively correlated with output fluctuations.

- There are no systematic patterns in the contemporaneous correlations between nominal effective exchange rates and industrial output; in addition, for a majority of the countries under study, these correlations are not significantly different from zero. Similar results are obtained for real effective exchange rates.

Overall, our results suggest the importance of supply-side shocks in driving business cycles in developing countries. Of course, using cross-correlation coefficients as indicators for evaluating the empirical relevance of demand-oriented, versus supply-oriented, macroeconomic theories can be problematic.[26] The results are also not uniform across countries. In particular, whereas negative price-output correlations in some countries provide support for "real" or supply-side interpretations of business cycles, countries where price-output correlations are positive would tend to support "nominal" or demand-side interpretations.

Our results are in line with those for a number of industrial countries and those obtained by Hoffmaister and Roldós (1997) for Asia and Latin America, and Hoffmaister, Roldós and Wickham (1997) for Sub-Saharan Africa, which suggest that in developing countries the main source of output fluctuations are supply shocks—even in the short run.[27]

6 Concluding Remarks

The purpose of this paper has been to study the cyclical properties of a large number of (seasonally-adjusted) macroeconomic time series for a group of 12 (mostly middle-income) developing countries, using four univariate detrending methods. We have provided a thorough discussion of the cross-correlation patterns between output and various macroeconomic time series, and attempted to organize our results on business cycle fluctuations in a systematic manner, in order to identify a set of relatively robust regularities that can be used as a "benchmark" to guide theoretical research in development macroeconomics. We have also highlighted similarities and differences between our results and existing studies on business cycle fluctuations in industrial and developing countries.

Because our main findings were summarized in the previous section, we conclude with several remarks on the methodological and analytical implica-

[26] For instance, Judd and Trehan (1995), using two econometric models, have shown that correlation coefficients for prices and output could easily be negative even if demand shocks were the primary source of cyclical fluctuations and prices were procyclical. Gavin and Kydland (1995) have also shown that alternative money supply rules can change the cyclical nature of prices in a flexible-price economy. More generally, covariation among a set of variables may depend not only upon the nature of the shocks that perturb the economy, but also (under rational expectations) upon how long the lag is between perception (announcement) and realization (implementation) of the shock.

[27] Hoffmaister and Roldós (1997) also emphasize that in Latin America external shocks—in particular, world interest rate shocks—and demand shocks affect output fluctuations more than in Asia. Hoffmaister, Roldós and Wickham (1997) suggest that external shocks—including changes in the terms of trade—tend to have a greater impact on output in CFA franc countries, compared to the rest of Sub-Saharan Africa.

tions of our analysis. First, our results suggest that, although the correlations derived from alternative filters were often very similar, there are several quantitative (as well as qualitative) features of the data that are not robust across detrending methods. This is similar to the results obtained by Blackburn and Ravn (1991), Canova (1998), and Park (1996). Because it is not generally possible to know *ex ante* when results will vary across filters, considering systematically an array of alternative detrending methods remains an important test of robustness in empirical research on business cycle regularities.

Second, as noted earlier, the unconditional correlations between various variables (such as exchange rates or prices) and domestic output could be small because they average across the effects of different types of shocks. It is therefore important to develop and estimate structural models, along the lines, for instance, of Ahmed and Park (1994), Hoffmaister and Roldós (1996, 1997), Hoffmaister and Végh (1996), and Rogers and Wang (1995), that attempt to separate out the effects of different types of macroeconomic shocks on variables such as prices, output and exchange rates in developing countries. However, existing methods remain controversial and we do not yet have models that convincingly separate out different types of shocks.

Third, the analysis in this paper has ignored the possible effects of measurement errors on the raw data. This is a potentially serious problem. For instance, in our analysis of the correlations between domestic output and foreign interest rate shocks, we did not account for the risk premium that borrowers from developing countries typically face on world capital markets. However, there is considerable evidence that such premia can be large (particularly for countries with a high external debt-to-output ratio) on average, and could change in unpredictable fashion in the short run—as a result of sudden shifts in market sentiment. This measurement problem—which has not been adequately addressed in other existing studies—suggests that caution should be exercised in judging the strength and direction of correlations between domestic output and a measure of world interest rates that does not capture movements in country-specific risk premia.

Finally, at the analytical level, a natural step forward is to build stochastic general equilibrium simulation models of small open developing economies in order to assess if such models (properly calibrated) can reproduce the stylized facts highlighted in this paper. As noted earlier, some of the correlations established in the present study (such as the countercyclical behavior of government spending) can indeed be explained within the framework of some existing theoretical constructs. Building more general quantitative models that are capable of accounting for the type of business cycle regularities highlighted here could prove important for the design of stabilization policies and for macroeconomic management in developing countries.

Appendix I
Detrending Techniques

This appendix provides a brief technical description of the three univariate detrending filters described in section 2.

The modified HP filter

By manipulating the first-order condition of the minimization problem shown in equation (2) (see King and Rebelo, 1993), a time domain representation of the standard HP filter can be developed in which the trend component x_t^* is represented by a two-sided symmetric moving average expression of the observed series:[28]

$$x_t^* = \sum_{h=-\infty}^{\infty} \alpha_{|h|} x_{t+h}, \tag{A1}$$

where the parameters $\alpha_{|h|}$ depend on the value of the Lagrange multiplier λ. Let $g_t^*(\lambda)$ be the smoothed series calculated from the "leave-out" procedure using the value of λ for the smoothing parameter. The cross-validation function is then defined as

$$CV(\lambda) = T^{-1} \sum_{t=1}^{T} \left(y_t - g_t^*(\lambda)\right)^2 \tag{A2}$$

which can be re-written as

$$CV(\lambda) = T^{-1} \sum_{t=1}^{T} \frac{\left(y_t - g_t(\lambda)\right)^2}{\left(1 - A_{tt}(\lambda)\right)^2} \tag{A3}$$

where $g_t(\lambda) = \sum_{s=1}^{T} A_{ts}(\lambda) y_s$. The cross-validatory choice of λ is then the value of λ that minimizes $CV(\lambda)$. For the HP filter, the minimization of $CV(\lambda)$ involves a great deal of computation, and for this reason we prefer a modified method called generalized cross-validation. The idea of generalized

[28]For the minimization problem, the first-order condition takes the form of the fourth-order difference equation:

$$\begin{aligned} 0 \quad = \quad & -2(x_t - x_t^*) + 2\lambda[(x_t^* - x_{t-1}^*) - (x_{t-1}^* - x_{t-2}^*)] \\ & -4\lambda[(x_{t+1}^* - x_t^*) - (x_t^* - x_{t-1}^*)] \\ & +2\lambda[(x_{t+2}^* - x_{t+1}^*) - (x_{t+1}^* - x_t^*)], \end{aligned}$$

which can be manipulated to give equation (A1); see King and Rebelo (1993) for details.

cross-validation is to replace the weighting matrix $A_{tt}(\lambda)$ by its average value over all t. This gives the generalized cross-validation function as

$$GCV = \frac{T^{-1}\sum_{t=1}^{T}(y_t - g_t(\lambda))^2}{(1 - T^{-1}\,\mathrm{tr}\,A_{tt}(\lambda))^2}. \tag{A4}$$

To use (A4), one still needs to find the trace of $A_{tt}(\lambda)$. To calculate $\mathrm{tr}\,A_{tt}(\lambda)$, one could use a singular value decomposition or the approximating algorithm of Silverman (1984). We use the latter method.

The band-pass filter

The band-pass filter developed by Baxter and King (1995) is constructed by combining various simple filters. The most basic building block is the low-pass filter which removes high frequency noise. An ideal low-pass filter which passes frequencies between $[-\underline{\omega}, \underline{\omega}]$ is defined as

$$x_t^* = LP_\infty(L, \underline{\omega})x_t,$$

where the polynomial in the lag operator L is defined as $LP_\infty(L, \underline{\omega}) = \sum_{h=-\infty}^{\infty} b_h(\underline{\omega})L^h$ and the weights are given by $b_0(\underline{\omega}) = \underline{\omega}/\pi$, and $b_h(\underline{\omega}) = \sin(h\underline{\omega})/h\pi$ for $h = 1, 2, ...$ Of course, in practice we must use a finite moving average to approximate this ideal filter, that is

$$LP_K(L, \underline{\omega}) = \sum_{h=-K}^{K} a_h(\underline{\omega})L^h,$$

with $a_0(\underline{\omega}) = \underline{\omega}/\pi$ and $a_h(\underline{\omega}) = \sin(h\underline{\omega})/h\pi$ for $h = 1, 2, ...K$.

A high-pass filter is defined similarly. That is, a high-pass filter, denoted $HP_K(L, \overline{\omega})$, has the same form as the low-pass filter but with weights equal to $1 - b_h(\overline{\omega})$. The band-pass filter (denoted $BP_k(\underline{\omega}, \overline{\omega})$) is then defined by subtracting the weights of the low-pass filter from the high-pass filter.

In addition, a constraint is added to the low-pass filter to ensure that the weight at the zero frequency is unity. This constraint is imposed by adding to each weight the term $(1 - \sum_{h=-\infty}^{\infty} b_h(\underline{\omega}))/(2K+1)$. A similar constraint is added to the high-pass filter to ensure that the weight at the zero frequency is zero. The resulting band-pass filter will then have weights that sum to zero, thus eliminating stochastic trends and deterministic polynomial time trends of order 2 or less. Baxter and King (1995) argue that for quarterly data one should chose $K = 12$ and frequency cut-offs that correspond to 6 quarters and 32 quarters. We adopt the same cut-offs in this paper.

The nonparametric filter

The nonparametric technique used in this paper can be described as follows. Consider the decomposition (1), and let the trend component x_t^* be an unknown smooth regression function. The Nadaraya-Watson estimator $\widehat{x}_t^h$ of x_t^* has the form

$$\widehat{x}_t^h(z) = \frac{T^{-1}\sum_{t=1}^{T} K_h(z - t/T)\, x_t}{T^{-1}\sum_{t=1}^{T} K_h(z - t/T)},$$

where $K_h(u) = h^{-1}K(u/h)$, h is the bandwidth parameter, T is the sample size, and $K(u)$ is a kernel with support $[-1, 1]$ that satisfies $\int K(u)\ du = 1$. The kernel we will use is the Epanechnikov kernel and is defined as

$$K(u) = 0.75\left(1 - u^2\right)\mathrm{I}\left(|u| \leq 1\right), \tag{A5}$$

where $\mathrm{I}(|u| \leq 1)$ is the indicator function. This parabolic shaped kernel is a commonly used kernel function which enjoys some optimality properties. The accuracy of kernel smoothers as estimators of x_t^* is a function of the kernel K and the bandwidth h, the most important factor being the choice of h. As in the case of the HP filter, we use a data dependent bandwidth selection procedure to determine the value of h that optimizes quadratic error measures for the regression curve. Specifically, we use the cross-validation method as described above. Another application of this filtering technique can be found in Coe and McDermott (1997).

Appendix II
Data Sources and Unit Root Tests

The primary sources of the data used in this study are the IMF's *International Financial Statistics* and Information Notice System, supplemented by various other sources. We provide below a description of the series, together with their IFS codes. All the data are available upon request.

- *Real output* is the industrial production index (series 66) for Mexico, Korea, India, Malaysia, and Tunisia, and the manufacturing production index (series 66ey) for Chile, Morocco, Nigeria, the Philippines, and Uruguay. The industrial production index was obtained from the IMF desk economist for Colombia, and from the OECD database for Turkey. For Turkey, Tunisia, and Uruguay, partial information was also provided by IMF desk economists.

- The *consumer price index* (CPI) is series 64 for all countries. For Tunisia, missing data in IFS were filled in for us by the IMF desk economist.

- The *nominal wage index* is series 65 for Mexico, Chile, the Philippines, and series 65ey for Korea. Data for Colombia were obtained from the IMF desk economist. Data for Turkey were obtained from the OECD database and the IMF desk economist. The *real wage index* is obtained by deflating the nominal wage series by the CPI.

- The *monetary base* (or reserve money) is series 14 for all countries. *Narrow money* is series 34 and *broad money* the sum of series 34 and 35, again for all countries. *Velocity* for each monetary indicator is calculated first by transforming the monetary aggregate into an index, and then by dividing by the product of the CPI and the real output index—which is used as a proxy for nominal output.

- *Private sector credit* is series 32d for all countries. The real credit variable is obtained by deflating the nominal aggregate by the CPI. For Tunisia, missing data in IFS were filled in for us by the IMF desk economist.

- *Government expenditure* in nominal terms is series 82 for Mexico, Korea, and the Philippines. Data for Chile were obtained from Chile's Ministry of Finance. The *expenditure index* is derived first by transforming the nominal series into an index, and then by dividing by the same proxy for nominal output used to derive velocity indicators.

- *Government revenue* in nominal terms is series 81. The *revenue index* is derived in the same way as the expenditure index.

- The *fiscal impulse measure* is derived by dividing series 82 by series 81.

- The *trade ratio* is measured as the ratio of merchandise exports at current prices (series 70) to merchandise imports at current prices (series 71), with both variables measured in U.S. dollar terms.

- Trade-weighted measures of *nominal* and *real effective exchange rates* are obtained from the IMF's Information Notice System.

- The *terms of trade* data are measured by the ratio of export unit values (series 74) to import unit values (series 75) for Colombia and Korea. For Mexico, export and import price indices were obtained from the OECD database.

- *World output* is proxied by the industrial production index for industrial countries (series 66, code 110). The *world real interest rate* is proxied by the difference between the nominal Euro-dollar rate in London (series 60d, country code 112), and the rate of inflation in consumer prices in industrial countries (series 64, code 110).

We performed a set of standard unit root tests, including the ADF test and the Phillips-Perron tests, on our raw data series (all of which were converted into logarithms for the empirical work, except for the world real interest rate). These tests indicated that virtually of the series were nonstationary in levels over the relevant sample period and, therefore, that computing correlations using the raw data would not be appropriate. We also used similar unit root tests to confirm that the cyclical components obtained with the filters employed in this paper were indeed stationary. In addition, we found that the inflation rate (measured as the four-quarter change in the price level) did not appear to be stationary in levels for several countries in our sample. To conserve space, the results of these unit root tests are not reported here but are available from the authors.

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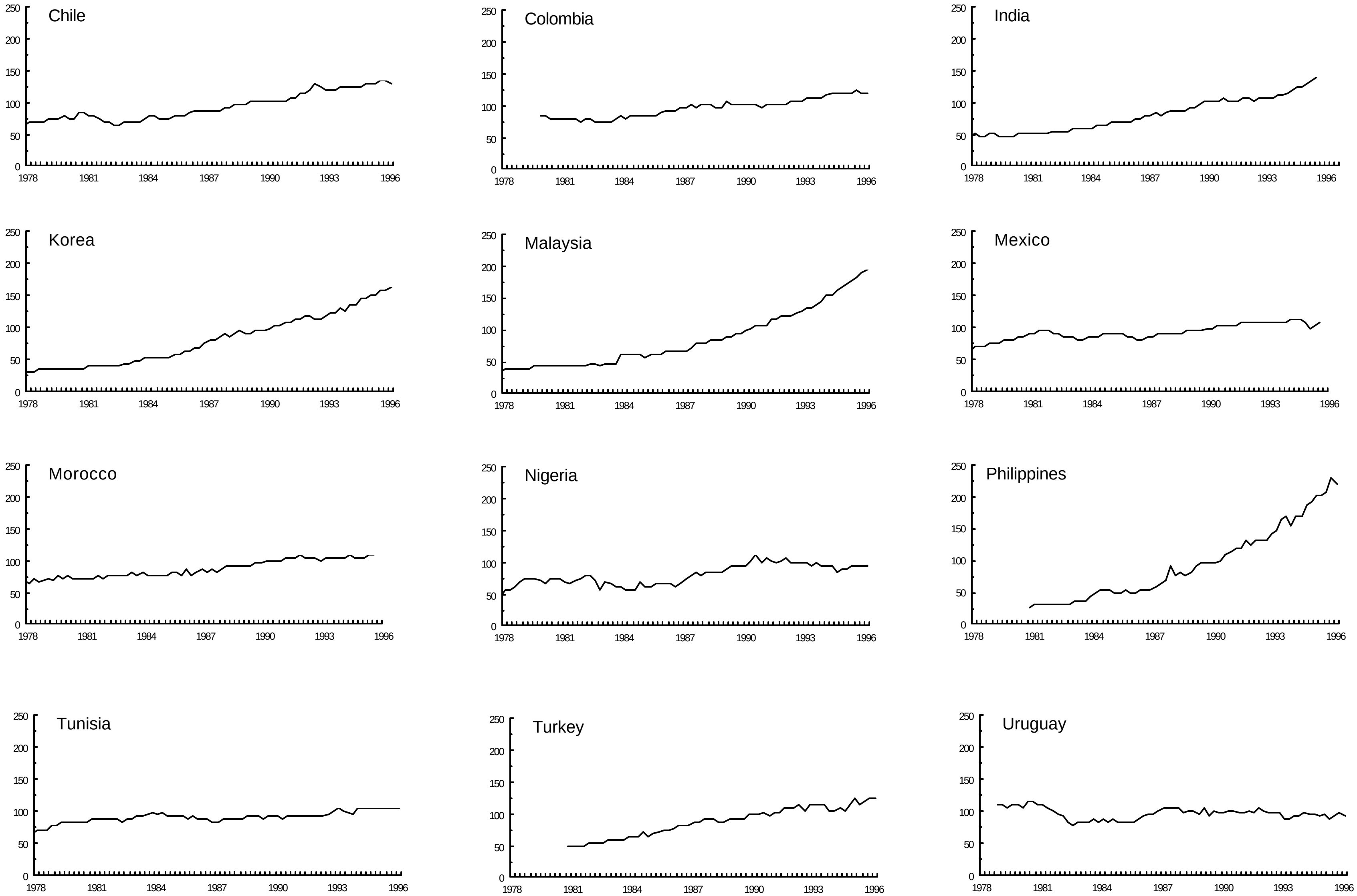
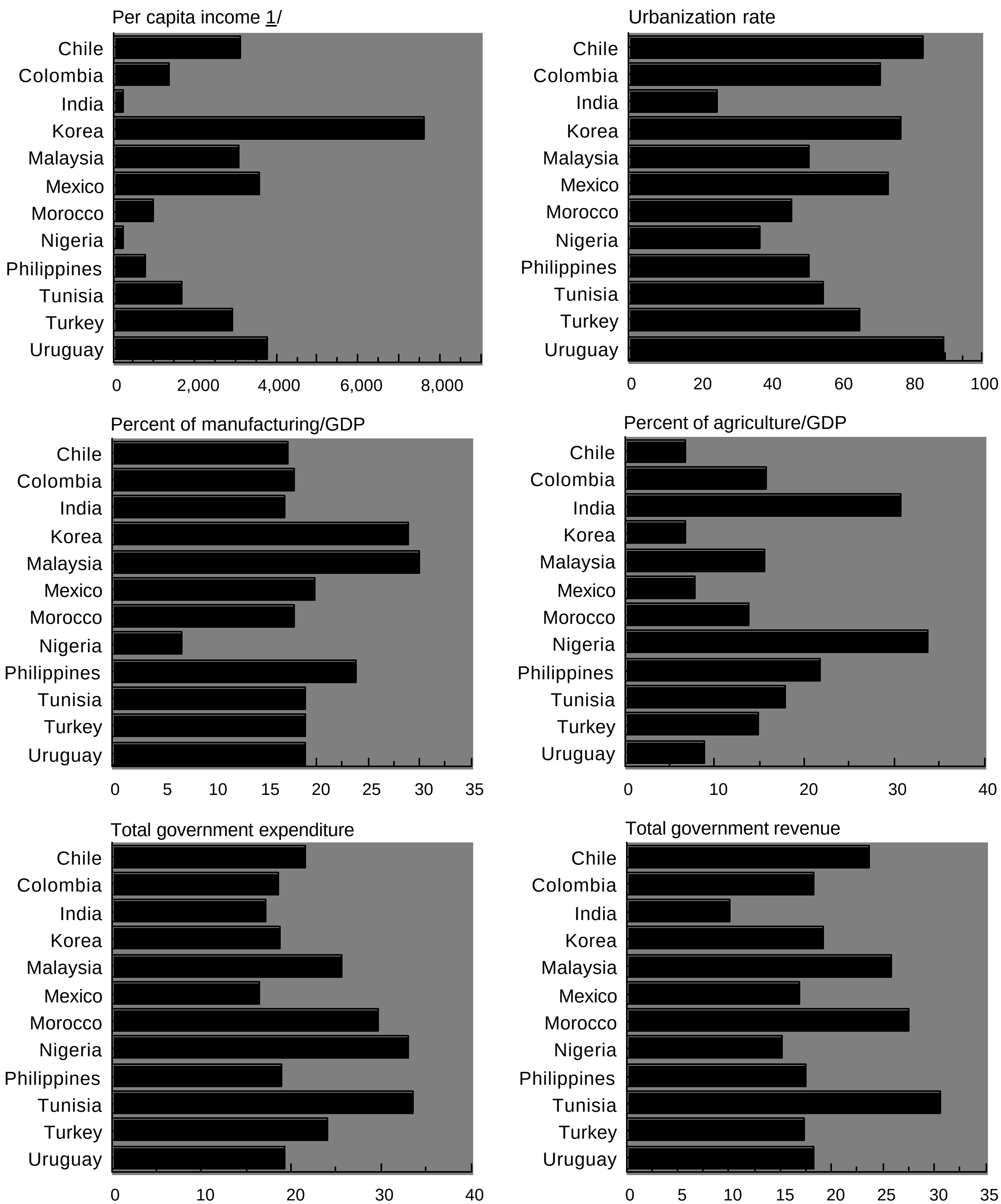
Figure 1
Industrial Production of Selected Developing Countries (1990 = 100)



Source: IMF and World Bank.

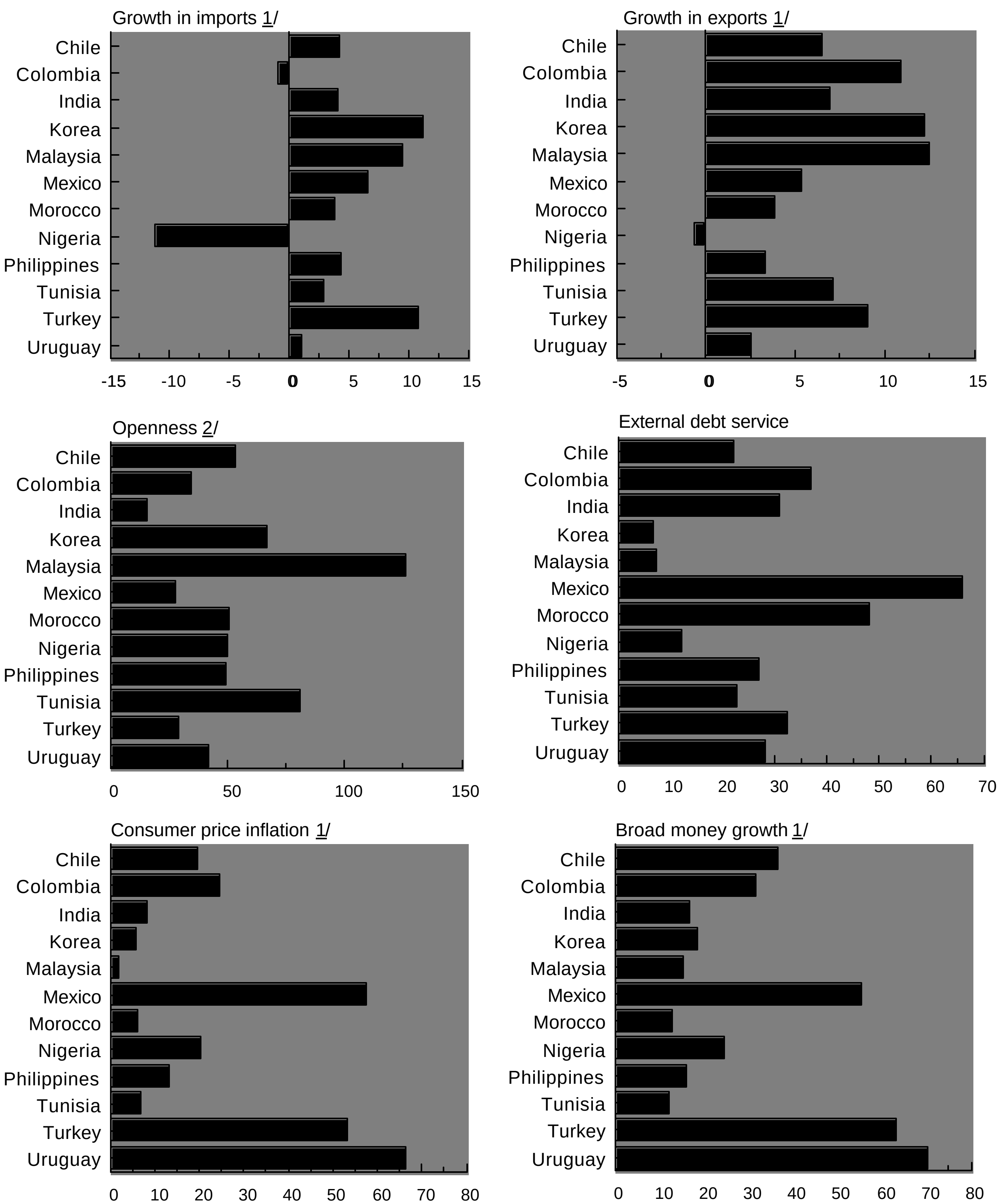
Figure 2
Selected Developing Countries: Economic Indicators
(Data refer to 1993, unless otherwise indicated)



Sources: IMF and World Bank.

1/ In U.S. dollars.

Figure 2 (concluded)
Selected Developing Countries: Economic Indicators
(Data refer to 1993, unless otherwise indicated)



Sources: IMF and World Bank.

1/ Average annual growth rate, in percent, 1980-93.
2/ Average annual ratio, in percent, 1980-93.