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09. December 2010

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December 4, 2010

Abstract

Applying nonstationary panel data econometric methods, this paper analyzes the major sources and transmission of inflation in the Gulf Cooperation Council (GCC) countries over the 1980-2008 period. We argue that, in GCC countries, money is essentially demand determined, so that the high collinearity between money and aggregate demand indicators such as non-hydrocarbon output is expected and should be dealt with accordingly. Several important results emerge from the analysis. First, the money supply stands out as a significant determinant of inflation both in short- and long-run. Both foreign prices and the nominal effective exchange rate are shown to be more successful in explaining inflation in the long-run than the short-run. The half-life of the speed of adjustment reveals that it takes about 2.9 years for 50% of a shock to the long-run equilibrium to dissipate. An implication of our results is the case it makes for more sovereign monetary policies in GCC countries.

JEL Codes: C32, C33, E31, E50, H30

Keywords: Inflation, Monetary policy, Fiscal policy, Exchange rates, Oil price, Panel data.

*This paper has benefitted from discussion with or comments from Khalid Alkhater, Katrin Assenmacher-Wesche, Alfred Haug, Rosmy Jean Louis, M.S. Rafiq and Joakim Westerlund. We particularly thank M.S. Rafiq for help with proofreading. The views expressed here are those of the authors and do not reflect the official view of the Qatar Central Bank. The usual disclaimer applies.

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1 Introduction

Understanding the sources of inflation is a prerequisite for devising effective economic policies to fight inflationary pressures. For example, if changes in the money supply contribute significantly to the changes in domestic price level, then central banks can devise appropriate monetary policies to keep inflation under control. In the standard case of a small open economy, if expansionary fiscal policy is to blame for the persistent inflationary pressure, in general, fiscal policy can be contracted to prevent inflation from rising. However, in the special case of the Gulf Cooperation Council (GCC) countries, tightening fiscal policy can be counterproductive given that fiscal policy is the main driver of growth of the non-hydrocarbon sector.[1] Rather, monetary policy aiming to sterilize the fiscal policy impact on inflation is more appropriate for GCC countries. Movements in global commodity prices can also significantly affect prices in domestic markets. This channel is particularly important for GCC countries, who rely heavily on imported goods (both hard and soft). Similarly, changes in global economic sentiment are also likely to play a major role in the determination of domestic price levels. There now exists significant evidence on the global determinants of domestic inflation,[2] and how increased globalization has affected local inflation (see IMF, 2006).

The main objective of this paper is to empirically analyze the major determinants of inflation in the six GCC countries namely, Bahrain, Kuwait, Oman, Qatar, Saudi Arabia and United Arab Emirates (UAE). The GCC countries exhibit many unique features that make it necessary to adapt (rather than adopt) an existing economic model to properly analyze the inflation dynamics. Care must be given to incorporate country (or region) specific characteristics when examining the determinants of inflation. GCC countries share common features such as high dependence on hydrocarbon (see Table 1), high exposure to global markets in terms of both trade and liberalized capital accounts, (almost) identical social and political structure, flexible labor markets, common exchange rate regimes and, above all, the six countries share a common market policy and are in the process of launching a monetary union.

Given the similarities, one would tend to think that domestic price levels in the GCC countries are likely to be affected by similar factors. Indeed, the commonality of economic and

[1]Furthermore, fiscal consolidation can significantly slow down the process of economic diversification, which refers to the development of the non-hydrocarbon sector and/or the reduction of the proportion of government revenue and export proceeds from the hydrocarbon (oil and gas) sector. See, Basher (2010) and the references therein for discussion on economic diversification in the GCC region.

[2]See, for example, Borio and Filardo (2007) and Ciccarelli and Mojon (2008).

Table 1: The role of hydrocarbon in GCC countries' government revenues, exports and GDP: 1980 and 2007 (percent)

	Government revenues		Exports		GDP	
	1980	2007	1980	2007	1980	2007
Bahrain	77.0	80.0	33.6	79.2	28.0	24.6
Kuwait	82.0	93.1	90.0	94.4	59.7	53.2
Oman	86.0	76.4	92.4	75.8	59.3	45.1
Qatar	94.0	60.0	95.0	80.8	64.0	56.4
Saudi Arabia	91.2	82.5	99.9	88.0	65.8	50.9
UAE	96.0	77.0	94.0	38.5	57.0	11.1

Source: Basher (2010).

non-economic features among the GCC countries leads naturally to a panel data approach, which treats the countries in the panel as a single economic block. One advantage of studying economic issues in panel is that by combining both time series dimension (T) and the cross-section dimension (N) of the data, the power of the statistical tests and estimators can be considerably improved over conventional tests. This is particularly desirable in a situation where it is problematic to conduct country-specific analysis due to a dearth in macroeconomic data.

Nevertheless, the use of panel data also poses several challenges that should be accounted for. First, while many panel estimators are based on pooling the data across cross sections, attention should also be given to unit-specific (here, country) heterogeneity. Within the GCC region, member countries differ considerably from each other in terms of their dependence on hydrocarbon revenues (Table 1), the level of financial development (IMF, 2008a) and the size of their respective market. These differences are likely to influence the inflation dynamics at least in the short-run and, therefore, must be accounted for in the econometric modeling. Second, despite their differences, the GCC countries are likely to be contemporaneously correlated (i.e. cross-section dependence) due to, for example, their exposure to common observed global shocks such as changes in oil prices. Accounting for such cross-section dependence is crucial in obtaining consistent estimates of the standard errors of the regression coefficients. In this paper, we consider the above two issues in selecting the suitable panel estimators and tests.

Rigorous analysis of the sources of inflation in GCC countries is important from several policy perspectives. First, the GCC countries are in the process of creating a monetary union.[3] Therefore, implementation of a common monetary policy requires the knowledge of common fac-

[3] As of 2009, Bahrain, Kuwait, Qatar and Saudi Arabia have agreed to join the monetary union, while Oman and UAE have temporarily opted out from the planned monetary union.

tors affecting price levels in all member countries. Second, fiscal policy remains (and is likely to continue to be) the major economic driver of growth – particularly for the non-hydrocarbon sector. While stimulative fiscal policies may be needed to create a well diversified non-hydrocarbon sector, too loose a fiscal policy can generate serious inflationary pressure thereby, ceteris paribus, undermining the goal of economic diversification. Hence, quantifying the impact of fiscal stance on inflation is crucial not only for its own sake, but also for helping to design an appropriate operational monetary policy framework for the common monetary policy. Third, knowledge of the determinants of inflation is often a key concern for foreign investors contemplating long-term investment in the GCC region. Finally, understanding the nature of inflation dynamics is a prerequisite for building a sensible inflation forecasting model for policy purposes.

Our main findings can be summarized as follows. The money supply is a significant determinant of inflation in both the short- and long-run. Both foreign prices and the nominal effective exchange rate are shown to explain inflation in the long-run, but not necessarily in the short-run. The half-life of the speed of adjustment reveals that it takes about 2.9 years for 50% of a shock to the long-run equilibrium to dissipate. The results in this paper make the case for more independent monetary policies in GCC countries.

The rest of the paper is organized as follows. Section 2 provides a historical overview of the inflation dynamics in GCC countries. Section 3 reviews related literature on the sources of inflation in GCC countries. Section 4 discusses the various channels of inflation transmission in GCC countries. Section 5 presents the empirical investigation and results. Section 6 concludes the paper. Some details on the econometric techniques are provided in the Appendix.

2 A Historical Background of Inflation in GCC States

Figure 1 plots annual inflation rates (percent) across the six GCC countries over the 1980-2008 period. Roughly, the sample can be divided into three episodic periods of inflation: *disinflation* (1980-1986), *moderate inflation* (1987-1999)[4] and *high inflation* (2000-2008). For example, the average annual aggregate GCC inflation rate declined from 7% in 1980 to 1.70% by 1986, indicating evidence of disinflation during the early part of the sample period. Likewise, aggregate GCC inflation ranged between 1.65% and 0.86% from 1987 to 1999, supporting the notion of moderate or stable inflation during the middle part of the sample period. In contrast, the

[4]For some countries, the incidence of two inflationary cycles within the 1987-1999 period is evident.

aggregate GCC inflation climbed from -1.47% in 2000 to 10.63% in 2008. Nevertheless, from Figure 1 it is also evident that inflation rates varied in terms of their magnitude and persistence across the six countries, and were affected by country-specific events (e.g. the Iraqi invasion of Kuwait in 1990). What could explain the episodic nature of inflation in GCC countries?

Figure 2 presents historical movements in the oil price, non-fuel commodity price index, federal funds rate and the trade-weighted U.S. dollar index over the 1980-2008 period. There is a close association in the timing of the episodic phases of these exogenous variables with that of GCC's inflation episodes. For example, the historical path of oil price movements can roughly be classified into three categories: *falling oil prices* (1980-1986), *stable oil prices* (1987-1999) and *rising oil prices* (2000-2008). Spot oil prices declined by about 60% between 1980 and 1986, hovered within a fairly narrow band of 18 to 20 dollars per barrel over the 1987-1999 period and skyrocketed by over 200% from 2000 to 2008.[5] It is not coincidental to observe the strong correlation in the timing of the persistence between oil prices and GCC inflation rates' given that hydrocarbon revenues determine the fiscal stance, hence play a major role in the economic development of the GCC countries (Table 1).

Similar to oil prices, the international non-fuel commodity price index exhibits episodes of declining, moderating (with some fluctuations in the intervening years) and rising price levels. For example, the non-fuel commodity price index had generally declined from 127 in 1980 to 98 in 1986 (with a base year in 2000), fluctuated between 95 and 128 over the 1987-1999 period and rose by about 90 points between 2000 and 2008. Given that the GCC countries are net importers of most types of non-fuel commodities, it is hardly surprising to see similarities in GCC's inflation and changes in international commodity prices.

In terms of the Federal Reserve's monetary policy stance, which is closely mimicked by GCC central banks, the 1980-1986 period can be characterized by a relatively tight monetary policy, followed by a period (1987-1999) of more neutral monetary stance and finally the more recent episode (2000-2008) of easy monetary policy by historical standards. Despite the declining trend in the federal funds rate during 1980-1986, its annual average during this episode was almost 11 percent. This tighter monetary stance by the Federal Reserve (and therefore by GCC central banks) is consistent with the falling price levels in GCC countries during 1980-1986. In contrast,

[5]Following the Iraqi invasion of Kuwait in 1990 (the Gulf War), spot oil prices doubled from below $15 a barrel in July 1990 to about $30 in August 1990 (monthly average). However, oil prices soon began to fall following the allied military victory. By March 1991, average spot prices were again around $15 a barrel. See Hutchison (1991) for further discussion.

the annual average federal funds rate during 1987-1999 was 5.76 percent, falling further to 3.26 percent over the 2000-2008 period. Therefore, despite the fluctuations, the episodes of moderate and easy monetary stance have the potential to explain much of the episodic nature of inflation in the GCC countries. Finally, the strong appreciation of the U.S. dollar over the 1980-1985 period is clearly consistent with the falling price levels in GCC countries, as a stronger dollar helps to lower the cost of imports in GCC countries. Following the sharp depreciation for a brief period in mid-1980s, the U.S. dollar maintained a relatively stable position over the 1988-2001 period, before falling precipitously from 2002 onwards.[6]

Allowing for caveats, the four foreign exogenous variables provide a reasonable approximation of the inflation dynamics in GCC countries. Besides these global factors, domestic factors such as the limited absorption capacity play a crucial role in the determination of inflation in GCC countries. Controlling for inflation is therefore a complicated task in the context of GCC due to (i) the lack of monetary independence as a result of the pegged exchange rate system coupled with liberalized capital accounts and (ii) the pro-cyclical nature of fiscal policies – further discussion is provided in Section 4. On a comparative scale, however, over the longer term GCC countries have had a significantly better record controlling inflation than other emerging oil-exporting countries such as Nigeria and Russia (see the discussion in Sturm et al. 2009).

3 Earlier Literature

The earlier literature has evolved along two directions: country specific studies and GCC area wide analysis. Given the economic similarities, it is useful to review the country specific studies to get a broader picture of the inflation dynamics in GCC countries. Based on quarterly data from 1964 to 1975, Keran and Al-Malik (1979) found that developments in the world inflation rate and domestic currency significantly explain the inflation rate in Saudi Arabia.[7] Darrat (1985) employed a polynomial distributed-lag model to empirically investigate the monetary sources of inflation in Saudi Arabia over the 1962Q1-1981Q4 period. Unlike Keran and Al-Malik (1979), the author considered M1 monetary aggregate and included quarterly data on

[6]The trade-weighted U.S. dollar index is only an approximation to the aggregate GCC area wide exchange rate movements. Nonetheless, given the *de jure/de facto* exchange rate peg between the U.S. dollar and GCC currencies, the former is expected to lead the latter.

[7]Currency, rather than M1 or M2 money, is used as the proper definition of money in Saudi Arabia—as back in 1970s, salaries, household and small business transactions were all operated through currency. In addition, interest payments on deposits were legally prohibited; as a result, citizens primarily held demand deposits instead of time and savings deposits.

real income (real GNP), which was obtained through a non-linear interpolation of the annual observations. Results revealed that both money supply growth (with a near unitary elasticity) and inflationary expectations exert a significant positive effect on inflation in Saudi Arabia. Moreover, the associated lag-adjustment of money supply growth is relatively short (three quarters), suggesting that monetary tightening can be an exceptionally effective measure to control inflation in the short run. Movements in foreign interest rates also affected Saudi inflation, whereas real income and bilateral U.S. dollar real exchange rate didn't yield expected results on inflation perhaps due to their questionable inclusion in the model.[8]

Salih (1993), using annual data over the 1970-1987 period, estimated an AR(1) model with an import price index and the money supply as the explanatory variables of inflation in Kuwait.[9] He found that a doubling of world inflation causes an approximately 23 percent increase in domestic inflation, while a 100 percent increase in the growth of broad money supply results in a 9.5 percent increase in domestic inflation.[10] Al-Mutairi (1995) estimated a four-variable VAR model – containing a domestic price index, narrow money supply, government expenditure and an index of import price – for Kuwait over the 1975Q1-1990Q2 period.[11] The impulse response functions show that shocks to the import price have a positive effect on the domestic price level, but the effect strengthens as the forecast horizon lengthens. Shocks to real government expenditure and the money supply induced positive movements in the price level only at intermediate horizons.

Hasan and Alogeel (2008) employed cointegration and error-correction techniques to analyze

[8]For example, the real exchange rate variable appears in the inflation equation through its appearance in the money demand function. Real exchange rate appears in the money demand function to capture the currency substitution effect, which is a quite reasonable assumption in the context of Saudi Arabia. However, when real exchange rate appears in the inflation model, due to the fixed exchange rate peg between Saudi riyal and U.S. dollar, movements in the Saudi-U.S. real exchange rate reflects nothing but the inflation differential between Saudi Arabia and U.S., which may be a good indicator in testing the viability of the riyal-dollar exchange rate peg, but not necessarily as a determinant of domestic inflation. As discussed in Section 4, a relevant exchange rate measure is the nominal effective exchange rate. Moreover, it is not clear whether real GNP is the appropriate measure of real income in Saudi Arabia (and for the rest of the GCC countries). Given that hydrocarbon income comprises a significant portion of GDP, private sector demand for money is more appropriately measured by the level of non-hydrocarbon income, which is the income accrues to the private sector directly through payments of wages and other income sources.

[9]The import price index was obtained by dividing the value of imports by the volume of imports. The unit value import indices based on custom data are deemed to be potentially unreliable and unsuitable due to practical errors and biases (IMF, 2009).

[10]Salih (1993) also estimated a quarterly model of nontradable inflation using similar explanatory variables over the 1970Q1-1990Q2 period. The quarterly model included four lags for the money supply and ten lags for the imported inflation, where the imported inflation is proxied by a Divisia import price index. Results revealed a delayed effect of the money supply growth on nontradable inflation (between third and fourth lagged quarters) but an immediate and persistent effect of imported inflation on domestic inflation. The cumulative effect of imported inflation is reported to be over 26%, slightly higher than the instantaneous effect reported by the annual data.

[11]The simultaneous use of narrow money and government expenditure in Kuwait can be problematic. As conjectured by Salih et al. (1989), the growth rate of a narrowly defined money supply essentially represents the growth rate of government expenditure.

the inflation process in Kuwait and Saudi Arabia. As possible determinants of inflation, they considered nominal effective exchange rate, trading partners price level, price of oil (to proxy transportation costs), money supply and domestic demand. In addition, both excess money supply and excess aggregate demand were included in the short-run model in order to capture the escalation of prices stemming from higher demand pressures.[12] Their results were remarkably similar for both Kuwait and Saudi Arabia. The pass-through of trading partners' inflation on domestic inflation appeared to be more stronger than effective exchange rate pass-through. The estimate of the speed of adjustment was similar with a half-life of only nine months, indicating that domestic (external) factors play a relatively limited (large) role in driving inflation in Kuwait and Saudi Arabia. Both excess demand and excess money supply appear as significant drivers of short-run inflation in Kuwait, while only excess money supply is found to be statistically significant in Saudi Arabia.

Kandil and Morsy (2009) also applied cointegration and error-correction methods to analyze the determinants of inflation in GCC countries using annual data from 1970 to 2007. The main variables considered are the nominal effective exchange rate, prices of major trading partners, broad money and government spending. Lagged excess demand to proxy structural bottlenecks on the supply side was included in the short-run model.[13] Results show that inflation in major trading partners' countries and the depreciation of nominal effective exchange rates have a significant inflationary effect in the long-run for most of the GCC countries, while the money growth channel is inflationary in the long-run in Bahrain and UAE only. Government spending targeted at improving supply-side bottlenecks (such as capacity building or infrastructure) eases long-run inflationary pressure in Kuwait, Oman and UAE; whereas higher government consumption indicated a significant long-run inflationary effect in the UAE. Excess demand exerted inflationary pressure in the short-run in Bahrain, Kuwait and UAE.

Finally, Saidi et al. (2009) analyzed the inflation behavior in the GCC region in framework of the quantity theory of money. They considered quarterly GCC simple-sum aggregated economic data covering the period 1991-2008. The empirical analysis is carried out in two steps: a pairwise

[12] Excess money supply is calculated by subtracting the actual money supply from the estimated long-run money demand. Excess aggregate demand is calculated via de-trending and Hodrick-Prescott filter techniques.

[13] As in Hasan and Alogeel (2008), excess demand variable is defined by the output gap. Hasan and Alogeel (2008) defined real domestic demand as the sum of public and private consumption plus investment, whereas Kandil and Morsy (2009) used real aggregate GDP to compute the output gap. Given the large contribution of hydrocarbon (and therefore exogenous) in the aggregate output, the accuracy of aggregate GDP-based output gap measure is questionable. A desirable alternative is to use non-hydrocarbon real GDP to proxy the output gap.

Granger (1969) causality tests followed by a VAR analysis of selected variables. The causality tests suggested that both the money supply and import prices Granger cause inflation (but not vice versa), while the oil price does not Granger cause inflation. The VAR model also supported the predictions of the quantity theory: inflation is affected mostly by money supply with the peak inflation response to a money supply shock comes about 3 quarters after the shock and strongly persists even after ten quarters. The authors undertook several sensitivity tests, which included additional variables (e.g. exchange rate) in the VAR model, but the role of money supply shocks remains predominant in explaining aggregate GCC inflationary pressure.

4 Channels of Inflation Transmission in GCC States

Given the fixed U.S. dollar peg, narrow domestic production base and heavy reliance on hydrocarbon revenues, there are several obvious external and internal channels through which shocks could propagate inflationary pressure in the GCC countries. This section attempts to summarize the major channels of inflation transmission in GCC countries. The discussion is based on economic theory and related empirical literature.

The foreign interest rate channel: Given the fixed peg to the U.S. dollar, under normal circumstances GCC central banks have to mimic the interest rate policy of the Federal Reserve due to the impossible Trinity (Mundell, 1963).[14] Therefore, any cut in the targeted federal funds rate will likely to put pressure for a similar cut in GCC policy interest rates, irrespective of domestic liquidity conditions in the banking system. Admittedly, inflation is imported through the interest rate parity implied by the exchange rate parity. This effect is accommodated in the analysis by the domestic money supply since a change in the interest rate differential between the anchor and pegging country would result in a change in the level of domestic money supply through, for example, the effect of capital flows on domestic money supply.

[14] According to Impossible Trinity, a country can choose only two of the following three options: free capital movement, autonomous monetary policy or a fixed exchange rate. Until the intensification of global financial crisis, the GCC central banks could not pursue sovereign monetary policy due to the fixed exchange rate systems and open capital accounts. However, following the collapse of Lehman Brothers in September 2008, movement of international capital dramatically fell, thereby creating an opportunity for the GCC central banks to undertake independent monetary policy, at least temporarily. However, not all GCC central banks were able to reap the benefits of autonomous monetary policy. Hence, when the Federal Reserve reduced its targeted federal funds rate twice in October 2008, the other GCC central banks matched the Fed's interest rates cuts, while the Qatar Central Bank did not follow suit with the intention to dampen high domestic inflationary expectations. This shows that—under such circumstances—independent monetary policy can be exercised to meet domestic requirements, even under a hard fixed peg regime.

The direct exchange rate channel: Due to the fixed U.S. dollar exchange regime any change (upward or downward) in the value of U.S. dollar vis-à-vis other currencies automatically produces an identical effect on the level of the Gulf currencies against trading partners' currencies. Thus, for instance, a depreciation of the dollar against the Japanese yen would result in a proportional depreciation of GCC currencies against yen.[15] The same process works in reverse when the dollar appreciates against yen or other currencies. Therefore, the GCC currencies are highly vulnerable to daily U.S. dollar volatility in the sense that their currencies would exhibit similar levels of volatility against third currencies (barring dollar). Often such external shocks are purely exogenous and uncorrelated to the GCC economies (save for those related to oil prices). Since GCC's nominal exchange rates (vis-à-vis dollar) cannot adjust in response to the movements in the dollar, domestic price level changes as a result. One appropriate way to capture this effect is by incorporating a measure of domestic nominal effective exchange rate (NEER) in the model.

The foreign prices channel: In addition to indirectly importing inflation via adopting the easy Federal Reserve's monetary policy stance and/or the depreciation of the U.S. dollar, the GCC countries also import foreign inflation directly. Due to a limited domestic production base (barring hydrocarbon), GCC countries rely heavily on imported consumer and producer goods. As a result, the GCC countries are quite vulnerable to changes in foreign goods' prices and, therefore, the pass-through from foreign to domestic import prices is expected to be strong. The ideal way to capture such pass-through effects is by a domestic import price index or a suitable foreign (export) price index.

The budgetary transmission channel: Given the sole state ownership of hydrocarbon resources, hydrocarbon export revenues have no *direct* impact on the domestic economy until the hydrocarbon income is channeled to the economy through government spending.[16] Given the historical tendency of pro-cyclical fiscal policy, inflation too has become pro-cyclical to government spending. For example, highly expansionary fiscal policies in 1970s and in 2000s led

[15] The effect would be less than proportional for Kuwait as it targets its exchange rate to an undisclosed basket of currencies.

[16] This assertion seems to be quite consistent with empirical results. In a panel regression of 10 oil-producing countries, Husain et al. (2008) found that oil prices do not independently influence the underlying non-oil output, the effect transmits through the impact on fiscal policy.

to strong excess demand, and coupled with emerging supply bottlenecks, resulted in strong inflationary pressures in GCC countries. Whereas, inflationary pressure was relatively benign between mid-1980s and 1990s as a result of, among other factors, tighter fiscal polices in most of the GCC countries.[17] Government spending is the major engine of economic growth for the non-hydrocarbon sector. The extent to which government spending generates pressure on inflation depends much on its composition: higher government consumption is generally associated with higher inflationary pressure in the short-run, whilst higher government investment aiming to ease capacity constraints may lower inflationary pressure over the long-run.[18]

Other domestic factors. Developments in nontradable sectors can significantly influence current and future price levels. In GCC countries, limited absorption capacity often creates a significant upward (cost) inflationary pressure in a number of nontradable sectors, especially in construction and real estate. For example, the shortage of residential and commercial housing units was a key factor behind higher nontradable inflation in several GCC countries (particularly Qatar and UAE) during the recent oil boom of the 2000s. This was the case because during economic upturns, the supply of foreign (skilled and unskilled) workers in GCC countries sharply increases primarily due to the shortages of domestic (skilled and unskilled) workers. The sudden increase in (expatriate) population is—due to construction lags—often not matched with a commensurate increase in residential and commercial housing units, thereby causing an imbalance in the demand and supply for housing units leading to a subsequent increase in rental prices. This process works roughly in reverse during economic downturns.

Overall, the contribution of domestic factors to local inflation can roughly be quantified with a suitable measure of aggregate demand. Several choices come to mind. The first is non-hydrocarbon GDP, the broadest measure of aggregate demand which includes all activities in the nontradable sector. An expansion of the non-hydrocarbon sector contributes to additional capacity buildup and helps reduce prices. Second, until now, government expenditure has been the main driver of growth in the non-hydrocarbon sector. It may be used as a proxy for aggregate demand. Finally, private spending—defined as the sum of household consumption and domestic investment—is also considered an alternative proxy for aggregate demand.

[17] Examining the fiscal policies in 19 oil-exporting countries over the 1965-2005 period, Sturm et al. (2009) find support of procyclical conduct of fiscal policy, including a more pronounced response during 1985-2005 sub-period. For reasons behind the phenomena of procyclical fiscal policy in GCC countries, see Basher (2010).

[18] Empirical evidence of this hypothesis for GCC countries can be found in Kandil and Morsy (2009).

5 Empirical Investigation

5.1 Data

Our data come from a number of sources. Domestic consumer price indices are taken from the IMF's "World Economic Outlook" database (October 2009 version). Broad money (M2) and the nominal effective exchange rate index were obtained from the IMF's "International Financial Statistics" database.[19] Non-hydrocarbon output for Kuwait, Oman, Qatar and Saudi Arabia were collected from respective national sources, while for Bahrain and UAE the data was obtained from United Nation's "National Accounts Main Aggregates" database. Data on government expenditure is obtained from the "Economist Intelligent Unit" and national sources. Data for household consumption, government consumption and domestic investment (measures by gross capital formation) are obtained from the UN database mentioned earlier. Government investment is calculated by subtracting government consumption from government expenditure. National sources include "Quarterly Statistical Bulletins" and "Annual Reports". The data are available from authors on request.

In the absence of a suitable domestic import price index, we proxy for foreign prices by using an index of non-fuel commodity price index published by the IMF. Although the use of a common foreign price index constrains the six GCC countries to identical exposure to foreign price shocks, this index comes closest to mimicking the developments in international commodity prices, particularly compared to using the trading partners' consumer price index (CPI) as is frequently used in the recent literature (e.g. Kandil and Morsy, 2009).[20] Furthermore, the non-fuel commodity price index includes tradable commodities, which is relevant for the GCC countries.

5.2 Model Specification

The discussion above leads us to the following panel data model of the sources of inflation in GCC countries:

$$p_{it} = \alpha_i + \beta_{1i} p_t^* + \beta_{2i} e_{it} + \beta_{3i} m_{it} + \beta_{4i} z_{it} + e_{it} \tag{1}$$

[19] The nominal effective exchange rate for Kuwait could not be obtained from IMF's database. It was taken from Hasan and Alogeel (2008). We are grateful to Maher Hasan for supplying the index.

[20] Foreign CPI merely reflects the consumer taste of the respective foreign country and is also subject to the corresponding government's tax and subsidy policies and country specific non-tradable goods.

where $i = 1, ..., 6$ and $t = 1980, ..., 2008$ are indices of the cross-section and time series dimension, respectively; p and $p*$ are domestic and foreign price indices, respectively; e denotes the nominal effective exchange rate (NEER) of the domestic currency vis-à-vis trading partners' currencies; m represents the money supply and z is a measure of aggregate demand. The coefficients α_i are country-specific effects; β_i are the $k \times 1$ coefficient vectors and e_{it} are disturbance terms. All variables are expressed in natural logarithm. Anticipating the sign of the slope parameters, an increase in p^*, m and z are likely to increase domestic prices, while an increase (appreciation) in e is expected to reduce domestic prices. Therefore, we expect the slope parameters to obey $\beta_1 > 0$, $\beta_3 > 0$ and $\beta_4 > 0$, while $\beta_2 < 0$. Our empirical strategy is to start with some descriptive statistical analysis followed by testing the order of integration, testing for non-cointegration and finally we estimate Equation (1) and conduct appropriate statistical inference.

5.3 Correlation Analysis

Suspecting strong collinearity between some regressors, Table 2 reports the pairwise correlation coefficients between all the candidate variables of the model. As can be seen, the money supply is strongly positively correlated with the three alternative measures of aggregate demand (non-hydrocarbon GDP, government spending and private consumption). Furthermore, the three aggregate demand measures are highly correlated with each other, suggesting that each variable is an equally likely candidate measure of domestic aggregate demand.

Table 2: Pairwise correlation coefficients

Variables	p_{it}	p_t^*	e_{it}	m_{it}	y_{it}	g_{it}	pc_{it}
p_{it}	1.000						
p_t^*	0.369**	1.000					
e_{it}	0.009	-0.402**	1.000				
m_{it}	0.306**	0.282**	-0.055	1.000			
y_{it}	0.333**	0.270**	0.003	0.962**	1.000		
g_{it}	0.193**	0.227**	0.044	0.917**	0.962**	1.000	
pc_{it}	0.311**	0.280**	0.066	0.940**	0.988**	0.968**	1.000

Notes: p (domestic CPI); p^* (foreign price); e (nominal effective exchange rate); m (money supply); y (non-hydrocarbon GDP); g (government spending) and pc (private consumption). All variables are expressed in natural logarithm. ** denotes statistical significance at the 5% level.

A main reason for the high correlation between money and the different measures of aggregate demand is that, in the GCC countries, money is primarily demand determined. The two primary sources of money accumulation (and deaccumulation) in GCC countries are government expenditure and capital flows. A sizable portion of government expenditure is financed by oil revenue, which is denominated in U.S. dollar. A higher government spending thus implies a higher public deposit of foreign currencies in the banking system, which are then converted into local currencies to finance government's current and capital expenditure. Likewise, a rise in net capital inflows increases foreign currency deposits in the banking system that must be converted into local currency before any transactions (e.g. investment) are undertaken. Given the *de jure/de facto* exchange rate peg to the U.S. dollar, GCC central banks have to intervene in exchange market on demand, thus accommodate changes in money demand. Hence, the actual level of money supply in the economy is demand determined; while lacking independent monetary policies, central banks merely serve as a provider of money on demand. Hence, the strong correlation between the money supply and government expenditure originates primarily from the monetization of the economic deficit.[21]

As government expenditure is the main driver of non-hydrocarbon output growth in GCC countries, both money and non-hydrocarbon output reinforce each other.[22] Higher government expenditure (and therefore higher money supply) allows the non-hydrocarbon sector to expand its capacity on the back of imported foreign workers. In turn, higher non-hydrocarbon income raises the demand for money. Likewise, during periods of fiscal consolidation (and consequently lower money supply), import demand for foreign workers declines and as a result non-hydrocarbon sector suffers from a loss in output, which in turn reduces the demand for money. The key factor behind the money-output nexus is the flexibility of the (expatriate) labor markets, which is pro-cyclical to government expenditure. The underlying dominance of fiscal policy and the absence of effective monetary policy results in a strong correlation between government expenditure, money, non-hydrocarbon output and its components of aggregate demand (e.g. consumption, investment).[23]

[21]By economic deficit, we mean the proportion of government expenditure financed by hydrocarbon revenues less the net private sector's transactions with the rest of the world.

[22]This hypothesis is confirmed by a Granger (1969) causality test between money and non-hydrocarbon output. From results not shown here, we find that non-hydrocarbon output Granger cause money with just 1 lag, while money Granger cause non-hydrocarbon output with 2 lags.

[23]Nonetheless, as an alternative measure, we have also employed the variance inflation factor (VIF), which is a more satisfactory way of detecting (multi)collinearity between explanatory variables. The VIF_i are given by $(1 - R_i^2)^{-1}$ where R_i^2 is the R^2 from regressing the ith independent variable on all other independent variables. The VIF shows how much the variance of the coefficient estimate is being inflated by multicollinearity. A common rule of thumb in applied econometrics is that a $VIF_i > 10$ indicates harmful collinearity (Kennedy, 2003). The VIF results are in line with that of the correlation coefficients, and therefore not presented here to save space. However, these results are available in a Supplement.

5.4 Testing for Non-stationairty

Our formal analysis begins by conducting panel unit root tests for all variables used in the model. Several panel unit root tests are available to examine the stochastic properties of economic variables.[24] One key issue to keep in mind while selecting the appropriate panel unit root test, is cross-sectional dependence among units (countries) in the panel. In many aspects the six GCC economies are related to each other, and it is very likely that these economies are simultaneously affected by common observed global shocks, such as changes in oil prices or due to specific domestic or sectoral shocks. The presence of common shocks is likely to generate dependence among the units in the panel, although their impact may not be the same across different cross-section units. Therefore, we need to select panel unit root tests that are robust to cross-section dependence in order to avoid size distortions of the tests. To this end, we apply the cross-sectionally augmented ADF (CADF) statistics proposed by Pesaran (2007) and the robust t-statistic test developed by Breitung and Das (2005). Pesaran's (2007) CADF test models cross-section dependence by augmenting the standard ADF regressions for the individual series with current and lagged cross-section averages of all the series in the panel. The Breitung and Das (2005) use a pre-whitening procedure to adjust for short-run serial correlation of the errors. A brief discussion on the construction of these tests is provided in the Appendix.

Table 3: Panel unit root tests

Series	Pesaran (2007)	Breitung and Das (2005)
p_{it}	-1.168 [0.933]	3.319 [0.999]
e_{it}	-0.917 [0.983]	-0.778 [0.218]
m_{it}	-1.792 [0.467]	6.487 [1.000]
y_{it}	-2.235 [0.115]	5.258 [1.000]
g_{it}	-1.922 [0.341]	3.364 [1.000]
pc_{it}	-1.798 [0.462]	5.353 [0.999]

Notes: p (domestic CPI); e (nominal effective exchange rate); m (money supply); y (non-hydrocarbon GDP); g (government spending) and pc (private consumption). All variables are expressed in natural logarithm. All tests test the null hypothesis of a unit root. P-values are reported in [ ]. P-value greater than 0.05 indicates that the null hypothesis of a unit root cannot be rejected at the 5% significance level.

Table 3 presents the results of the panel unit root tests. Due to small sample observations per unit, a maximum of two lags is allowed. The deterministic component includes an intercept

[24]See Breitung and Pesaran (2008) for a recent overview of this growing literature.

only. Results indicate that the null hypothesis of a unit root cannot be rejected for all variables at the conventional levels of statistical significance, suggesting evidence of non-stationarity.[25] As for the foreign price variable, which is common to all cross-section units, we apply the DF-GLS unit root test of Elliott et al. (1996), which is generally more powerful than the standard Augmented Dickey-Fuller (ADF) unit root test. Following Ng and Perron's (2001) recommendation, we use the modified Akaike criterion in order to select the lag length in the test regression. The DF-GLS test statistics is -1.416, indicating that the null hypothesis of a unit root in foreign price cannot be rejected at any conventional levels of significance.[26] Given the overwhelming evidence of non-stationarity of all the variables, we next proceed to compute the panel cointegrating relationship among the variables in equation (1).

5.5 Testing for Non-Cointegration

Given the small cross-section and time dimensions, we have to be careful in selecting the appropriate panel cointegration tests that are able to fully discriminate between the null and alternative hypotheses. Furthermore, the cointegration tests should be robust to the presence of cross-section dependence. To this end, we apply several recently proposed panel non-cointegration tests that are robust to cross-section dependence. These are the four error-correction based panel non-cointegration statistics developed by Westerlund (2007) and the two residual-based stationary panel bootstrap non-cointegration statistics developed by Di Iorio and Fachin (2009). The Westerlund (2007) tests avoid the problem of common factor restrictions[27] and are designed to test the null hypothesis of no cointegration by inferring whether the error-correction term in a conditional error-correction model is equal to zero. Therefore, a rejection of the null hypothesis of no error-correction can be viewed as a rejection of the null hypothesis of no cointegration. In comparison, Di Iorio and Fachin (2009) use the nonparametric residual-based stationary bootstrap (RSB) resampling method of Parker et al. (2006) to develop panel cointegration tests. Their tests are designed to test the null hypothesis of no cointegration against four different alternative hypothesis. Nevertheless, for practical purpose only two of the four alternative

[25] Results are almost similar when unit root tests are conducted using both intercept and trend as deterministic components. The Pesaran's (2007) CADF test rejects (at 5% level) the null hypothesis of a unit root only for the non-hydrocarbon GDP; while all variables appear non-stationary according to the Breitung and Das (2005) test. These results are available in a Supplement.

[26] The DF-GLS critical values are -2.653 (1% level); -1.953 (5% level) and -1.609 (10% level).

[27] In the cointegration literature, common factor restriction suggests that the long-run parameters for the variables in their levels are equal to the short-run parameters for the variables in their first differences. A failure to satisfy the restriction can cause a significant loss of power for residual-based cointegration tests – see Kremers et al. (1992) for further discussion.

hypothesis are relevant. These are the (i) median-based ADF test which test the alternative hypothesis that the panel is cointegrated as a whole and (ii) the mean-based ADF test which describes the average of the individual statistics as panel cointegration statistic. One similarity between the Westerlund (2007) and Di Iorio and Fachin (2009) tests is that they both rely on the bootstrap methods to model dependence across cross-sectional units. The bootstrap methods allow to make inference under a very general form of cross-section dependence. A brief discussion on the construction of these tests is provided in the Appendix.

Table 4: Panel non-cointegration tests

Tests	Test statistics	Robust P-value
Westerlund (2007)	G_τ: -0.952	0.912
	G_α: -3.638	0.444
	P_τ: -1.944	0.798
	P_α: -2.459	0.458
Di Iorio and Fachin (2009)	Median (ADF): -4.090	0.024
	Mean (ADF): -4.230	0.015

Note: The panel cointegration tests are conducted on the following specification:

$$y_{it} = \alpha_i + \beta_i x_{it} + \epsilon_{it}$$

where y_{it} is the domestic consumer price index (CPI); x_{it} is a $k \times 1$ vector of explanatory variables including foreign prices, nominal effective exchange rate, money supply and a measure of aggregate demand; α_i is the country-specific effect; β_i are the $k \times 1$ coefficient vectors and ϵ_{it} are disturbances. All variables are expressed in natural logarithm. All tests are designed to test the null hypothesis of no cointegration. P-value greater than 0.05 indicates that the null hypothesis of no cointegration cannot be rejected at the 5% significance level. P-values are robust to cross-section dependence.

The Di Iorio and Fachin (2009) test is implemented using a block size of $0.10T$ (where T is the sample size) with 5000 bootstrap redrawings. The Westerlund (2007) is computationally intensive and therefore only 500 bootstrap replications are used.[28] Due to an increase in the number of estimated parameters, the maximum lag length allowed is one. Both tests allow for a deterministic component (intercept) in the cointegration relation and are designed to test the null hypothesis of non-cointegration.

Table 4 presents the panel cointegration results of equation (1), where aggregate demand is captured by non-hydrocarbon output. The results are mixed. The Westerlund (2007) tests are unable to reject the null hypothesis of non-cointegration at the conventional levels of significance;

[28]The Westerlund (2007) tests are estimated using Stata's `xtwest` routine. On Stata/IC the maximum number of bootstrap replications allowed is less than 800. Besides, with T=29 per cross-sectional unit, the maximum permissable lag length is 1. See Persyn and Westerlund (2008) for further discussion.

whereas the Di Iorio and Fachin (2009) tests reject the null hypothesis of no-cointegration at the 5% level of significance, indicating the existence of a long-run equilibrium relationship between the variables of the model.[29] One likely explanation for the conflicting results is the different ways the alternative hypothesis of these panel non-cointegration tests are formulated. For example, a rejection by Westerlund's (2007) P_τ and P_α statistics is interpreted as evidence of cointegration for the panel as a *whole*; whereas a rejection by Westerlund's (2007) G_τ and G_α statistics is interpreted as evidence of cointegration for *at least one* of the cross-sectional units. By comparison, a rejection by Di Iorio and Fachin's (2009) tests is taken as evidence of cointegration for *most* of the units. Thus, the conflicting test outcomes as reported in Table 4 could be a reflection of the variations in test construction considered in this study. In addition, the Westerlund (2007) tests require comparatively large T dimension than that of Di Iorio and Fachin (2009).[30]

This leads us to examine the cointegration relationship for each cross-section by employing conventional univariate time series cointegration tests. To this end, we apply the well-known residual-based non-cointegration test of Engle and Granger (1987) and the system-based tests of Johansen (1988). The former tests the null hypothesis of no cointegration against the alternative of at least one cointegrating relationship, whereas the latter test tests the null hypothesis of no cointegration against r cointegrated relationships. For comparison sake, we consider $r = 0$ and employ the maximum eigenvalue test of Johansen (1988). Again, the lag length is set at 1 and only an intercept is included as the deterministic component.[31]

Table 5 reports the test results of time series non-cointegration tests, where the aggregate demand is captured by non-hydrocarbon output. We get conflicting results. The Engle and Granger (1987) test shows no evidence of cointegration, while the Johansen (1988) test strongly indicates the presence of cointegration. However, this kind of conflicting outcomes among individual cointegration tests is not surprising. For example, conducting various residual- and system-based cointegration tests on 134 data sets, Gregory et al. (2004) find that p-values of different tests are typically not perfectly correlated. Furthermore, Pesavento (2004) demonstrates that the tests differ in their power in different parts of the parameter space. Such mixed

[29] This finding is not affected when the aggregate demand is measured by either government spending or private demand. These results are reported in a Supplement.

[30] Test outcomes can significantly differ due to different choices of adjustment methods (e.g. treatment of serial correlation) used in cointegration testing. For a demonstration of this issue in the context of panel data, see Westerlund and Basher (2008).

[31] Test results with both intercept and trend are available in a Supplement.

Table 5: Time series tests of non-cointegration

Country	Engle-Granger	Johansen	Engle-Granger & Johansen combination test
Bahrain	-2.406	44.546***	12.985**
Kuwait	-1.809	39.126**	9.207*
Oman	-4.135*	55.788***	60.014***
Qatar	-2.358	38.195**	8.624*
Saudi Arabia	-2.792	34.599**	7.121
U.A.E.	-3.599	48.117***	17.481***

Note: Individual or combine cointegration tests are conducted on the specification:

$$y_t = \alpha + \beta x_t + \epsilon_t$$

where y_t is the domestic CPI; x_t is a $k \times 1$ vector of explanatory variables including foreign prices, nominal effective exchange rate, money supply and a measure of aggregate demand; α is the deterministic component (constant); β is the $k \times 1$ coefficient vectors and ϵ_t are disturbances. All variables are expressed in natural logarithm. The Engle-Granger & Johansen combination test is the Fisher χ^2 test developed by Bayer and Hanck (2009). All tests test the null of no cointegration. ***, ** and * denote statistical significant at the 1%, 5% and 10% level, respectively.

signals among cointegration tests are problematic since it is not clear how to interpret test outcomes.

Against the backdrop of conflicting test results, Bayer and Hanck (2009) devise an innovative method to propose a combined test by merging the Engle-Granger and Johansen tests. By combining the non-cointegration tests, they are able to control for the size of the test and obtain an unambiguous test decision. The two tests are combined using the well-known Fisher aggregator (the famous Fisher's (1932) χ^2 test). The authors propose both asymptotic and bootstrap combination tests. Simulation results show that the proposed Fisher-type test performs very well and follows closely (or even exceeds) the power envelope of the underlying single tests. In particular, the bootstrap versions of combination tests appear to converge to the nominal size somewhat more quickly, while the (size-adjusted) power of the bootstrap versions is very similar to that of the asymptotic tests throughout. The last column in Table 5 reports the Bayer and Hanck (2009) Fisher-type combination test. As can be seen, save for Saudi Arabia, the null hypothesis of no cointegration is rejected at the 10% significance level or lower for all countries, indicating the existence of a long-run relationship between domestic CPI and its determinants. For Saudi Arabia, the Johansen test strongly rejects the null hypothesis of non-cointegration. These individual-based cointegration results complement the finding of the panel cointegration based on the Di Iorio and Fachin (2009) statistics.[32,33]

5.6 Estimation Results

Several panel estimators are available in the literature to estimate multiple cointegrating regressions. Our interest centers on error-correction type models that allow one to obtain both short- and long-run estimates. In particular, the estimated error-correction term of a properly specified model reveals the expectations and adjustment processes of an economy that is crucial for the effective evaluation of, say, the impact of fiscal and monetary policies on prices. To that end, we apply the mean group (MG) estimator of Pesaran and Smith (1995) which is designed to estimate both short- and long-run relationships in dynamic heterogenous panels. The MG estimator is obtained by estimating separate cointegrated equations for each group followed by a simple arithmetic average of the estimated coefficients. Thus, under the MG estimator the intercepts, slope coefficients and error variances are allowed to differ across cross-sectional units in the panel. The MG estimator performs reasonably well for large T relative to N, which is fairly satisfied by our data.[34] A brief discussion of the MG estimator is presented in the Appendix.

Table 6 presents the estimation results for three different regression specifications. In all cases, the dependent variable is the natural logarithm of domestic CPI. The level of significance of the estimated coefficients are based on a one-tailed test, as we have strong a priori that the relationship between inflation and its determinants (e.g. foreign prices, money supply, aggregate demand and the negative of the exchange rate) is positive and, thus, can run $H_0 : \beta \leq 0$ against $H_1 : \beta > 0$. Whereas for the exchange rate variable, the null and alternative hypothesis is defined the opposite way. Note that, for all three models the LR statistic strongly rejects the null hypothesis of cointegrating vector homogeneity, suggesting that pooling is not feasible.

In Model 1, the estimated long-run coefficients bear the correct signs, but only the exchange

[32] The fact that cointegration is evident individually but wiped out after aggregation (i.e. panel), is quite well-known – see Trapani and Urga (2010) and the references therein for related discussion. Pesaran and Smith (1995) show that aggregation of heterogenous cointegrating equations does not imply cointegration in the aggregate relationship unless some specific conditions are satisfied.

[33] The individual cointegration results are quite similar when the aggregate demand is measured by either government spending or private consumption. However, the evidence of cointegration is slightly more stronger for the former indicator than the latter. These results are reported in a Supplement.

[34] Pesaran, Shin and Smith (1999) also propose a pooled mean group (PMG) estimator that combines both pooling and averaging. However, the PMG estimator constrains the long-run coefficient to be equal across groups (just like a conventional fixed effects estimator) and henceforth not used in this paper. Nonetheless, a limitation of the MG estimator is that it does not allow for residuals to be contemporaneously correlated. The common correlated estimator (CCE) of Pesaran (2006) avoids this problem, but requires N to be substantially larger than T.

Table 6: Mean group (MG) estimation of long- and short-run equations

Series	Model 1	Model 2	Model 3
$\ln p^*_{t-1}$	0.049	0.519**	0.626**
$\ln e_{it-1}$	-0.560**	-0.729***	-0.705***
$\ln m_{it-1}$	0.272*	0.161	0.245**
$\ln y_{it-1}$	0.029		
$\ln g_{it-1}$		0.177***	
ϕ_i	-0.180***	-0.212***	-0.211***
$\Delta \ln p^*_{t-1}$	-0.136	-0.126	-0.116
$\Delta \ln e_{it-1}$	-0.074	-0.008	0.010
$\Delta \ln m_{it-1}$	0.028*	0.036*	0.046*
$\Delta \ln y_{it-1}$	0.074*		
$\Delta \ln g_{it-1}$		0.009	
Constant	0.394**	0.515**	0.521***
LR statistic	63.785***	69.7600***	52.848***
Log-likelihood	472.182	460.943	444.108
SBIC	-918.745	-896.270	-862.597
$N \times T$	168	168	168

Note: The dependent variable is domestic CPI. p^*_t represents foreign prices; e_{it} is the nominal effective exchange rates; m_{it} denotes money supply; y_{it} is the non-hydrocarbon output and g_{it} denotes government spending. ϕ_i is the error correction speed of adjustment term. Δ denotes the first difference operator. All variables are expressed in natural logarithm. ***, ** and * denote statistical significant at the 1%, 5% and 10% level, respectively.

rate and money supply appear as statistically significant at 5% and 10% level, respectively. This is likely because of – as elaborated in Section 5.1 – the high collinearity between money and non-oil output, which makes it difficult to disentangle the effect of money and output on prices. By comparison, the estimated short-run coefficients are more illuminating. The error-correction coefficient has the expected negative and highly significant sign. Both money supply and aggregate demand (non-oil output) stand as significant determinants of inflation in the short-run, whereas foreign prices and the exchange rate do not appear as significant sources of inflation in the short-run.

Model 2 replaces non-oil output with government spending, which is considered here as an alternative measure of aggregate demand. The results of the long-run coefficients are noticeably different than Model 1. The sign of the estimated coefficients of all four explanatory variables are correct, and save for the money supply, they are statistically significant at the 5% level.[35] The difference in results between Model 1 and 2 is possibly due to the lower degree of collinearity

[35]The specification where aggregate demand is measured by private consumption is quite similar to Model 2 and therefore not presented here to save space. These results are presented in a Supplement.

between money and government spending, relative to the correlation between money and non-hydrocarbon output. For example, the variance inflation factor (VIF) for non-oil output is 14.36 in the regression of money supply on non-oil output and other explanatory variables, which is well above the threshold level of 10 (Kennedy, 2003). Whereas, the VIF for government spending is 6.74 (within the cut-off limit of 10) when the money supply is regressed on government spending and other independent variables. Generally, the confidence interval of the parameters are very wide when regressors are highly collinear leading to high standard errors or very low t-ratios. The error correction coefficient suggests a quicker speed of adjustment, relative to Model 1. The short-run parameters are almost identical in direction and magnitude as in Model 1, however unlike non-oil output, government spending does not exert a significant effect on domestic prices in the short-run.[36]

Nonetheless, the results of Model 2 should be interpreted carefully. In the GCC context, the money supply encompasses the monetization of the net impact of government spending and interest rate changes by the Federal Reserve, thereby, making the money supply a key "resource" variable of the economy. Hence, the simultaneous use of money supply and government spending in the regressions can be problematic. Moreover, as discussed earlier that the money supply is primarily demand determined implies that the money supply independently represents the bulk of domestic demand in the GCC region. These economic features along with the statistical evidence of high collinearity between the money supply and aggregate demand measure(s) have led us to drop the aggregate demand variable from the model.[37]

Accordingly, Model 3 in Table 6 presents estimation results of a very parsimonious model. Results reveal that foreign prices, the exchange rate and money supply are the most significant long-run determinants of inflation in GCC countries. Some remarks are in order. First, ceteris paribus, a one percentage point increase in foreign prices cause a 0.62 percentage points increase in the domestic price level. The comparable estimate documented by Hasan and Alogeel (2008) is 0.78 for Kuwait and 0.83 for Saudi Arabia; whereas, Kandil and Morsy (2009) obtained a coefficient of 0.48 for Kuwait and 1.92 for Saudi Arabia—the results for remaining GCC states stand somewhere between Kuwait and Saudi Arabia. One similarity between the two studies is that they both use trading partners' consumer price indices (CPIs) to measure the pass-

[36]Following Kandil and Morsy (2009), we examined how the composition of public spending affect the long-run price levels. As expected, public consumption (investment) is positively (negatively) related to price levels, although these effects are not statistically significant. These unreported results are available in a Supplement.

[37]Results deteriorate when money supply (instead of non-oil output) is dropped from the model as both foreign prices and exchange rates become statistically insignificant.

through effect. As foreign CPIs include foreign nontradable goods, these estimates should be interpreted with caution. Nonetheless, one caveat with our choice of non-fuel commodity price index is that although it is a fully tradable price index, it essentially represents the export (or producer) prices of tradable commodities and does not include transportation costs and the cost of freight insurance. These features can play a significant role in magnifying the pass-through effect into domestic prices. For example, examining the import market for corn in Qatar, Raboy et al. (2010) observe that over the 2004-2008 period the average difference between the f.o.b. (any destination) per tonne price of Argentine corn and the c.i.f. import price was about US$200/tonne higher in Qatar than any other destinations.[38] These elements are likely to be present in other GCC States (albeit in varying degree) given the comparable small market size and strong reliance on imports. Hence, an appropriately constructed domestic import price index is the ideal variable for quantifying the impact of international prices on domestic prices.

Second, all else constant, a one percentage point appreciation of the domestic currency vis-à-vis trading partners' currencies causes a 0.70 percentage points decrease in the domestic price level. In contrast, Hasan and Alogeel (2008) obtained a very moderate pass-through effect of around 0.20 for Kuwait and Saudi Arabia, which they consider is comparable to the estimated pass-through effects in developed countries (Goldberg and Campa, 2010) or emerging market economies (Mihaljek and Klau, 2008). It is well-known that the extent of pass-through of exchange rate changes to consumer prices is lower than to import prices.[39] However, there are several reasons why this result may not hold in the context of GCC economies. First, the lack of CPI responsiveness to exchange rate changes is partly because of the ability of the monetary authorities in developed/emerging countries to use contractionary monetary policy in the event of a domestic currency depreciation (Gagnon and Ihrig, 2004). This monetary policy flexibility is largely absent in GCC countries due to the *de jure/de facto* exchange rate peg to the U.S. dollar. Second, the extent of local-value-added components[40] to the final consumption value of imports generates a wedge between border and retail prices, implying that CPI inflation reacts less to changes in exchange rates – see, among others, Burstein et al. (2002).

[38] For example, according to U.S. Department of Agriculture (USDA, 2009), the per-tonne freight charges of shipping corn from Argentina (Rio Plate) to Algeria were US$40/tonne during the second quarter of 2009. USDA also reported that shipping rates for corn from the Gulf Coast of the U.S. to Japan were on the order of US$49/tonne, and US$27/tonne to Japan from the U.S. Pacific Northwest.

[39] See Goldberg and Knetter (1997) for a comprehensive survey of the literature on goods prices and exchange rates.

[40] These include expenditures on transportation, storage, finance, insurance, marketing, wholesaling and retailing.

This happens partly because of the substitution in consumption away from costly imports to lower quality local goods and partly because distributors/retailers, to some extent, absorb exchange rate fluctuations in their own margin in order to maintain stable prices or expand market share at the retail level – see, e.g. Goldberg and Campa (2010). However, domestic substitutes of imported goods are not a feasible option for GCC countries due to the very limited domestic production base.[41] Moreover, even though distribution-related costs incorporate a significant share of local value added in GCC countries, local prices may not be insulated from exchange rate movements due to the observed departure from perfect competition and/or the prevailing monopolistic competition in the distributors/retailers segment (as assumed by these open-economy macroeconomic models). The limited availability of domestically produced goods or a larger share of imports in total industry supply leads to a substantially higher pass-through effect of exchange rates into import prices and subsequently into consumer prices. As the share of tradable goods to nontradable goods in the local consumer baskets is much larger, coupled with the market deficiencies just mentioned, the direct transmission of import price movements into consumer prices is predicted to be much stronger in GCC economies than advanced or emerging countries.[42] These non-standard features of the GCC economies play a role in magnifying (rather than damping) the pass-through into consumer prices. The fact that our point estimate of 0.70 falls short from complete pass-through effect can be explained by the prevalence of subsidies or administered control prices in the GCC countries.[43]

Third, all else constant, a one percentage point increase in domestic money supply results in a 0.24 percentage points increase in the domestic price level. This monetary phenomenon of inflation is expected and is consistent with the findings of previous studies. To re-emphasize the dominant role of money supply in GCC economies, recall that, as government oil revenues are generated in foreign exchange, money supply essentially captures the monetization of government spending on domestic economy. Secondly, as already discussed in Section 4, the money supply also incorporates the interest rate channel as the effect of changes in Federal funds rate is transmitted to domestic money supply through identical changes in domestic interest rates.

[41] It should be noted that, domestic production diversification differs across the six GCC states with a comparably more facilities in Oman and Saudi Arabia than other member states.

[42] Of course, this claim needs to be complemented by empirical evidence. Nevertheless, the fact that exchange rate pass-through is generally higher in highly dollarised economies (this is related to the idea of producer currency pricing) is consistent with our high estimate of the exchange rate pass-through effect. See, Campa et al. (2007) for a partial list of papers concerning currency invoicing and exchange rate pass-through.

[43] For example, the pass-through from international to domestic retail gasoline prices between end-2003 and mid-2008 was highest for low-income countries (70%) and recorded lowest for GCC countries (25%). See IMF (2008b) for further details.

Thirdly, the money supply also captures the impact of net capital flows in GCC countries as, for example, an inflow of capital may cause an increase in domestic money supply. On balance, the role of money supply as a resource variable is deemed far more crucial to the GCC economies, as compared to more diversified developed and developing economies.[44]

Finally, the error correction speed of convergence parameter is negative and significant. The error correction parameter becomes easier to interpret when we transform it to the half-life, $\ln(0.5)/\ln(1+\phi_i)$, implying that it takes about 2.9 years for 50% of a shock to the long-run equilibrium to dissipate. This appears a reasonable estimate of the speed of adjustment when we look at the average duration of inflationary episodes in GCC countries. For example, during the 2004-2008 episode of high inflation when the price level persistently increased on the back of high oil prices, the almost 6% annual inflation during the 2004-2008 period was well above GCC's long-run average of 2.82% over the 1980-2008 period. Inflation came down in 2009, with the 3% average annual inflation settled roughly close to the long-run level.[45] Furthermore, the fact that inflation in GCC countries are routinely influenced by the exogenous factors also appear consistent with the half-life statistic. Overall, the estimated coefficients of the short-run determinants of inflation did not produce economically interesting results.

6 Concluding Remarks

The goal of this paper has been to empirically analyze the sources of inflation in GCC countries. Several important results emerge from the analysis. First, the money supply is reported as a significant determinant of inflation both in short- and long-run. This suggests that GCC countries can benefit from having their own monetary policies rather than the importing monetary policy of the Federal Reserve. This implication is further complemented by the finding that the exchange rate plays a decisive role in the determination of inflation in the long-run. Therefore, sovereign monetary policies would allow the GCC central banks to use their own

[44]Following the recent literature on GCC, we have augmented Model 3 with a measure of excess money supply or excess demand in the short-run equations. These excess measures are obtained by applying the Hodrick and Prescott (1997) filter, with a penalty parameter of 6.25 for annual data. Unreported results show that excess money supply has an incorrect negative sign and is statistically insignificant. Conceptually, because of the fixed exchange rate system, excess money supply cannot persist over the course of a year. Excess money supply, relative to its demand, is corrected by an increase in inflation. By comparison, excess demand (defined as the sum of private and public consumption and domestic investment) positively affects prices in the short-run and is statistically significant at the 1% level. The inclusion of excess demand in Model 3 does not change the overall results qualitatively. These results are reported in a Supplement.

[45]It is worth mentioning that, to a considerable degree, the 2008-2009 financial crisis expedited the speed of convergence. Nonetheless, our estimate of the speed of convergence is different from Hasan and Alogeel (2008) and Kandil and Morsy (2009), who documented faster estimates of the error correction coefficient.

exchange rate policies to effectively deal with the deviations from long-run price equilibrium. One would expect a much quicker speed of convergence under a more independent monetary policy environment than that of the estimated 2.9 years under the existing pegged exchange rate system. The finding that foreign prices positively contributed to domestic price level provides further support for using exchange rate policies to partially sterilize the impact of international price shocks.

We wish to emphasize, at this point, that the results presented in this paper are in no way intended to be definitive estimates of inflation dynamics in GCC countries. There is a great deal of room for further research, particularly in the areas of (1) improvements in the econometric methods for inferring robust estimates for panels with small cross-section and time dimension; (2) construction of import price indexes to accurately gauge the effect of international price shocks; and (3) the testing and use of alternative model specifications to unearth the precise sources of inflation in the constituent GCC countries. We offer in this paper a sample analysis of what can be achieved with available published data and existing methodologies.

A Statistical appendix

A.1 Pesaran (2007) CADF Panel Unit Root Test

To briefly summarize the construction of the CADF test, let us consider a heterogenous, linear panel data model

$$Y_{it} = (1 - \rho_i)\mu_i + \rho_i Y_{it-1} + u_{it}, \tag{A.1}$$

where the error term u_{it} has a common factor structure: $u_{it} = \gamma_i f_t + e_{it}$, where f_t is the unobserved common factor, γ_i is the corresponding factor loading and e_{it} is the individual-specific (idiosyncratic) error term. It is convenient to rewrite as

$$\Delta Y_{it} = \alpha_{0i} + \alpha_{1i} Y_{it-1} + \gamma_i f_t + e_{it}, \tag{A.2}$$

where $\alpha_{01} = (1 - \rho_i)\mu_i$ and $\alpha_{1i} = (\rho_i - 1)$. To account for the cross-sectional dependence induced by the common factor, Pesaran (2007) suggests to cross-sectionally augment the test equation (A.2) with cross-sectional averages of the first differences and the lagged levels. The cross-sectionally augmented DF (CADF) equation is then given by

$$\Delta Y_{it} = a_i + b_i Y_{it-1} + c_i \overline{Y}_{t-1} + d_i \overline{Y}_t + \varepsilon_{it}, \tag{A.3}$$

where $\overline{Y}_{t-1} = \sum_{i=1}^{N} Y_{it-1}$, $\Delta\overline{Y}_t = \sum_{i=1}^{N} \Delta Y_{it}$ and ε_{it} is the regression error. The individual specific test statistic for the hypothesis $H_{0i} : \rho_i = 1$ for a given i is now the t-statistic of b_i in (A.3), denoted by $CADF_i$. The panel unit root for the hypothesis $H_0 : \rho_i = 1$ for all i against the heterogenous alternative $H_1 : \rho_i < 1$ for some i is given by the cross-sectional average of the $CADF_i$ tests, such that

$$\overline{CADF} = N^{-1} \sum_{i=1}^{N} CADF_i.$$

As the $CADF_i$ statistics has no finite first and second moments, Pesaran (2007) suggests replacing extreme values of the test statistics by K_1 and K_2 such that $Pr[-K_1 < CADF_i < K_2]$ is sufficiently large, namely in excess of 0.9999. Pesaran (2007) simulates K_1 and K_2 for 3 types of model depending on the deterministic part (no deterministic; intercept; intercept and linear trend) and provides critical values for the test statistics obtained via stochastic simulation.

A.2 Breitung and Das (2005) Robust t-Statistic Panel Unit Root Test

Breitung and Das (2005) study the behavior of several panel unit root tests when cross-sectional dependence is present in the data. They consider that the data is generated by a simple $AR(1)$ process

$$\Delta y_{it} = \phi y_{it} + \varepsilon_{it}, \tag{A.4}$$

where $\Delta y_{it} = y_{it} - y_{it-1}$ and $\phi = \alpha - 1$. The robust t-statistic tests for the unit root null hypothesis $\phi = 0$ against the homogenous alternative $\phi < 0$. The robust t-statistic is given by

$$t_{rob} = \frac{\sum_{t=1}^{T} y'_{t-1} \Delta y_t}{\sqrt{\sum_{t=1}^{T} y'_{t-1} \widehat{\Omega} y_{t-1}}},$$

where $\widehat{\Omega} = \frac{1}{T} \sum_{t=1}^{T} (\Delta y_t - \widehat{\phi} y_{t-1})(\Delta y_t - \widehat{\phi} y_{t-1})'$.

To adjust for short-run serial correlation of the errors, Breitung and Das (2005) relied upon a pre-whitening procedure similar to Chang (2004). For the short-run dynamics, they assume that y_{it} has an $AR(p)$ representation. Following Chang (2004), the individual specific autoregressive parameters were removed within a first step regression such that $e_{it}(v_{it-1})$ are the residuals from a regression of $\Delta y_{it}(y_{it-1})$ on $\Delta y_{it-1}, ..., \Delta y_{it-p}$. In the second step, the common parameter ϕ is estimated from a pooled regression: $e_{it} = \phi v_{it-1} + v_{it}$. However, the short-run parameters are not removed from v_{it-1} in the first step regression. To cope with this difficulty, Breitung and Das (2005) suggest, under the null hypothesis, pre-whitening y_t as

$$y_t^* = y_t - \Gamma_1 y_{t-1} - ... - \Gamma_p y_{t-p}$$

is a vector random walk and Δy_t^* is white noise with $E(y_t^* y_t^{*'}) = \Omega$; and Γ_j for $j = 1, ..., p$, are the diagonal matrices containing the autoregressive parameters associated with $\Delta y_{it-1}, ..., \Delta y_{it-p}$. Accordingly, the robust t-statistic is applied to the "pre-whitened" series y_t^*.

A.3 Westerlund (2007) Error Correction Based Panel Cointegration Tests

Westerlund's (2007) error correction tests assume the following data generating process

$$\Delta y_{it} = \delta_i' d_t + \alpha_i (y_{it-1} - \beta_i' x_{it-1}) + \sum_{j=1}^{p_i} \alpha_{ij} \Delta y_{it-j} + \sum_{j=-q_i}^{p_i} \gamma_{ij} \Delta x_{it-j} + e_{it}, \tag{A.5}$$

where d_t contains the deterministic components (none, intercept, intercept and trend) and the parameter α_i is the error-correction term. The construction of the group means tests proceeds in three steps. The first step is to estimate (A.5) by least squares for each unit i and obtain $\widehat{e}_{it}$ and $\widehat{\gamma}_{ij}$. The second step is to compute

$$\widehat{u}_{ij} = \sum_{j=-q}^{p_i} \widehat{\gamma} ij \Delta x_{it-j} + \widehat{e}_{it}, \tag{A.6}$$

which is then used to obtain $\widehat{\alpha}_i(1) = \frac{\widehat{\omega}_{ui}}{\widehat{\omega}_{yi}}$, where $\widehat{\omega}_{ui}$ and $\widehat{\omega}_{yi}$ are the usual Newey-West long-run variance estimators based on $\widehat{u}_{it}$ and Δy_{it}, respectively. The third step is to compute the group mean tests in the following way

$$G_\tau = \frac{1}{N} \sum_{i=1}^{N} \frac{\widehat{\alpha}_i}{SE(\widehat{\alpha}_i)}, \quad \text{and} \quad G_\alpha = \frac{1}{N} \sum_{i=1}^{N} \frac{T\widehat{\alpha}_i}{\widehat{\alpha}_i(1)},$$

where $SE(\widehat{\alpha}_i)$ is the conventional standard error of $\widehat{\alpha}_i$.

The first step in computing the panel tests is the same as for the group mean tests. Regressing Δy_{it} and y_{it-1} onto d_t, the lags of Δy_{it} as well as the contemporaneous and lagged values of Δx_{it} yields the projection errors

$$\Delta \widetilde{y}_{it} = \Delta y_{it} - \widehat{\delta}_i' d_t - \widehat{\lambda}_i' x_{it-1}) - \sum_{j=1}^{p_i} \widehat{\alpha}_{ij} \Delta y_{it-j} - \sum_{j=-q_i}^{p_i} \widehat{\gamma}_{ij} \Delta x_{it-j}, \tag{A.7}$$

and

$$\widetilde{y}_{it-1} = y_{it-1} - \widetilde{\delta}_i' d_t - \widetilde{\lambda}_i' x_{it-1}) - \sum_{j=1}^{p_i} \widetilde{\alpha}_{ij} \Delta y_{it-j} - \sum_{j=-q_i}^{p_i} \widetilde{\gamma}_{ij} \Delta x_{it-j}. \tag{A.8}$$

The second step is to make use of $\Delta\widetilde{y}_{it}$ and $\widetilde{y}_{it}$ in estimating the common error correction parameter α and its standard error. In particular, we compute

$$\widehat{\alpha} = \left(\sum_{i=1}^{N}\sum_{t=2}^{T}\widetilde{y}_{it-1}^{2}\right)^{-1}\sum_{i=1}^{N}\sum_{t=2}^{T}\frac{1}{\widehat{\alpha}_i(1)}\widetilde{y}_{it-1}\Delta\widetilde{y}_{it}.$$

The third step is to compute the panel statistics as

$$P_\tau = \frac{\widehat{\alpha}}{SE(\widehat{\alpha})}, \quad \text{and} \quad P_\alpha = T\widehat{\alpha},$$

where $SE(\widehat{\alpha}) = \left((\widehat{S}_N^2)^{-1}\sum_{i=1}^{N}\sum_{t=2}^{T}\widetilde{y}_{it-1}^2\right)^{-1/2}$, where $\widehat{S}_N^2 = \frac{1}{N}\sum_{i=1}^{N}\frac{\widehat{\sigma}_i}{\widehat{\alpha}_i(1)}$ with $\widehat{\sigma}_i$ being the estimated regression standard error in equation A.5.

A.4 Di Iorio and Fachin (2009) Bootstrap Panel Cointegration Tests

Consider the following data generating process

$$y_{it} = \alpha_i + \beta_i x_{it} + \epsilon_{it}, \tag{A.9}$$

with $\epsilon_{it} = \rho_i\epsilon_{it-1} + \nu_{it}$. It is immediately seen that when H_0 : no cointegration holds $\rho = 1$, while it does not $|\rho| < 1$. Denoting by G a summary statistic of the no cointegration tests for the individual units (e.g. $G = N^{-1}ADF_i$, or $G = Median(ADF_1, ..., ADF_n)$, where ADF_i is the ADF statistic computed on the residuals of the i-th cointegrating regression. The proposed bootstrap procedure include eight simple steps:

1. Compute $\widehat{\nu}_{it} = \widehat{\epsilon}_{it} - \widehat{\rho}_i\widehat{\epsilon}_{it-1}$, where $\{\widehat{\epsilon}_{it}\}$ are the residuals and $\widehat{\rho}_i$ are the OLS estimate of ρ_i.
2. Resample the series $\{\widehat{\nu}_{it}\}$ via the stationary bootstrap to the entire $T \times N$ matrix of the residuals, obtaining $\{\widehat{\nu}_{it}^*\}$.
3. Compute $\{\widehat{\nu}_{it}^*\}$ obtaining pseudoresiduals $\{\widehat{\epsilon}_{it}^*\}$ obeying the null hypothesis of no cointegration.
4. Compute $y_{it}^* = \widehat{\alpha}_i + \widehat{\beta}_i x_{it} + \epsilon_{it}^*$.
5. Estimate the cointegrating regression on the dataset $\{y_{it}^*, x_{it}^*\} : y_{it}^* = \widehat{\alpha}_i^* + \widehat{\beta}_i^* x_{it} + \epsilon_{it}^*$.
6. Estimate the AR(1) coefficient ρ_i^* for the residuals $\{\widehat{\epsilon}_{it}^*$.
7. Repeat 2-6 a large number (say, B) of times.
8. Compute the bootstrap significance level, assuming that the rejection region is the left tail of the distribution, $p^* = prop(G^* < \widehat{G})$.

A.5 Pesaran and Smith (1995) Mean Group Estimator

Assume an autoregressive distributive lag (ARDL)$(p, q_1, ..., q_k)$ dynamic panel specification of the form

$$y_{it} = \sum_{j=1}^{p}\lambda_{ij}y_{it-j} + \sum_{j=0}^{q}\delta_{ij}' x_{it-j} + \mu_i + \epsilon_{it} \tag{A.10}$$

where y_{it} is the domestic CPI; x_{it} is a $k \times 1$ vector of explanatory variables; δ_{ij} are $k \times 1$ coefficient vectors; λ_{ij} are scalars; μ_i represents the country-specific fixed effect and ϵ_{it} are disturbances. It is common to reparameterizing (A.10) to obtain the error correction model

$$\Delta y_{it} = \phi_i(y_{it-1} - \theta_i^{'} x_{it-1}) + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta y_{it-j} + \sum_{j=0}^{q-1} \delta_{ij}^{'*} \Delta x_{it-j} + \mu_i + \epsilon_{it} \tag{A.11}$$

where $\phi_i = -(1 - \sum_{j=1}^{p} \lambda_{ij})$, $\theta_i = \sum_{j=0}^{q} \delta_{ij}/(1 - \sum_k \lambda_{ik})$, $\lambda_{ij}^{*} = -\sum_{m=j+1}^{p} \lambda_{im}$ $j = 1, 2, ..., p-1$, and $\delta_{ij}^{*} = -\sum_{m=j+1}^{q} \delta_{im}$ $j = 1, 2, ..., q-1$. Notice the foregoing error correction equations are written in terms of current, rather than lagged levels of the exogenous regressors. This allows an ARDL(1,0,0) as a special case. The parameter ϕ_i is the error correction speed of adjustment term. If $\phi_i = 0$, there is no error correction and thus no evidence of a long-run relationship. ϕ_i is expected to be significantly negative under the prior assumption that the variables show a return to a long-run equilibrium. The vector $\theta_i^{'}$ contains the long-run relationships between the variables. The MG parameters are simply the unweighted means of the individual coefficients. For example, the MG estimate of the error correction coefficient, ϕ, is

$$\hat{\phi} = N^{-1} \sum_{i=1}^{N} \hat{\phi}_i \tag{A.12}$$

with the variance

$$\hat{\Delta}_{\hat{\phi}} = \frac{1}{N(N-1)} \sum_{i=1}^{N} (\hat{\phi}_i - \hat{\phi})^2 \tag{A.13}$$

The means and variance of other short-run coefficients are similarly estimated.

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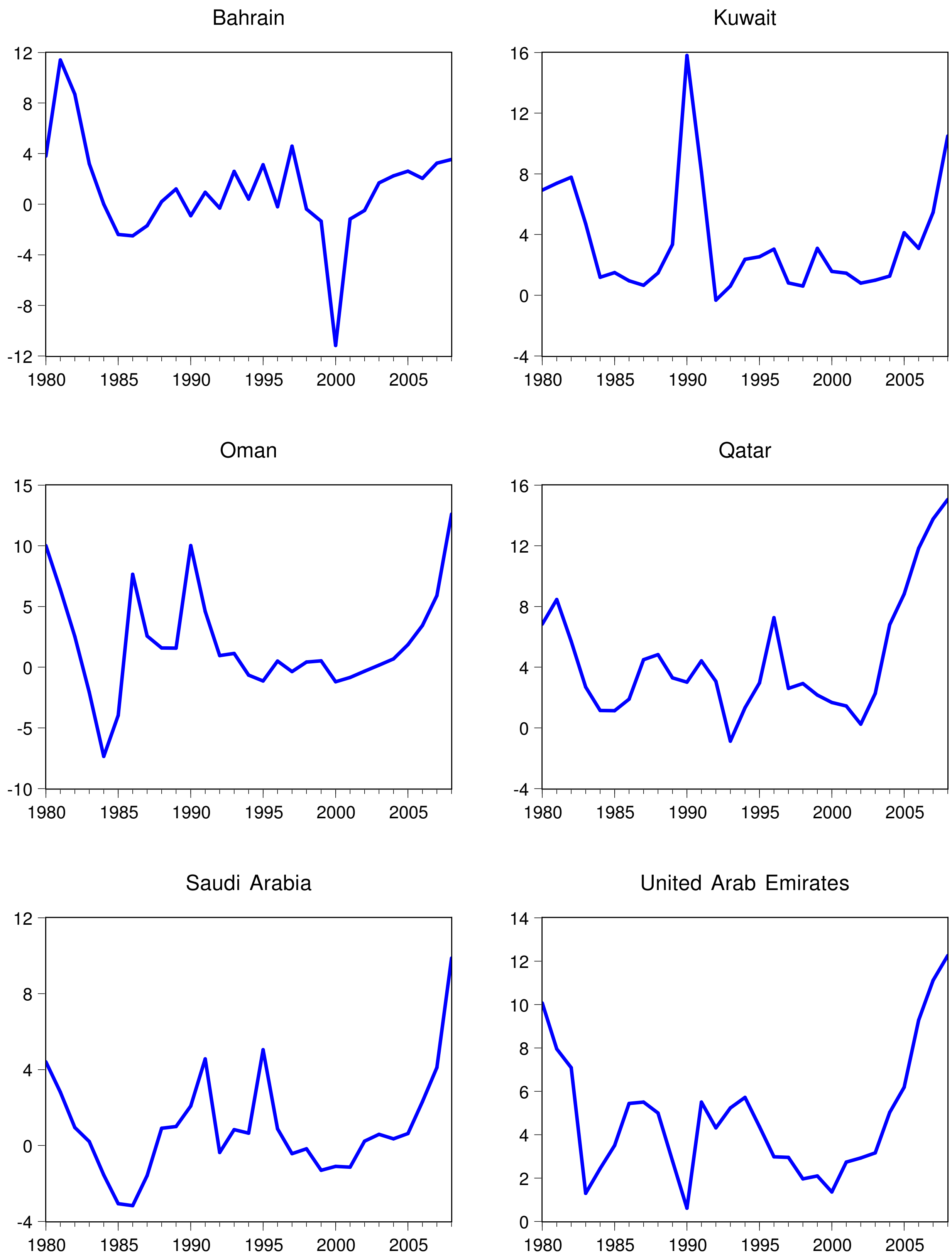
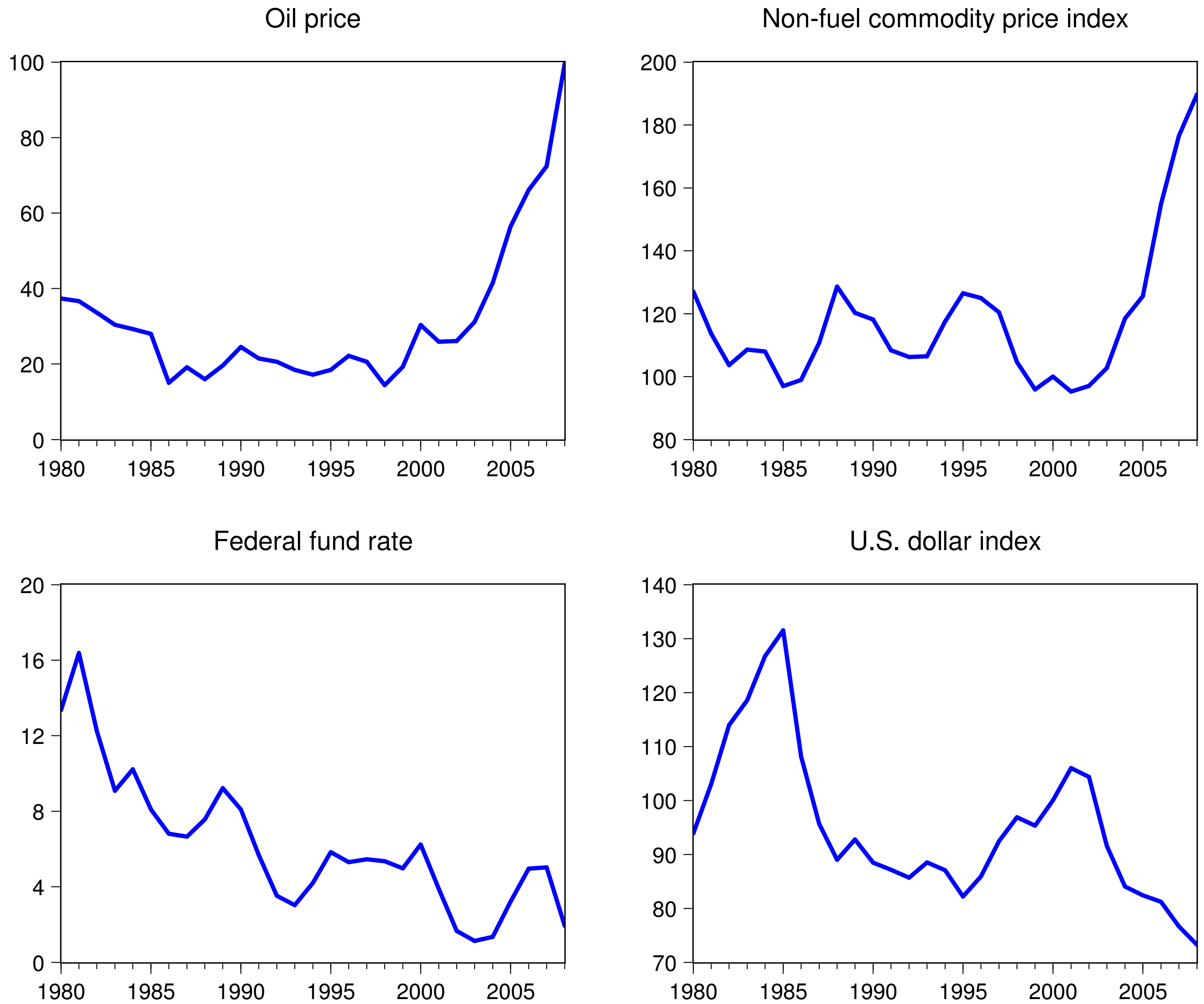
Figure 1: Inflation rates in GCC countries: 1980-2008



Note: Inflation, average consumer prices (annual percent change). World Economic Outlook Database (October 2009 version), International Monetary Fund.

Figure 2: Oil price, commodity price, federal fund rate and U.S. dollar index: 1980-2008



Note: Spot oil price per barrel, West Texas Intermediate; federal fund rate (percent) and U.S. dollar trade weighted index of major currencies are obtained from Economic Data – FRED, Federal Reserve Bank of St. Louis. Non-fuel commodity price index is taken from IMF's Primary Commodity Prices.