# Egyptian and Syrian commodity markets after the dissolution of the Ottoman Empire: a Bayesian structural VECM analysis

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**Abstract**

The disruption of the Ottoman Empire caused dramatic changes to the economic and political structure of the Middle East. The newly established nations, incorporated into British and French formal and informal empires, actively implemented a range of protectionist policies, thus disrupting the regions traditional trade flows and patterns. This paper investigates the impact of this new economic setting on commodity market integration in Syria and Egypt, using Bayesian inference. After testing for co-integration through the calculation of Bayes factors and computing impulse response functions, our results point to the absence of cross border market integration.

*Keywords:* Market integration, Middle East, interwar, Cointegration, Posterior probability
*JEL Classification:* N15, N75, N95, C11, C12, C32,

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*The ancestors of the London bankers were still roaming the wilds with clubs in their hands, when the Phoenician sails were plying a prosperous trade between Syria and Egypt. The Phoenician sails have long since gone beyond the horizon but the Syro-Egyptian trade continues. Twenty-five centuries of commercial relations bind the two countries together.*

*Burns (1933, p.82)*

# 1 Motivation

After the end of World War I, the former unification of the Middle East under a single imperial authority was substituted by a series of separate states with their own tariffs, custom regulations and currencies. Thus, the core of the Ottoman Empire was fragmented into nine countries: Egypt, Syria, Lebanon, Transjordan, Iraq, Palestine, Turkey, Saudi Arabia and Yemen. Only the latter three exercised full sovereignty during the interwar era; Britain retained control over Egypt, which was declared a protectorate; the other Arab nations became administered by Britain and France, in accordance with the so called Mandate system, established by the League of Nations.[1]

Previous studies indicate that the Middle East became progressively more integrated with the international economy during the so-called first wave of globalisation (1850-1914), thus following the same path of many other regions of the world (see Issawi, 1966; Owen, 1981; Islamogu-Inan, 1987; Kasaba, 1988; Pamuk, 1987, 2004; Panza, 2013; İnalcık and Quataert, 1996). At the time, global commodity and factor markets became increasingly more integrated, primarily thanks to the dramatic improvements in transportation technologies (Foreman-Peck, 1983; O'Rourke and Williamson, 1999).

This period was followed by an anti-global and autarkic interwar era, when commodity and factor price convergence came to a halt (Findlay and O'Rourke, 2007; Federico, 2012). One of the major causes for such reversal of globalisation has been identified with the failure to dismantle the system of protectionist trade policies put in place during the Great War.[2] In fact, in its aftermath, average tariff rates continued rising globally: from Europe to the US,[3] from Latin America to Asia.[4] Furthermore, other forms of protectionist policies, such as quantitative restrictions to trade, government sanctioned trade monopolies, import licences, antidumping legislations and competitive devaluations remained a widespread practice of interwar commercial policy internationally.[5]

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[1]The League of Nations sanctioned the division of the Ottoman Empire and granted Britain the right to administer Transjordan, Palestine and Iraq and France the right to administer Lebanon and Syria (Cleveland, 2004).

[2]Eichengreen (1992) provides an excellent overview of the global economy during the interwar period.

[3]In the mid-1920s tariffs were unambiguously higher than they had been in 1913 in Bulgaria, Czechoslovakia, Germany, Hungary, Italy, Romania, Spain, Switzerland and Yugoslavia (Findlay and O'Rourke, 2007). In the US, the Fordney-McCumber Tariff Act and later the Smoot-Hawley Tariff considerably raised overall protection (Irwin, 1998).

[4]In Latin America average tariffs increased from 22.2 percent in 1929 to 27.3 percent in 1932. In Asia tariffs were on the rise, too: between 1918 and 1929 average tariffs climbed from 4.8 percent to 14.7 percent in India and from 10.8 percent to 25.3 percent in Burma (Clemens and Williamson, 2004).

[5]During the interwar years, protectionism's damaging impact on trade volumes was not counteracted by a decline in the costs of international transport. While some productivity growth had been achieved in ocean shipping during the interwar, this was diminished by rising factor prices and thus was not sufficient to overcome the effect

Over the past decades, economic historians have made a concerted effort in examining the dramatic changes that affected national economies worldwide in the period between the First and Second World Wars, using different empirical approaches.[6] However, with a few exceptions, the countries of the Middle East have yet to be fully incorporated into this research agenda, with most of the existing economic history literature being predominantly of qualitative nature (see, e.g. Hershlag, 1964; Mead, 1967; Issawi, 1982; Tignor, 1989; Quataert, 1994; Owen and Pamuk, 1998).[7]

Focusing on the experiences of Egypt and Syria, this paper draws attention onto some fundamental questions in order to bring new insights on Middle Eastern economies during the interwar era: Did the region share the same anti-global developments of most countries? Did the Egyptian and Syrian commodity markets, once united by the same custom union under the aegis of the Ottoman Empire, disintegrate? Or did the incorporation of the two nations into the British and French colonial systems lead to the establishment of improved linkages within the region?

We address these issues through an empirical investigation of the process of market integration between Syrian and Egyptian commodities using Bayesian inference. We contribute to the scholarly debate on two fronts: first, we bring new insights to the growing market integration literature focusing not only on a period that has strikingly received very little attention, the interwar;[8] but also on a relatively unexplored region, the Middle East.[9] Secondly, the analytical tools chosen represent an important methodological contribution to the field, due to their specific suitability in dealing with small datasets within a dynamic multivariate framework. To this purpose, we explore the evolution of Middle Eastern market price relationships across time (1923-1939) and space (Aleppo, Beirut, Cairo and Alexandria), applying state of the art developments in Bayesian structural vector error correction models.

The paper is structured as follows: we first provide background information on historical trade patterns in the Middle East before the disruption of the Ottoman Empire (Section 2) and during the interwar (Section 3), drawing attention on the commercial ties between Syria and Egypt (Section 4). We then analyse the factors which impacted on the degree of market integration between the two countries, focusing on the role of trade policy, transport networks and commercial institutions (Section 5). After presenting our dataset (Section 6) and describing the methodology used in the empirical analysis (Section 7), we present and discuss the co-integration results (Section 8). Section 9 concludes.

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of widespread rampant protectionist measures (Mohammed and Williamson, 2004).

[6]Despite the existence of a voluminous literature on the interwar era, only very few works have explored empirically the disruption of commodity market integration and hence the deterioration of the process of international price transmission, which brought to a halt the globalizing trends of the previous four decades. An important exception is provided by Findlay and O'Rourke (2004:461-5), who measure commodity price differentials between different cities of the world to attest the disintegration of global markets.

[7]Notable exceptions are Hansen (1991); Pamuk and Williamson (2002); Yousef (2002); Karakoç (2012).

[8]Important contributions are Hynes et al. (2012); Trenkler and Wolf (2005); Federico and Persson (2007).

[9]We are not aware of the existence of any study on market integration in the Middle East in the interwar era.

## 2 Trade patterns in the Ottoman Empire before WWI

Before the dissolution of the Ottoman Empire all its regions belonged to the same custom union.[10] From the mid-19$^{th}$ century most Ottoman provinces experienced a spectacular increase in trade flows, spurred on by declining transport costs (Harlaftis and Kardasis, 2000).[11] The larger size of imports and exports is emblematic of the Empire's participation to the first wave of globalisation, which determined a partial shift in the patterns of exchange from within the region itself to trade with Europe. Such changes were particularly dramatic in Egypt, whose openness and integration with the world economy were the highest in the whole Ottoman realm.[12]

While trade with Europe grew, intra-Ottoman commerce continued to represent a larger portion of trade of most Middle Eastern states during the 19$^{th}$ and early 20$^{th}$ Century, despite the absence of well-developed infrastructure and transport system (İnalcık and Quataert, 1996, p.770).[13] For example, in 1862 the value of Ottoman imports in the province of Damascus was five times greater than that of non-Ottoman goods (İnalcık and Quataert, 1996, p.836).[14] Moreover, 80 per cent of all Damascus exports were directed to the Empire in 1892 (Peter, 2004, p.418). In 1910 about 45 per cent of Syria's [15] exports and 19 per cent of its imports went to and came from Egypt and other parts of the Empire (Musrey, 1969; Peter, 2004, p.418).[16] Trade between the various administrative divisions of Syria was substantial, too.[17]

Egyptian regional trade figures are less impressive, but still not negligible. Imports from other parts of the Ottoman Empire covered about one fifth and around 11 per cent of Egyptian average annual imports in 1884 and in 1909-1913 respectively. On the other hand, trade with Europe was much more conspicuous, with about two-thirds of Egypt's exports going to Britain and over one-third of its imports coming from there at the turn of the century (Musrey, 1969, p.200, footnote 9). Egypt's linkages with Great Britain were strengthened after colonization in 1882, when it de facto withdrew from the Ottoman custom union and signed a separate trade treaty with the Empire. This imposed a reciprocal 8 per cent *ad valorem* import tax [18]

[10] Import and export duties were fixed at 8 percent and 1 percent per respectively by a series of trade treaties signed in 1861/2. Some minor alterations occurred between 1870 and 1914 Tunisia became a French protectorate in 1884 and Libya an Italian colony in 1912.

[11] Trade rose from 9 million Turkish Lira in 1830 to 45.9 million in 1910-13 (Owen and Pamuk, 1998, p.4).

[12] See Panza (2013) for a study of market integration comparing the Egyptian and the western Anatolian cotton markets.

[13] Ottoman international exports formed around 25 per cent of Ottoman agricultural production, so that the remaining 75 per cent stayed within the Empire (İnalcık and Quataert, 1996, p.834).

[14] See İnalcık and Quataert (1996, pp.836-7) for a detailed account of intra-Ottoman trade flows.

[15] The region generally referred to as Syria before WWI (*Bilad-el-Sham*) included present day Syria, Lebanon, Jordan, Israel and Palestine.

[16] These exports included items such as barley, millet, sheep and other livestock, dried apricots, legumes, wine, etc., but also silk and cotton textiles.

[17] For example cereals were sent in large amounts from Homs and Hama to Aleppo, Tripoli and Beirut; from the Hauran to Damascus and Haifa; and from Gaza and Beersheba to Jerusalem and Jaffa. Fruits, vegetables, oil, soap and textiles and leather were also important trade items.

[18] Despite the Ottoman imperial tariff increase to 11 per cent in 1907, Egyptian goods continued to be subject to the lower 8 per cent rate.

The outbreak of the First World War, together with large numbers of casualties, forced migrations, famine and disease, led to a dramatic disruption of trade flows, compounded by the embargo of the Allies on the Mediterranean.[19] Its aftermath brought about the political and economic dismantlement of the Empire, marking the end of its large free trade area and the beginning of significant economic divisions within the Middle East.

# 3 Middle Eastern trade during the interwar era

The dissolution of the Empire gave origin to a set of countries with separate customs and different units of monetary systems. Palestine, Transjordan and Iraq became British mandates and had their currencies tied to the British pound. France obtained a mandate over Syria, a large region comprising the states of Syria, Greater Lebanon, Jabal al-Duruz, the Government of the Alawis (Latakia) and the Sandjak of Alexandretta. The official currency became the Syrian pound, tied to the French franc.[20]

Egypt, which became a British protectorate in 1914, was unilaterally declared independent by Britain in 1922. However, the economic and political ties between the two countries remained very strong during the whole interwar period: the British High Commissioner held powers with a strong potential for intervention in Egyptian economic matters since London reserved rights over four areas: defence, imperial communications, the Sudan and the protection of foreign interests (Tignor, 1989, pp.4 and 44).[21] Furthermore, the Egyptian pound remained pegged to the Sterling.

Despite the economic and political fragmentation, inter-Arab trade still constituted a substantial share of the total trade of most countries during the 1920s, aided by moderate tariff rates. Over one third of Syrian exports went to and around one tenth of its imports came from Egypt, Palestine, Transjordan, Iraq, the Hejaz and the Nejd, with Egypt and Palestine being its most important customers. Over two-fifths of Palestine's trade was conducted with Syria and Egypt, and most of Transjordan's exports went to Syria, Egypt and Palestine. Despite Egypt's lower level of engagement in the region (only about one-twentieth of its trade was carried on with other Arab countries), it imported a substantial quantity of goods from Palestine, Syria and Sudan and exported to those countries a considerable amount of cotton and other commodities (Musrey, 1969, p.15).[22]

Trade with the mandatory/occupying powers (France and Great Britain) became increasingly more important after the dissolution of the Empire, despite the initial absence of preferential commercial agreements. It was facilitated by tied currencies, foreign investments and foreign control. During the late 1920s over one third of Egypt's exports went to and around one fifth of

[19]In 1916, the volume of Ottoman foreign trade decreased to one fifth of the level before the war, 90 percent of which was undertaken with Germany and Austria (Pamuk, 2005, p.118).

[20]On April 1, 1920 the French High Commissioner emitted a decree for the establishment of a new Syrian paper currency based on the French franc. Thus, the Syrian pound, equivalent to 20 francs and divisible in 100 piasters, became the unit of currency, replacing the Turkish gold pound (Himadeh, 1936, p.264).

[21]In Egypt, a Department of Foreign Affairs was created in the Ministry of Interior to safeguard foreign interests, which benefitted from a series of tax exemptions allowed by the so called capitulations (Tignor, 1989, p.47). Moreover, British officials continued to play a fundamental role in the upper strata of the bureaucracy.

[22]Many of the export figures could be biased upwards since they include also re-exports. However, it is important to highlight that not all the trade conducted via land route was recorded and a great deal of smuggling took place within the region (Musrey, 1969, p.207).

its imports came from Great Britain. France was one of Syria's leading trade partners, accounting for about one-sixth of Syrian imports and exports.

However, despite the disruption of the Ottoman Empire and the establishment of closer linkages with Great Britain and France, there was still a semblance of a regional market in the Near East by 1930, which constituted an important outlet for foodstuffs and other agricultural commodities, as well as for a small number of manufactured goods produced in the region (Musrey, 1969, p.16).

It was during the 1930s that this market shrank, owing to a series of intertwined global and domestic factors, namely the depression, tariff escalation and monetary policy developments. The dramatic reduction in prices and output after the Great Depression led to an intensification of protectionist trade policies worldwide. Economic nationalism, which had not previously been a significant factor in inter-Arab trade relations, began to assert itself, mirroring a global trend. The division of the world in currency blocs (dollar, sterling, franc) had repercussions on the Middle East, weakening trade linkages among countries belonging to different blocs. The region was shaken by the same de-globalisation forces common the rest of the world, characterised by shrinking capital flows and declining commerce, both regionally and internationally (see Figure 1).

## 4 Egyptian and Syrian commercial ties in the interwar

The post-war commercial conventions between Syria and Egypt were based upon the latter's position as a non-member of the League of Nations: being non-contiguous to Syria, Egypt could not be granted preferential tariffs normally allowed for bordering countries (Burns, 1933, p.82). However, a provisional most-favoured-nation agreement was formally established (Accord of November 1, Burns, 1933, p.83), but it did not involve the granting of preferences.

In Syria the import duty applied to most goods remained the old Ottoman rate of 11 per cent *ad valorem* until 1923. Duties were raised progressively, ranging between 15 per cent (normal duty) and 30 per cent (maximum duty) in 1924 and between 25 per cent (normal duty) and 50 per cent (maximum duty) in 1926, mainly for the purpose of raising revenues.[23]

Until 1930, Egypt applied a uniform 8 per cent *ad valorem* tariff on most imports. Once tariff autonomy was gained, a set of protectionist measures were taken in order to encourage industry (Hansen, 1991, p.87). Tariffs on raw materials were lowered and those on manufactured goods raised (Hansen and Nashashibi, 1975, p.4).[24] A new general duty of 15 per cent was put in place together with specific duties applicable to a series of goods, reaching 25 per cent: duties rose particularly on fruit and vegetables, which represented most of Syrian exports to Egypt, and continued to grow over time.[25] With the devaluation of the Egyptian pound in 1931, due to British abandonment of the Gold standard, duties on many agricultural and industrial goods kept on rising and did so throughout the 1930s.[26] For example in 1931 the duties on bananas

[23]Normal duties were applied to the members of the League of Nations, while maximum duties to non-members.

[24]In addition to the tariff reform, the government supported manufacturing through subsidies, cheap loans and other industrial policy measures.

[25]Duties were raised also on household soap, artificial silk and other articles (Musrey, 1969, p.21).

[26]In addition to imposing high tariffs, the government of Egypt also extended funds to encourage the establishment of new industries and the modernization of old ones, and it purchased carpets, footwear, clothing and other articles from Egyptian producers only (Musrey, 1969, p.17). While such policies had a positive impact on

rose to 45 per cent and on oranges to 110 per cent (Burns, 1933, p.85). Consequently the Syro-Egyptian trade relationship deteriorated remarkably, with Syrian exports experiencing an 85 per cent decline between 1930 and 1933 (see Figure 2).

In Syria, too, import duties were raised substantially on many articles during the early 1930s. As retaliation to Egypt's trade policy, the tariff rate on rice, which constituted the main import from Egypt, increased, too. In 1930, Syrias ten main exports to Egypt were subjected to an average weighted duty of 21.1 per cent, whereas the ten main Syrian imports from Egypt bore an average weighted duty of 14.6 per cent (see Table 1).

The tariff war continued during the 1930s. In April 1933 Egypt imposed a surtax of 100 per cent on Syrian imports and on August 1933 Syria subjected Egyptian imports to its maximum duties, which were twice the normal rate. While a formal provisional most-favoured-nation agreement was signed in 1934, Egypt's tariff increase at the end of the 1930s and Syria's multiple devaluations did not aid a reinstatement of pre-depression trade relations. The Syrian pound (tied to the French franc) was devalued in 1936, and additional devaluations followed at intervals during 1937, so that it depreciated by around 50 per cent by the end of 1937. After the first devaluation in 1936, Syria raised most of its duties by 15 per cent, followed by an additional 20 per cent in May 1938.

While Syrian imports of Egyptian goods do not seem to have been affected extensively by tariff escalation, its exports to Egypt dropped significantly from 1930 (see Figure 2). For example, while during 1929 and 1930 around 17 per cent of Syrian exports went to Egypt, this figure dropped to around 5 per cent for the remainder of the 1930s. While during the 1920s Syria had exported to Egypt a substantial amount of different agricultural commodities, as well as various types of textiles, by 1939 Egypt imported only a limited range of Syrian goods (Musrey, 1969, p.22). The reason for such different repercussions of increased tariffs on bilateral trade relations is related to the timing of the onset of protectionism. The drop of Syrian exports to Egypt after the depression was so dramatic since until 1930 Egypt kept a comparatively low tariff rate (which did not hamper foreign trade). On the other hand, Syria embraced a protectionist stance much earlier, with its tariff rising progressively from 1923: in fact, it was in the 1920s that Egypt's exports declined the most (Figure 2).

The deterioration of the Egypt-Syria trade relation was paralleled by an analogous worsening of the whole Middle Eastern regional commerce. Similar policies of protectionism coupled with competitive devaluations were adopted by most countries of the region. By 1939 Egypt's inter-Arab trade, excluding Sudan, constituted less than 3 per cent of its total trade (Musrey, 1969, p.25). The only semblance of a regional market in the Near East during the 1930s was limited to Palestine, Transjordan and Syria. Within such a framework, almost no attempt was made to develop regional economic ties.[27]

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industrialization, they also had adverse effects on Egypts inter-Arab relations, especially with Palestine and Syria.

[27] The only exception was the maintenance of free trade relations between Egypt and Sudan and those between Syria and Palestine. The latter were terminated in 1939.

# 5 Factors influencing commodity market integration in the Middle East

To measure the degree of integration between Egyptian and Syrian markets, and its evolution over time, we explore the relationship between commodity prices. Specifically, this paper investigates the issue of market integration by looking at the mechanisms of price transmission between locations, as explained in detail in the methodology section. To this purpose, we draw from the classic paradigm of the Law of One Price (LOP), as specified by the standard spatial price determination model of Takayama and Judge (1972). The LOP postulates that markets are integrated when the price difference between two homogeneous goods sold in different locations equals to transfer costs. This implies that changes in demand or supply in one location are translated into the same degree of price change in each trading market. On the other hand, policies that impede the pass-through of price signals, such as government-induced distortions (tariffs, quotas, subsidies, etc.) or the presence of imperfectly competitive market structures, weaken the linkages between two trading economies, hindering, or *in extremis* preventing, markets from integrating.

Before making an empirical assessment on the degree of price transmission between Syrian and Egyptian markets, we are able to identify two opposing forces that influenced arbitrage opportunities between them. One the one hand, a series of factors may have acted against market integration: rising protectionism, particularly tariff escalation in the 1930s; the practice of competitive devaluations first of the Egyptian pound (1931) and successively of the Syrian pound (1936); Egypt/Britain and Syria/France preferential trade agreements; the rise of nationalism. All these forces caused an increase of the price differentials between trading markets, potentially leading to their disintegration. At the same time, other factors may have favoured integration: the relative low rates of protection in Egypt until the early 1930s; the expansion and improvements of physical infrastructure in both countries, which had a positive impact on transport costs; and the development of better commercial institutions which lowered transaction costs. Sections 5.1 to 5.3 will have a closer look at these factors.

## 5.1 Protectionism and exchange rate policy

The end of the Empire and the heavy mark caused by WWI, both in terms of physical destruction and of economic and monetary dislocation, left the Middle East facing severe inflationary pressures, like most belligerent countries. The whole region was affected by the intrinsic fragility of the international system made of reparations, war debts and lack of cooperation which crumbled with the Great Depression. The downward spiral of competitive devaluations, tight monetary policy and protectionism, which triggered an unprecedented fall in global trade, hit negatively Middle Eastern commodity markets, too. The world-wide spread of protectionist practices lowered international trade flows; widening price wedges between trading economies arising from declining export prices and increasing import prices caused a disintegration of the global market (Hynes et al., 2012, p.120). In Section 4 we have illustrated how Syria and Egypt's trade policies mirrored this global trend, being directed towards increasing the levels of protection of the respective domestic markets. These were reinforced by exchange rate policy choices. In fact, currency pegs to the franc and the sterling implied a renewed commitment to the gold standard. Recovery was particularly hard for the countries of the sterling bloc like Egypt, as the British pound fixed its value at the pre-war gold parity, despite the considerable change in financial strength and competitiveness. On the other hand, the French franc's devaluation at one fifth of its per-war parity gave Syria an initial competitive advantage over other countries in the region. While the adherence to the gold standard ensured some degree of exchange rate stability, this did not last for long. Changes in exchange rate policies in Britain and France had a direct impact on Syro-Egyptian commercial relationships: the devaluation of the pound (1931) and subsequently of the franc (1936) was likely to contribute to increased price fluctuations in both markets.

## 5.2 Transport

Both Syria and Egypt underwent a process of infrastructure development during the interwar, particularly the transport sector, which led to a reduction of transport costs both domestically and regionally. In Egypt, shipping costs dropped due to a series of improvements of the transport system (Issawi, 1963, p.32). The railway network increased[28] and cost of rail transport declined constantly, in response to an increase in motor competition.[29] A further downward pressure on railway rates was the result of a government policy aimed at encouraging both exports and local production.[30] Also international transport costs were reduced, through a series of subsidies granted to Egyptian shipping companies, which expanded their merchant fleet.

In Syria, the French administration embarked on an extensive program of transport development.[31] Road building was expanded through a systematic program, based on the construction of three longitudinal trunk lines, each traversing one of the plains running parallel to the coast.[32] This was complemented with the development of a series of transverse lines joining the plains and valley by connecting them across the mountain ranges, for example from Beirut to Damascus, from Tripoli to Homs, from Latakia to Aleppo (Himadeh 1936:179). Moreover, rail tracks expanded from 525 to 950 miles between 1914 and 1938 (Grunwald and Ronall, 1960, p.55) and rail rates experienced a sharp decline since 1928, due to increased competition from motor vehicles.[33]

Shipping facilities improved, too: in particular, the port of Beirut expanded, doubling in size, and was endowed with larger warehouses. Postal and telegraphic services experienced con-

[28] Railways track length increased from 1,900 miles in 1914 to 2,268 miles in 1939 (Hansen, 1991, p.43;Grunwald and Ronall, 1960, p.55).

[29] While water transportation along the Nile represented another source of competition, most commodities were transported by train due to the quicker delivery time and to the fact that river transport was not much available in Lower Egypt (Fahmy, 1931).

[30] Agricultural and industrial bulky products for local consumption and export were offered reduced rates (Issawi, 1963, p.202).

[31] French policymakers became particularly active from 1933/4 with the appointment of Count Damien de Martel as new high commissioner, who established a six-year plan to promote the development of roads, railways, ports and irrigation (Gates, 1998, p.31).

[32] The new road network connected Râs al-Nâkûra to Aleppo via Tyre, Didon, Beirut, Tripoli, Latakia and Antioch; Tyre to Aleppo via Zahlah, Baalbak, Homs and Hama; Dar'a to Homs via Damascus. For a description of the extension of the road network see Himadeh (1936, p.180).

[33] Himadeh (1936, pp.184-5) reports that from 1928 railway rates were modified from week to week to meet this competition and that the freight rate dropped from 5.62-8.10 Syrian piasters per ton in the late 1920s to 1-2 piasters per ton in the mid-1930s.

siderable progress, strengthening regional communications. Moreover, the first telephone lines were installed both within the Syrian territory and in connection with Palestine, Transjordan and Egypt.

The development of national transport systems and the reduction in domestic transport costs was likely to have reduced price differentials within both Syria and Egypt, thus favouring internal market integration. On the other hand, better shipping facilities and increased access to information may have improved the process of price transmission between the two countries.

### 5.3 Commercial institutions

The interwar period saw the consolidation and expansion of a series of institutions focusing on trade, particularly commercial banks. In Egypt, some of the gaps of the credit system were filled by the creation of specialised, government-sponsored banks which, among other things, facilitated trade transactions (Issawi, 1963, p.33). The foundation of the Egyptian Chamber of Commerce in Cairo was followed by the formation of other commercial banks in the 1920s easing both domestic and international trade. A particularly important role was played Bank Misr, the first purely Egyptian owned and managed institution, mirroring the rise of the Egyptian merchant and business community. The increase in the capital base of the bank facilitated the availability of credit for import-export activities (Tignor, 1989). Its special linkages with Syria consolidated trade relations between the two countries. Moreover, new multinational bank branches dealing with domestic and international trade were opened: British and French banks, already widespread before WWI, were joined by Italian and Belgian ones.

In Syria both foreign and domestic banks expanded the scope of their operation in the 1920s and 1930s, with commercial banking representing a major component of their activities. French banks opened new branches in different Syrian cities, all dealing with foreign trade: the Banque Française de Syrie, established in 1919; the Crédit Foncier d'Algérie et de la Tunisie opened its first branches in Beirut (1921) Aleppo (1930) Damascus and Tripoli (1931); the Compagnie Algérienne expanded to Beirut (1931) and Tripoli (1932); the Banco di Roma, confined to purely commercial banking, established three branches in Beirut, Aleppo and Damascus after WWI (Himadeh, 1936, pp.287-8).

Another chief banking establishment contributed to improve commercial operations, particularly between Syria and Egypt: the Banque Misr-Syrie-Liban. It was founded by the Bank Misr in collaboration with a group of Syrian financiers with the aim of improving trade and economic relations between the two countries (Himadeh, 1936, pp.290).

## 6 Data

One of the contribution of this paper stems from the creation of a new dataset for Syrian prices, using a completely unexplored primary source: the *Bulletin Économique Trimestriel des pays sous Mandat Français.* Egypt's prices are taken from another rarely utilised primary source, the *Annuaire Statistique.* We collected quarterly wholesale prices for seven commodities commonly used by Egyptian and Syrian consumers in Alexandria, Cairo, Beirut and Aleppo: wheat, olive oil, barley, flour, sugar, rice and soap. Our final data have been converted in £GB per kg and are presented in Figure 3. All Egyptian goods were expressed in Egyptian piasters (100 piasters equal to 1E£) and were converted in £GB using the following exchange rate: 1£E= 1.025 £GB (El Imam 1962). Prices were reported in the following units: wheat (type Zawati) in *ardeb* of

150kg; barley (type baladi Beheri) in *ardeb* of 120 kg; rice (type de Damiette, mahsous) in *kadah* of 1.835 kg; sugar (type granulé), olive oil (type de Candie) and flour in *oke* of 1.248 kg; soap (type baladi, Kafr el Zayat) in *rotl* of 0,449 kg.

Syrian goods were originally either in Turkish or Syrian piasters and converted in £GB using the quarterly exchange rates published in the various issues of the *Bulletin Économique* (1923-1939). The following units were used for prices in Beirut: *kantar* of 256 kg for wheat and barley; *rotol* of 2.564 kg for soap, flour and sugar; kg for rice. Prices in Aleppo were reported in kg or quintals.

# 7 Methodology

Using the framework offered by cointegration analysis, our study of market integration explores the evolution of the relationship between commodity prices in Syria and Egypt, both in the long and short-run. This approach is widely used in the literature as attested by the large body of theoretical and empirical studies, including those within the field of economic history.[34] In recent years the debate on how market integration should actually be measured has been quite lively, with scholars resorting to sophisticated econometric tools, mainly based on cointegration or price dispersion (see for example Uebele, 2011; Jacks, 2005; Federico and Persson, 2007; Sharp, 2008). This paper's empirical strategy contributes to this literature by drawing from the so-called *weak form* of the LOP, identifying the following relationship between prices:

$$y^A_{i.t} - y^B_{i.t} \leq c^{AB}_{i.t}, \tag{1}$$

where subscript $i$ stands for different commodities, $t$ denotes a time period, A and B refer to two locations and $c^{AB}_{i.t}$ to the cost of transferring $i$ from A to B. The condition of full price equality embodied by the so-called *strong form* of the LOP (i.e. $y^A_{i.t} - y^B_{i.t} = c^{AB}_{i.t}$) is unlikely to occur in the short-run, and even more so in our historical setting. Nevertheless, the relationship described in equation (1) constitutes an equilibrium condition, since spatial arbitrage will ensure that $y^A_{i.t} - y^B_{i.t}$ will move towards $c^{AB}_{i.t}$.

In our empirical analysis we will investigate the degree of price transmission between Syrian and Egyptian commodity markets focusing on two specific aspects of the LOP. Firstly, the fulfilment of equation (1), which represents the existence of price co-movement or completeness of price adjustments between locations. This is investigated through cointegration analysis. And secondly, the dynamics of the price relationships, represented by the speed of adjustment to the equilibrium defined in (1), after a price change in one location. This will highlight the degree of market efficiency (Federico, 2012) and will be explored via computing impulse response functions. A practical tool for performing these analyses is the structural vector error correction model (SVECM hereafter).

## 7.1 Testing market integration with SVECM

A SVECM is defined by:

$$A_0 \Delta y_t = \mu + \alpha \beta' y_{t-1} + B_1 \Delta y_{t-1} + \cdots + B_p \Delta y_{t-p} + u_t, \tag{2}$$

[34]See Federico (2012) for a review of the existing economic history literature on market integration.

for $t = 1, \dots, T$ where $\Delta y_t = y_t - y_{t-1}$ is a $N$-dimensional vector of the first differences of vector $y_t$; $A_0$ is a $N \times N$ matrix of contemporaneous effects; $\mu$ is a $N$-dimensional vector of constant terms; $\alpha$ and $\beta$ are $N \times r$ full rank matrices of so called loading coefficients and cointegrating vectors, respectively; $B_l$ for $l = 1 \dots, p$ are $N \times N$ matrices of the short-run dynamics; and $u_t$ is a $N$-dimensional vector of error terms that is assumed to follow a $N$-variate normal distribution with mean set to a vector of zeros, and an identity covariance matrix.

Several features of this model are especially useful in the analysis of the integration of commodities markets. First of all, SVECMs explicitly model the long-run equilibrium relationship between variables, represented by the product $\beta' y_{t-1}$. This term embodies the cointegration relationship and is stationary, despite containing a vector of unit-root nonstationary variables ($y_{t-1}$). $\beta' y_{t-1}$ constitutes a generalization of the left-hand side of equation (1), where the price difference between two locations is multiplied by the parameters from matrix $\beta$. The constant terms in the cointegrating relations, corresponding to trade costs $c_{i.t}^{AB}$, on the right-hand side of (1), are captured by vector $\mu$. Matrix $\beta$ is assumed to be semi-orthonormal, that is $\beta\beta' = I_r$, which assures the global identification of the system $\alpha\beta'$.[35] Thus, this formulation of the cointegration relationship encompass the LOP, as expressed in (1).

While the SVECM in equation (2) is commonly used in applied studies on cointegration, this has not yet been analysed in a Bayesian setting.[36] In this paper we utilise Bayesian inference, following the informative prior distribution proposed by Strachan and Inder (2004) for the cointegrating space spanned by matrix $\beta$. The estimation algorithm adapts the collapsed Gibbs sampling of Koop et al. (2008) to the structural form model, where the matrix of contemporaneous effects $A_0$ is estimated by the algorithm derived by Villani and Warne (2003) and Villani (2009). All the details regarding the specification of the model, its identification and the required restrictions, prior distributions, estimation algorithms and inference methods are explained in the Statistical Appendix.

The essential information to assess the presence of integration (or the absence thereof) in equation (2) is contained in the cointegration rank, $r$, which indicates the number of cointegrating relationships. In the context of our analysis, the number of cointegrating relations needs to be greater than one, but lower than $N$ in order to signal the existence of a long-run equilibrium between the markets in particular cities.[37] That case represents the hypothesis of market integration. If the number of cointegrating relations is equal to zero ($r$=0), then the commodities prices are unit-root non-stationary, but not cointegrated, thus pointing to the absence of long-run integration among markets. If the number of cointegrating relations is equal to the number of variables in the system *($r$=$N$)*, then the variables are unit-root stationary, and thus cointegration analysis is not a proper tool to investigate the integration of markets.

In order to test the cointegration rank for the commodities of our dataset we compute the so called Bayes factors, based on the Savage-Dickey Density Ratio (hereafter SDDR) proposed by Dickey (1971) and Verdinelli and Wasserman (1995).[38] The SDDR is the ratio of the posterior probability of the hypothesis that the cointegrating rank, $r$, is equal to zero to the posterior probability that the cointegrating rank is equal to some number $r_0$, where $1 \le r_0 \le N$. To make

[35]See Johansen (1988) for the solutions to a global and local identification of the VECM as well as Strachan and Inder (2004) for the analysis in the Bayesian context.

[36]We are not aware of any applied study using a Bayesian model in this form.

[37]$N$ represents the total number of locations analysed for a particular commodity.

[38]This Bayesian cointegrating rank test was proposed by Koop et al. (2008).

this interpretation feasible we assume that all the models have equal probability *a priori*. All the computational details are presented in the Statistical Appendix.

After having computed the cointegration rank for each commodity, we further investigate the cointegration relationships between cities. To this purpose we develop a new test to identify which variables enter the cointegration relationship. Specifically, we use the SDDR to test the hypothesis that some parameters of matrix $\beta$ are equal to zero. To be clear, we calculate the SDDRs setting zero restrictions on rows of the matrix of cointegrating vectors, accounting for all of the possible combinations. Finding evidence in favor of the zero restrictions implies that the corresponding variables do not enter the long-term equilibrium equation. This test allows us to identify which locations are integrated in each commodity market. We report all the computational details of this testing technique in the Statistical Appendix.

## 7.2 Impulse Response Functions

The impulse response function (hereafter IRF) analysis constitutes another useful tool to investigate the timing patterns of commodity market integration. The study of IRFs allows us to identify the impact that a one unit value price shock (£GB per kg) in one city has to all other variables in the system. This type of analysis provides information about the time patterns of the persistence of a price shock and its effect on the prices in other locations.

Specifically, we analyse the responses to orthogonal shocks, $u_t$, identified through the $(N-1)N/2$ just-identifying restrictions imposed on the elements of matrix $A_0$. This matrix of contemporaneous effects points to the degree to which a particular location reacts instantaneously to a price shock in the remaining locations. The choice of the exclusion restrictions in our models has been guided by the geographical structure, trade patterns and supply chains for each commodity. The zero restrictions imposed on the matrix of contemporaneous effects, $A_0$, is determined by the ordering of the variables in vector $\Delta y_t$. These considerations are presented for each particular commodity market in Section 8.

## 7.3 Cointegration analysis using Bayesian inference

Given our small dataset containing $T = 65$ observations, our objective is to perform a reliable investigation of commodity market integration. By employing the SVECM we perform a multivariate analysis where all variables are treated as endogenous. This choice, however, exposes statistical inference to small sample problems. To illustrate this point, consider a four-variate time series (a total number of 260 observations) and a SVECM with four lags ($p = 4$), and cointegrating rank $r = 1$. Such a model requires the estimation of 86 parameters based on just 3 observations per parameter. This illustrates how multivariate estimation with a small database is constrained by the limited number of observations to parameter, making asymptotic inference cumbersome (see e.g. Cheung and Lai, 1993; Nielsen, 2004).[39] Due to these computational difficulties we apply Bayesian inference methods to our modelling framework, in order to obtain feasible estimates while imposing as little arbitrary assumptions as possible.

The objective of Bayesian inference is to characterise the posterior distribution of the parameters of a model, collected in vector $\theta$, given observable data: $p(\theta|data)$. The posterior

[39]Multiple small-sample corrections for the Johansen's test for cointegration have been developed (e.g. Johansen, 2002).

distribution is proportional, up to a normalising constant, to the product of the likelihood function, $L(\theta; data)$, and a prior distribution of the parameters, $p(\theta)$. In our empirical setting the form of the likelihood function is determined by equation (2) and the conditional normality assumption of the error term, $u_t$. The prior distribution reflects the knowledge of the researcher about the parameters prior to seeing the data. The inference based on the posterior distribution is always valid given the model, even in small samples. In what follows, we explain how the choice of the identification of the cointegrating system, as well as other elements of our Bayesian model, provide a reliable cointegration analysis.

Consider the identifying assumption that $\beta$ is a semi-orthonormal matrix that spans the cointegrating space. As argued by Strachan and Inder (2004) and Villani (2005), Bayesian inference on such a VECM parametrisation allows a very flexible modeling of the cointegration space. We choose the prior distribution of the cointegrating space of our model following Strachan and Inder (2004) that allows for a proper posterior analysis. An alternative and commonly used identification constraint of the cointegrating system is to restrict the top $r \times r$ elements of matrix $\beta$ to an identity matrix. However, contrary to the solution we use, the constrained matrix $\beta$ does not allow for the full flexibility in spanning the cointegration space. Moreover, these restrictions lead to a nonstandard small-sample distribution of the maximum likelihood estimator (see e.g. Phillips, 1994) and to a posterior distribution being improper, multimodal or without specified moments depending on the particular specification of the model (see Strachan and Inder, 2004, and references therein).

Many different premises can determine the choice of a prior. In this study, it is motivated by the LOP, as explained in Section 8. In our setting the prior distribution covers the whole space between the cointegration plane (spanned by a matrix representing the LOP) and its orthogonal complement. The concentration of the prior probability mass in a space outstretched between these two cointegrating planes is determined by hyper-parameter $\tau$, for which we assume a hierarchical prior distribution and estimate it (see Statistical Appendix and Strachan and Inder, 2004; Koop et al., 2008, for the technical details). The estimation of $\tau$ makes the prior distribution interpretable, informative and flexible at the same time.

Another element of our Bayesian model which is particularly suitable for making inference in a small sample is the specification of the prior distribution for the VECM parameters $\mu$, $\alpha$ and $B_1, \ldots, B_p$. For these parameters we assume a multivariate normal distribution with a mean of zero vector and a diagonal covariance matrix. The diagonal elements of the covariance matrix are set to hyper-parameter $\nu^{-1}$ for which we specify a hierarchical prior and estimate it. Hyper-parameter $\nu$ shrinks the prior distribution towards the prior mean, thus increasing the precision of the estimation. Since the level of shrinkage is is not chosen arbitrarily but estimated, this technique is often referred to as adaptive shrinkage. In the econometric literature, the application of Bayesian shrinkage decided on a successful performance of large Bayesian vector autoregressions (see e.g. Bańbura et al., 2010; Koop, 2013) both in-sample and out-of-sample. Moreover, as shown by Giannone et al. (2014) although adaptive shrinkage introduces a bias in the estimates towards the prior mean, the implied IRFs are still capable of maintaining the correct shape and scaling. We apply these techniques to the SVECM analysis.

To summarise, our Bayesian model is designed to perform a reliable market integration analysis given the small dimensions of our data due to its ability to allow for: (1) flexibility in the modeling of all variables; (2) the unconstraining identification of the cointegrating system; and (3) the use of adaptive shrinkage techniques.

# 8 Empirical results

## 8.1 Details of the estimation procedure

In order to investigate the degree of integration between Egyptian and Syrian markets we estimate a series of Bayesian SVECMs with lag order from 1 to 4 and cointegrating rank from 1 to $N$, where $N$ is the number of variables available for each commodity. The detailed specification of the prior distributions is as follows. The Gamma prior distribution for the overall shrinkage parameter, $\nu$, has the mean parameter equal to $\underline{\mu}_{\nu} = 1$ and the degrees of freedom parameter set to $\underline{\nu}_{\nu} = 1 + Nr$. This choice makes this probability density function as disperse as possible, thus expressing our lack of knowledge *a priori* about the values of $\nu$. The specification of parameter $\underline{\nu}_{\nu}$ allows for proper computations of the SDDRs (see Koop et al., 2008). The Gamma prior distribution for the cointegration-specific shrinkage parameter $\tau$ has the mean parameter equal to $\underline{\mu}_{\nu} = 5$ and the degrees of freedom parameter set to $\underline{\nu}_{\nu} = 15$. With this choice we emphasise the role of the LOP in our investigation, giving it reasonable support *a priori* (see Koop et al., 2008).

In our prior specification the LOP is represented by choosing an appropriate setting for matrix $\tilde{H}$, for each particular model. $\tilde{H}$ spans the cointegrating space in the prior distribution of the matrix of cointegrating relationships $\beta$ (see more details in the Statistical Appendix). The specification of $\tilde{H}$ is guided by our knowledge of the nature of market ties in Syria and Egypt during the interwar and by the related assumptions on the locations where the LOP was most likely to hold.[40] Specifically, when setting the matrix, we take into account the close relation between the markets in Alexandria and Cairo (and Beirut and Aleppo) due to the small geographical distance between the cities, and to their belonging to the same nation state. Therefore, all our models (with the exception of soap, where data for Alexandria are not available) include the relationship: $y_{Cairo.t} - y_{Alexandria.t} = ecm_{1.t}$, in the specification of matrix $\tilde{H}$. This is done by setting the elements of one of its column to: 1 for $y_{Cairo.t}$; -1 for $y_{Alexandria.t}$; and 0 for others. Similarly, when data are available for both Syrian cities (barley and wheat), we emphasise the potential existence of internal market integration. This is done by specifying in a column of matrix $\tilde{H}$ the relation: $y_{Beirut.t} - y_{Aleppo.t} = ecm_{2.t}$, for models with $r \geq 2$. For models with higher cointegrating ranks, an additional column represents the link between the ports of Alexandria and Beirut (due to the availability of data we use the Aleppo series for olive oil and the Cairo series for soap). The detailed specification of matrices $\tilde{H}$ for all the models is available on request.

The value of the hyper-parameter $\tau$ establishes whether the probability mass *a priori* is closer to the cointegrating plane defined by matrix $\tilde{H}$ (as specified above), or to a plane that is orthogonal to $\tilde{H}$. For all the estimated models, the posterior means of this parameter range between 0.24 and 0.32, with the standard deviations not exceeding the value of 0.14. This result confirms our intuition about the degree of integration between cities, since our $\tilde{H}$ receives a higher weight than the matrix orthogonal to it (this is the case for $\tau < 1$, see Koop et al., 2008).

We performed the posterior simulations by launching the estimation algorithms at the following starting values: we set all the parameters of the SVECM to zeros beside the diagonal elements of matrix $A_0$ which we set to 1. The missing observations (see Figure 3) were linearly

[40]The empirical work of Yousef (2002) points to the existence of domestic integration within various regions of Egypt during the interwar. We follow his findings for the ordering of our variables in the specification of matrix $\tilde{H}$.

interpolated with Stata function ipolate and treated as given. After 10.000 preliminary draws, the algorithm converged to a stationary posterior distribution for of all the models. We then simulated a sample of 50.000 more draws from the posterior distribution and used it to make inference. The computer codes for the reproduction of all the results are available on request.

## 8.2 Market integration testing

Table 2 reports the results of the Bayesian cointegrating rank testing for the SVECMs with lag order $p = 4$. The table includes the logarithms of the Bayes factors, which is equal to the logarithms of the ratio of the posterior probabilities of two models: $\ln(\Pr[M_r|data]/\Pr[M_0|data])$, where $M_i$ denotes the SVECM with the cointegration rank $i$. The Bayes factors are computed on the basis of the SDDR. A much easier tool for interpreting these results is provided by the posterior probabilities of the ranks calculated on the basis of the Bayes factors, reported in the bottom panel of Table 2. These probabilities provide clear evidence that a cointegrating rank equal to zero is preferred for all commodities. The no cointegration hypothesis ($r = 0$) gains the largest probability *a posteriori*, thus suggesting that the series were likely not to be cointegrated. Moreover, our results show that the hypothesis that all variables are jointly unit-root stationary within a particular systems ($r = N$) gains no support at all. We also performed the same testing with other lags ($p = 1, 2, 3$); the results are consistent with our estimations, except for sugar.[41]

While $r = 0$ has the largest posterior probability, Table 2 also points to the existence of non negligible probabilities of having a cointegration rank $\geq 1$ and $< N$ for most commodities: barley, flour, rice, sugar and wheat. In these cases we cannot rule out *tout court* the possibility of integrated markets. This is true especially for barley and sugar, whose posterior probabilities of $r > 0$ and $r < N$ are 0.31 and 0.58, respectively. These results make it compelling to explore further the possibility of commodity market integration. Which locations were the drivers of our results? In which cities was the process of price pass-through stronger?

To answer these questions, and to better understand the dynamics of cointegration, we identify the variables belonging to the cointegrating relations for the SVECM with $r = 1$, using the test described in section 7.1. For barley, we find that the posterior probability that Aleppo and Beirut do not enter the cointegrating relation jointly is over 4.5 times greater than the posterior probability of such an event for Alexandria and Cairo. A similar ratio is found for wheat, thus pointing to the existence of larger degree of integration within Egypt rather than within Syria or between the two countries. A similar conclusion can be drawn for flour and rice. For these two commodities, the probability that Beirut is excluded from the cointegrating relation is nearly 2 times higher (equal to 1.86 and 1.97 respectively) than the probability of excluding either Alexandria or Cairo. These results also hold for models with one lag.

To summarise, our empirical findings provided strong evidence for the lack of cross-border market integration between Egypt and Syria. While we could not exclude the possibility of cointegration for some commodities, we identified stronger linkages within Egyptian cities. The following section will contribute to the interpretation of these outcomes.

[41] Our results indicate that in the case of sugar the lower the lag order the greater the support for cointegration (for $p = 1$ the probabilities for the cointegrating rank being equal to 0, 1 or 2 are 0.07, 0.65 and 0.27 respectively). We, however, use the results for $p = 4$ due to the potential overestimation of $r$ due to too low a $p$ (compare with e.g. Boswijk and Franses, 1992).

### 8.3 Timing patterns in price shocks transitions

While the posterior probabilities do not provide support for the existence of long-run linkages between Egyptian and Syrian prices, we investigate further the timing patterns of the price shocks transitions, in order to consolidate our findings. For this purpose we plot the IRFs for each commodity, which allows us to gain a more thorough understanding of the mechanism of price pass-through between locations. The identification strategy needed for the computation of the IRFs requires imposing exclusion restrictions on matrix $A_0$, chosen according to bilateral trade patterns and to the nature of supply linkages between Syrian and Egyptian cities.

First of all, it is important to highlight that trade between Syria and Egypt took place mainly via steam and sail ships through the ports of Beirut and Alexandria.[42] Data on trade patterns and composition are only sporadically available for each year and for each commodity type. Table 3 provides some evidence on the direction and magnitude of bilateral trade for some of the goods included in our study. While most of these commodities were generally both imported and exported bilaterally, Syria was predominantly a supplier of wheat, barley, flour and sugar to Egypt and Egypt an exporter of rice to Syria. Furthermore, based on historical records, we treat Syria as a supplier of olive oil and soap to Egypt. We used these criteria to determine the identifying restrictions presented in Table 4. For instance, in the case of barley and wheat, the $A_0$ matrix restrictions assume that: 1) the prices in Beirut were likely to affect Egyptian markets, whereas Egyptian cities did not impact the prices in Beirut within one quarter; 2) the prices in Aleppo had contemporaneous effects on the prices in Beirut and Alexandria; 3) the prices in Alexandria affected contemporaneously the prices in Cairo, but not conversely.

Figures 6 to 12 present the IRF for each commodity and each city. Some important common patterns emerge from the IRFs of barley, flour, rice, soap, sugar and wheat: first of all, we can observe that the responses to own price shocks die out quickly, as the IRFs are significant only for 1 or 2 quarters. Thus, they highlight the relatively transitory nature of shocks. Secondly, the impact of orthogonal price shocks between locations is significant within Egyptian markets, with exogenous shocks to Alexandria affecting Cairo's barley, flour, sugar and wheat prices. The contemporaneous effects of shocks from other locations are only a half (for barley and wheat) or a third (for flour and sugar) of the magnitude of own shocks. All these effects fade away at most 3 quarters after the occurrence of the shock. Other contemporaneous effects for these commodities, with the exception of Aleppo's shock on Beirut's prices for wheat, are economically or statistically negligible. We also note that the price shocks for olive oil are more persistent. The effects of own shocks die out after over a year in Egyptian cities and are quite persistent (for at least 3 years) in Aleppo. As expected, Aleppo played a leading role in the regional olive oil market, transmitting the impact of a price shock to the Egyptian markets, in the short-run. Although these IRFs exhibit more persistent effects, they are by nature short-lived, and thus not visible in the long-run relationship between markets. Thus, the results provided by the IRFs are consistent with our previous findings on the integration of Egyptian prices: we can in fact observe that in general price shocks were not completely transmitted and occurred only within Egypt.

To summarise, consistently with the testing results provided by the Bayes factors, the IRF analysis confirmed that the strongest market linkages occurred within Egypt only. Thus, the lack of long-lasting price transmission between Syrian and Egyptian cities is indicative of the

[42]Specific data on the number of ships cleared in Beirut and Alexandria are provided by the *Bulletin Économique* and the *Annuaire Statistique* respectively.

absence of regional market integration. We can infer that the previously documented reduction in domestic and regional transport and telecommunication costs and the consolidation of bilateral ties through commercial banking were not enough to counter the negative impact of the protectionist practices embraced by both countries. These findings highlight the deterioration of economic relations between the two nations, well supported by available historical evidence from contemporary manuscripts and newspapers emphasising the negative impact of tariff escalation on Syro-Egyptian commercial ties (see e.g. De Monicault, 1936, p.93).

Another key component of this paper's results, emphasised by the IRFs, is the contrast between the presence of internal integration in Egypt and the lack thereof in Syria. The different geographical characteristics of the two countries are likely to have played an important role in defining the dynamics of price pass-through. Egypt's Nile valley, where both Alexandria and Cairo are located, was a very compact and well connected area since the 19$^{th}$ century. Hence, the improvement of transportation networks and the related reduction in trade costs described in section 5.2 are likely to have facilitated the process of price integration.[43] On the other hand, Syria's geography continued to represent an obstacle for the integration of local markets, despite the improvement of the road network connecting Beirut to inland Aleppo (via Latakia) and the creation of a rail connection between the two cities (via Hama and Homs) in 1933 (De Monicault, 1936, pp.80-81). Syria's uneven physical topography is likely to have contributed to maintaining segmented markets, allowing wholesale merchants to exploit arbitrage opportunities, resulting in the absence of price transmission domestically.

## 8.4 Taking Turkey into account

As a robustness check, we also collected price data for the same commodities in Istanbul, the capital of the Ottoman empire, which historically had strong trade linkages with both Egypt and Syria. These data are available only from 1926, thus resulting in the loss of 12 observations from our original sample.[44] The addition of Turkish prices maintained the baseline results of the cointegration rank testing unaltered. The hypothesis of no cointegration gained the dominant posterior probability mass for all commodities. In particular, the posterior probability supporting $r = 0$ was larger than that of our main analysis, with the exception of barley and wheat. The posterior probabilities of $r$ being 0, 1 or 2 are 0.54, 0.28 and 0.16, respectively, for barley, and 0.75, 0.16 and 0.08, respectively, for wheat. For these two commodities, the results of the testing for zero restrictions on the cointegrating relationship for models with $r = 1$ show similar results to those of our baseline sample. They have in fact signalled the existence of internal integration within Egyptian markets. For both commodities, placing zero restrictions for Aleppo, Beirut and Istanbul leads to a higher posterior probability than those for Alexandria and Cairo.[45] These results confirm that the possible evidence of cointegration is likely to hold only within Egypt. For the remaining commodities (flour, olive oil, rice, soap and sugar) our findings do not support the hypothesis of cross border integration between Egyptian, Syrian and Turkish markets, at least, in our small dataset.

[43] Our findings are in line with those of Yousef (2002) who tests for the presence of price transmission for a series of commodities in various Egyptian towns and concludes that these were integrated.

[44] Since having a shorter time series further exposes multivariate cointegration to small sample problems, we decided not to use the extended database as our baseline analysis.

[45] In particular, the zero restrictions imposed on Aleppo and Beirut jointly have a posterior probabilities 4.6 times larger for barley and nearly 4.4 times larger for wheat than the joint zero restrictions for Alexandria and Cairo.

The analysis of the IRFs for the data including Istanbul does not reveal any significant change from our baseline estimations. The dynamic transitions between cities are to a certain extent lower in magnitude, and most cross-city effects are not statistically significant. Both features could be attributed to the smaller sample size and the larger number of parameters estimated in the models. All the quantitative results for this section are available on request.

# 9 Conclusion

This paper has explored the issue of integration between various important Syrian and Egyptian commodity markets after the disruption of the Ottoman Empire. Using multivariate time series analysis based on structural VECMs, we computed Bayesian factors to test for cointegration and to be able to select the appropriate model for each commodity. After identifying the locations entering a cointegration relationship, the analysis was matched with the computation of impulse response functions for each market, in order to measure the impact of orthogonal shocks in one city onto the others. This methodology allowed us to investigate the dynamics of the relationship among prices in Cairo, Alexandria, Aleppo and Beirut and their evolution over time.

After the incorporation of Syria and Egypt into the French and British spheres of influence, trade linkages between the two countries deteriorated considerably, particularly after the Great Depression, mirroring a global trend of advancing economic nationalism. This generated an escalation of protectionist policies, which, coupled with the use of competitive devaluations, further damaged the economic relationship between the two countries. While at the same time both Egypt and Syria saw an improvement of their infrastructure and commercial institutions, our analysis indicates that these developments were not enough to outweigh the impact of protectionist trade policy. Our empirical results suggest that the induced increase in price wedges between the two economies inhibited regional price transmission, thus leading to the absence of cross-border price pass-through. While cointegration testing did not exclude the possibility of market integration, a deeper investigation of the data through additional testing and IRF analysis, showed that the transmission of price signals was strongest between the Egyptian cities of Cairo and Alexandria. These results were robust to the inclusion of Istanbul prices into all models.

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# A Statistical Appendix: Bayesian estimation and inference for SVECM

This appendix contains all the derivations required to perform the computations used in this paper. First, we present all the assumptions regarding the identification of the model; the model is rewritten in two different matrix notations, useful for posterior derivations. Then, we provide details on the prior specification, and finally, we discuss the algorithm used for the simulation of the posterior distribution of the parameters given the data. The estimation algorithm adapts the collapsed Gibbs sampling of Koop et al. (2008) (when the panel data dimension is neglected) to the structural model, where the matrix of contemporaneous effects $A_0$ is estimated by the algorithm derived by Villani and Warne (2003) and Villani (2009). In the last part, the details of computations of the Savage-Dickey density ratio for the cointegration rank testing are explained.

## A.1 Model and likelihood function.

The analyzed model is the structural vector error correction model:

$$A_0 \Delta y_t = \mu + \alpha\beta' y_{t-1} + A_1 \Delta y_{t-1} + \cdots + A_p \Delta y_{t-p} + \epsilon_t, \tag{3}$$

$$\epsilon_t \sim i.i.d.\mathcal{N}\left(\mathbf{0}_N, I_N\right), \tag{4}$$

for $t = 1, \ldots, T$, where $A_i$, for $i = 0, \ldots, p$ are $N \times N$ matrices of contemporaneous effects (for $i = 0$) and autoregressive terms (for $i = 1, \ldots, p$), $\alpha$ and $\beta$ have dimensions $N \times r$ and are the so called loading matrix and the matrix of cointegrating relationships, respectively. The error term, $\epsilon_t$, follows a multivariate standard normal distribution, as in equation (4), where $\mathbf{0}_N$ denotes a $N \times 1$ vector of zeros, and $I_N$ is an identity matrix of order $N$.

The model as stated in equations (3) and (4) is not identified. In order to uniquely identify the parameters of the model the identifying restrictions need to be imposed on the matrix of contemporaneous effects, $A_0$, and on the cointegrating system consisting of matrices $\alpha$ and $\beta$. The identifying restrictions for $A_0$ are problem-specific and here we only report in this appendix how they can be imposed for the purpose of the estimation. In this paper, we only consider exclusion restrictions. Let $1 \times n$ vector $A_{0.n}$ denote $n$th row of matrix $A_0$. Then, it can be expressed in terms of its unrestricted elements collected in a $1 \times q_n$ vector $a_n$:

$$A_n = a_n U_n, \tag{5}$$

where $U_n$ is an appropriate $q_n \times N$ matrix of zeros and ones.

The problem of identification of the cointegrating system for Bayesian inference was considered by Strachan (2003), Strachan and Inder (2004) and Villani (2005). The problem consist of the lack of uniquely identified matrices $\alpha$ and $\beta$ in a matrix product $\alpha\beta'$. It can be presented by stating the following equalties:

$$\alpha\beta' = (\alpha\kappa)(\beta\kappa^{-1})' \equiv \boldsymbol{\alpha}\boldsymbol{\beta}', \tag{6}$$

that hold for any nonsingular matrix $\kappa$. We specify this matrix following Koop et al. (2010) as:

$$\kappa = \left(\alpha\alpha'\right)^{1/2} = \left(\boldsymbol{\beta}\boldsymbol{\beta}'\right)^{1/2}, \tag{7}$$

which defines a convenient reparametrization in equation (6). The cointegrating system as defined by equations (6) and (7) is then identified by assuming that matrices $\beta$ and $\boldsymbol{\alpha}$ are semi-orthogonal (see Strachan and Inder, 2004):

$$\beta'\beta = I_r, \quad \text{and} \quad \boldsymbol{\alpha}'\boldsymbol{\alpha} = I_r. \tag{8}$$

We describe two different matrix notations for our model that later are used for describing the estimation algorithm. The first proposed notation takes the form of a likelihood function that is convenient to derive the steps of the Gibbs sampler for parameters $A_0$, $\mu$, $A_i$ for $i = 1, \dots, p$. Let a $N \times T$ matrix stack vectors $\Delta y_t$ for $t = 1, \dots, T$: $\Delta y = \begin{bmatrix} \Delta y_1 & \dots & \Delta y_T \end{bmatrix}$. Define a $k \times 1$ vector, where $k = r + 1 + Np$, such that: $x_t = \begin{pmatrix} y_{t-1}'\beta & 1 & \Delta y_{t-1}' & \dots & \Delta y_{t-1}' \end{pmatrix}'$, and a $k \times T$ matrix that collects $x_t$: $X = \begin{bmatrix} x_1 & \dots & x_T \end{bmatrix}$. Accordingly, define a $N \times T$ matrix collecting error terms: $E = \begin{bmatrix} \epsilon_1 & \dots & \epsilon_T \end{bmatrix}$, and a $N \times k$ coefficients matrix $B = \begin{bmatrix} \alpha & \mu & A_1 & \dots & A_p \end{bmatrix}$. Then the model from equations (3) and (4) can be written as:

$$A_0 \Delta y = BX + E, \tag{9}$$
$$\text{vec}(E) \sim \mathcal{N}(\mathbf{0}_{NT}, I_{NT}), \tag{10}$$

or it can be written in terms of its $n$th equation as:

$$a_n U_n \Delta y = B_n X + E_n, \tag{11}$$

where similarly as $a_n U_n$ is a $n$th row of matrix $A_0$, $B_n$ and $E_n$ are the $n$th rows of matrices $B$ and $E$ respectively. Then the likelihood function is:

$$L(A_0, B, \beta; \mathbf{y}) \propto |\det(A_0)|^T \exp\left\{-\frac{1}{2}\sum_{n=1}^{N}(a_n U_n \Delta y - B_n X)(a_n U_n \Delta y - B_n X)'\right\}. \tag{12}$$

The second notation is used for a convenient representation of the sampling algorithm for the parameters of the cointegrating system. Define a $\tilde{k}\times 1$ vector $\Delta\tilde{y}_t = \begin{pmatrix} \Delta y_t' & -1 & -\Delta y_{t-1}' & \dots & -\Delta y_{t-p}' \end{pmatrix}'$, where $\tilde{k} = Np + N + 1$, and a $T \times \tilde{k}$ matrix that stacks vectors $\Delta\tilde{y}_t$: $\Delta\tilde{y} = \begin{bmatrix} \Delta\tilde{y}_1 & \dots & \Delta\tilde{y}_T \end{bmatrix}'$. Further, define a $T \times N$ matrix $y_{-1} = \begin{bmatrix} y_0 & \dots & y_{T-1} \end{bmatrix}'$ and a $\tilde{k} \times N$ matrix of coefficients $C = \begin{bmatrix} A_0 & \mu & A_1 & \dots A_p \end{bmatrix}'$. By defining $\hat{y} = \text{vec}(\Delta\tilde{y}C)$, $\hat{x} = \boldsymbol{\alpha} \otimes y_{-1}$, $\boldsymbol{b} = \text{vec}(\boldsymbol{\beta})$, and $\hat{\varepsilon} = \text{vec}(E)$ the model from equation (3) can be written as:

$$\hat{y} = \hat{x}\boldsymbol{b} + \hat{\varepsilon}. \tag{13}$$

The likelihood function becomes then:

$$L(\boldsymbol{\alpha}, \boldsymbol{b}, C; \mathbf{y}) \propto |\det(A_0)|^T \exp\left\{-\frac{1}{2}\left[\boldsymbol{b} - \hat{b}\right]' V_{\boldsymbol{b}}^{-1}\left[\boldsymbol{b} - \hat{b}\right]\right\}, \tag{14}$$

where $V_{\boldsymbol{b}} = \left(\hat{x}'\hat{x}\right)^{-1} = \left(\boldsymbol{\alpha}'\boldsymbol{\alpha}\right)^{-1} \otimes \left(y_{-1}'y_{-1}\right)^{-1}$, and $\hat{b} = V_{\boldsymbol{b}}\hat{x}'\hat{y}$.

## A.2 Prior distribution

Let a $N \times r$ matrix $\tilde{H}$ state prior cointegrating relations. Then $H = \tilde{H}\left(\tilde{H}'\tilde{H}\right)^{-1/2}$ is a $N \times r$ semi-orthogonal matrix that spans the prior cointegrating space. Let a $N \times r$ matrix $Z$ be such that $Z_{ij} \sim i.i.d.\mathcal{N}\left(0, \frac{1}{\nu}\right)$ for $i = 1, \dots, N$ and $j = 1, \dots, r$. Following Strachan and Inder (2004), write $\boldsymbol{\beta} = HH'Z + \sqrt{\tau}H_\perp H_\perp' Z = P_{\sqrt{\tau}}Z$, where $P_{\sqrt{\tau}} = HH' + \sqrt{\tau}H_\perp H_\perp'$, and $H_\perp$ is an orthogonal compliment of $H$. Then, each column of $\boldsymbol{b}$, denoted by $\boldsymbol{b}_j$ for $j = 1, \ldots, r$, is conditionally normally distributed given hyper-parameters $\tau$ and $\nu$, with mean $\mathbf{0}_N$ and covariance matrix equal to $\frac{1}{\nu}P_\tau$. $\tau$ is a cointegrating space specific shrinkage hyper-parameter, whereas $\nu$ is an overall-shrinkage hyper-parameters (see Koop et al., 2008, for their detailed interpretation).

Each row of matrix $\alpha$, denoted by $\alpha_n$ for $n = 1, \ldots, N$, is *a priori* conditionally normally distributed given $\beta$, $\nu$ and $\tau$ with mean $\mathbf{0}_r$ and covariance matrix $\frac{1}{\nu}\left(\beta' P_\tau^{-1}\beta\right)^{-1}$. A joint prior distribution is assumed for the rows of matrix $B$, denoted by $B_n$ for $n = 1, \ldots, N$. The conditional prior distribution of $B_n$ given $\beta$, $\nu$ and $\tau$ is a multivariate normal distribution with mean $\mathbf{0}_k$ and covariance matrix $\nu^{-1}\underline{V}_{B_n}$. Note that vector $B_n$ includes also $n$th row of matrix $\alpha$. Therefore, matrix $\underline{V}_{B_n}$ is a diagonal matrix with its diagonal elements equal to one, except for the top-left-hand-side $r \times r$ block that is equal to $\left(\beta' P_\tau^{-1}\beta\right)^{-1}$.

A hierarchical structure is assumed for the parameters of the model. Therefore, prior distributions are assumed also for the hyper-parameters $\nu$ and $\tau$, and these parameters are also estimated. The overall shrinkage parameter $\nu$ has a gamma prior distribution with the mean parameter set to $\underline{\mu}_\nu$ and the degrees of freedom parameter set to $\underline{\nu}_\nu - Nr$, where the correction by term $Nr$ is used to make the comparison of models of different cointergration rank $r$ correct (see Koop et al., 2008). The inverse of the cointegration related hyper-parameter $\tau$ follows *a priori* gamma distribution with the mean parameter set to $\underline{\mu}_\tau$ and the degrees of freedom parameter set to $\underline{\nu}_\tau$. Finally, the unrestricted elements of the rows of the contemporaneous effects matrix, denoted by $a_n$, follow a flat improper prior distribution as long as the resulting matrix $A_0$ is invertible.

To summarize, the prior specification for the considered model has a detailed form of:

$$p(\theta) = p(\nu)p(\tau^{-1})p(a_n)\prod_{j=1}^{r} p(\boldsymbol{b}_j|\tau, \nu)\prod_{n=1}^{N} p(B_n|\tau, \nu, \beta)p(\alpha_n|\tau, \nu, \beta)p(a_n), \tag{15}$$

where $\boldsymbol{b} = \text{vec}(\boldsymbol{\beta})$, $\beta = \boldsymbol{\beta}(\boldsymbol{\beta}'\boldsymbol{\beta})^{1/2}$ and each of the prior distributions is assumed:

$$\begin{aligned}
\boldsymbol{b}_j|\tau, \nu &\sim \mathcal{N}\left(\mathbf{0}_N, \frac{1}{\nu}P_\tau\right),\\
\alpha_n|\tau, \beta, \nu &\sim \mathcal{N}\left(\mathbf{0}_r, \frac{1}{\nu}\left(\beta' P_\tau^{-1}\beta\right)^{-1}\right),\\
B_n|\nu, \tau, \beta &\sim \mathcal{N}\left(\mathbf{0}_k, \nu^{-1}\underline{V}_{B_n}\right),\\
\nu &\sim \mathbf{G}\left(\underline{\mu}_\nu, \underline{\nu}_\nu - Nr\right),\\
\tau^{-1} &\sim \mathbf{G}\left(\underline{\mu}_\tau, \underline{\nu}_\tau\right),\\
p(a_n) &\propto \mathbf{1}(\det(A_0) \neq 0),
\end{aligned}$$

for $n = 1, \ldots, N$ and $j = 1, \ldots, r$, where $\mathbf{1}$ denotes an indicator function that takes the value of 1 when the condition in brackets hold and 0 otherwise.

## A.3 Gibbs sampling algorithm

Gibbs sampling algorithm is a tool to draw a sample from a joint posterior distribution of the parameters given data by an iterative sampling scheme from the full conditional posterior distributions (see Casella and George, 1992, and references therein). In what follows, we characterize the full conditional posterior distributions for parameters $a_n$, $B_n$, $\beta$, $\alpha$, $\tau$ and $\nu$ consecutively.

**Sampling matrix of contemporaneous effects.** The full conditional posterior distribution of matrix $A_0$ is proportional to:

$$A_0|B,\beta,\mathbf{y} \propto |\det(A_0)|^T \exp\left\{-\frac{1}{2}\sum_{n=1}^{N}(a_n-\overline{a}_n)\,\overline{V}_{a_n}^{-1}\,(a_n-\overline{a}_n)'\right\}, \tag{16}$$

where:

$$\overline{V}_{a_n} = \left[U_n\Delta y\Delta y' U_n'\right]^{-1},$$
$$\overline{a}_n = B_n X\Delta y' U_n'\overline{V}_{a_n}.$$

The full conditional posterior distributions of $a_n$ are for $n = 1,\dots,N$ in a form as presented by Villani and Warne (2003) and Villani (2009) who derive an appropriate sampling algorithm. The algorithm draws independently from the full conditional distribution of the unrestricted parameters of each row of $A_0$, $a_n$.

**Sampling autoregressive parameters.** Combining the likelihood function (12) with the prior distribution results in the full conditional posterior distribution for $B_n$ being a multivariate normal distribution. The draws are obtained by independent sampling from the following distribution for $n = 1,\dots,N$:

$$B_n'|a_n,\beta,\nu,\tau,\mathbf{y} \sim \mathcal{N}\left(\overline{B}_n',\overline{V}_{B_n}\right),$$

where:

$$\overline{V}_{B_n} = \left[XX' + \nu\underline{V}_{B_n}^{-1}\right]^{-1},$$
$$\overline{B_n} = a_n U_n\Delta y X'\overline{V}_{B_n}.$$

**Sampling cointegration parameters.** Given a draw for matrix $\alpha$ from the previous step, compute $\boldsymbol{\alpha} = \alpha(\alpha'\alpha)^{-1/2}$. Then, parameter $\boldsymbol{b}$ is drawn from a multivariate normal distribution:

$$\boldsymbol{b}|C,\boldsymbol{\alpha},\nu,\tau,\mathbf{y} \sim \mathcal{N}\left(\overline{\boldsymbol{b}},\overline{V}_{\boldsymbol{b}}\right),$$

where:

$$\overline{V}_{\boldsymbol{b}} = \left[\hat{x}'\hat{x} + I_r\otimes\nu P_\tau^{-1}\right]^{-1},$$
$$\overline{\boldsymbol{b}} = \overline{V}_{\boldsymbol{b}}\hat{x}'\hat{y},$$

where $\otimes$ denotes the Kronecker product. Given the draw for $\boldsymbol{b}$, compute $\beta = \boldsymbol{\beta}(\boldsymbol{\beta}'\boldsymbol{\beta})^{1/2}$ and $\alpha = \boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\beta})^{1/2}$, i.e. draws of matrices $\alpha$ and $\beta$ obtained by the collapsed sampling algorithm (see Koop et al., 2008, and references therein for more detailed explanations about the collapsed sampling algorithm).

**Sampling shrinkage parameters.** The overall shrinkage parameter is drawn from the gamma distribution:

$$\nu|B, \tau, \mathbf{y} \sim \mathbf{G}\left(\overline{\mu}_\nu, \overline{\nu}_\nu\right),$$

where:

$$\begin{aligned}
\overline{\nu}_\nu &= \underline{\nu}_\nu + Nk, \\
\overline{\mu}_\nu &= \overline{\nu}_\nu \left[\frac{\underline{\nu}_\nu - Nr}{\underline{\mu}_\nu} + \sum_{n=1}^{N} B_n \underline{V}_{B_n}^{-1} B_n'\right]^{-1}.
\end{aligned}$$

Similarly, the inverse of the cointegration-specific shrinkage parameter is drawn from the gamma distribution:

$$\tau^{-1}|\boldsymbol{\beta}, \nu, \mathbf{y} \sim \mathbf{G}\left(\overline{\mu}_\tau, \overline{\nu}_\tau\right),$$

where:

$$\begin{aligned}
\overline{\nu}_\tau &= \underline{\nu}_\tau + Nr - r^2, \\
\overline{\mu}_\tau &= \overline{\nu}_\tau \left[\frac{\underline{\nu}_\tau}{\underline{\mu}_\tau} + \nu \operatorname{tr}\left(\boldsymbol{\beta}' H_\perp H_\perp' \boldsymbol{\beta}\right)\right]^{-1},
\end{aligned}$$

where tr denotes a trace of a matrix.

## A.4 Testing cointegrating rank

In order to test for the cointegrating rank of the model we perform a Savage-Dickey density ratio (SDDR) test (see Dickey, 1971; Verdinelli and Wasserman, 1995). For a chosen lag order of the SVECM model we estimate models with different cointegrating ranks $r = 1, \ldots, N$. Then, for each such model, we compute a SDDR to compare the posterior probabilities (we assume equal prior probability for each cointegrating rank, $r = 0, 1, \ldots, N$, equal to $(N+1)^{-1}$) of a model with cointegrating rank $r$ and a model with the cointegrating rank equal to zero. The SDDR is defined for a hypotheses given by $\alpha = \mathbf{0}_{N\times r}$, where $\mathbf{0}_{N\times r}$ is a $N \times r$ matrix of zeros, as follows:

$$SDDR = \frac{p\left(\alpha = \mathbf{0}_{N\times r}|\mathbf{y}\right)}{p\left(\alpha = \mathbf{0}_{N\times r}\right)}, \tag{17}$$

where the numerator of equation (17) includes an ordinate of the marginal posterior distribution of $\alpha$ evaluated at $\mathbf{0}_{N\times r}$, whereas the denominator includes an ordinate of the marginal prior distribution of that parameter evaluated under the null hypothesis.

The marginal posterior distribution $p(\alpha|\mathbf{y})$ is computed from the full conditional posterior distribution of parameter $B_n$ by its marginalization (in order to obtain the full conditional posterior distribution of $\alpha_n$ for $n = 1, \ldots, N$) and by applying the Rao-Blackwell tool (see Gelfand and Smith, 1990). Given a sample of $S$ draws from the posterior distribution of parameters given data, $\left\{a_n^{(s)}, \beta^{(s)}, \nu^{(s)}, \tau^{(s)}\right\}_{s=1}^{S}$, the exact computations are performed according to:

$$\hat{p}\left(\alpha = \mathbf{0}_{N\times r}|\mathbf{y}\right) = S^{-1} \sum_{s=1}^{S} \prod_{n=1}^{N} \hat{\phi}\left(\alpha_n = \mathbf{0}_r | a_n^{(s)}, \beta^{(s)}, \nu^{(s)}, \tau^{(s)}, \mathbf{y}\right), \tag{18}$$

where $\hat{\phi}$ denotes an ordinate of a pdf of a $r$-dimensional normal distribution with the mean set to the first $r$ elements of $\overline{B_n}$ and the covariance matrix set to the top-left-hand-side $r \times r$ block of $\overline{V}_{B_n}$ (as specified in Section A.3) evaluated at $\mathbf{0}_r$.

The marginal prior distribution $p(\alpha)$ is computed from the marginal prior distribution of $\alpha_n$ given $\beta$, $\nu$ and $\tau$ by applying the Rao-Blackwell tool. Given a sample of $Q$ draws from the prior distributions of parameters, $\left\{\beta^{(q)}, \nu^{(q)}, \tau^{(q)}\right\}_{q=1}^{Q}$, the exact computations are performed according to:

$$\hat{p}(\alpha = \mathbf{0}_{N\times r}) = Q^{-1} \sum_{q=1}^{Q} \prod_{n=1}^{N} \hat{\phi}\left(\alpha_n = \mathbf{0}_r | \beta^{(q)}, \nu^{(q)}, \tau^{(q)}\right), \tag{19}$$

where $\hat{\phi}$ denotes an ordinate of a pdf of a $r$-dimensional normal distribution with the mean set to a vector of $r$ zeros and the covariance matrix equal to $\nu^{-1}\left(\beta' P_\tau^{-1} \beta\right)^{-1}$ (as defined in Section A.2) evaluated at $\mathbf{0}_r$. Note that $\hat{p}(\alpha = \mathbf{0}_{N\times r})$ for $r = 1, \dots, N-1$ could be computed by the analytical formula derived by Koop et al. (2008).

In order to estimate a model with cointegrating rank $r = N$ the algorithm presented in Section A.3 requires adjustments. Note that cointegrating rank $N$ means that all the variables are unit-root stationary, and thus that each of the variables represent a cointegrating relation. Therefore, $\beta$ is set to an identity matrix of order $N$. Consequently, the steps for simulating parameters $\beta$ and $\tau$ are dropped from the Gibbs sampling algorithm. Other steps require some straightforward adjustments as well.

## A.5 Testing zero restrictions on rows of matrix $\beta$

In order to test zero restrictions for the rows of matrix $\beta$ we test the restrictions imposed on matrix $\boldsymbol{b}$. For that purpose, we compute a SDDR. Let $\mathcal{K}$ be a set of $K$ indicators, $k$, that correspond to the elements of matrix $\boldsymbol{b}$ to be restricted, such that the whole chosen rows of corresponding matrix $\beta$ are restricted. Of course, $0 \leq K \leq Nr$. Thus, vector $\boldsymbol{b}_{k\in\mathcal{K}}$ denotes the restricted parameters. Then the SDDR is defined by:

$$SDDR = \frac{p(\boldsymbol{b}_{k\in\mathcal{K}} = \mathbf{0}_K | \mathbf{y})}{p(\boldsymbol{b}_{k\in\mathcal{K}} = \mathbf{0}_K)}. \tag{20}$$

The marginal posterior distribution $p(\boldsymbol{b}_{k\in\mathcal{K}} = \mathbf{0}_K | \mathbf{y})$ is computed using the Rao-Blackwell tool:

$$\hat{p}(\boldsymbol{b}_{k\in\mathcal{K}} = \mathbf{0}_K | \mathbf{y}) = S^{-1} \sum_{s=1}^{S} \hat{\phi}\left(\boldsymbol{b}_{k\in\mathcal{K}} = \mathbf{0}_K | C^{(s)}, \boldsymbol{\alpha}^{(s)}, \nu^{(s)}, \tau^{(s)}, \mathbf{y}\right), \tag{21}$$

where $\hat{\phi}$ is an ordinate of a pdf of a $K$-variate normal distribution with the mean equal to $\overline{\boldsymbol{b}}_{k\in\mathcal{K}}$ and the covariance matrix set to $\overline{V}_{\boldsymbol{b}.k\in\mathcal{K}}$. The ordinate of the marginal prior distribution is computed by:

$$\hat{p}(\boldsymbol{b}_{k\in\mathcal{K}} = \mathbf{0}_K) = Q^{-1} \sum_{q=1}^{Q} \hat{\phi}\left(\boldsymbol{b}_{k\in\mathcal{K}} = \mathbf{0}_K | \nu^{(q)}, \tau^{(q)}\right), \tag{22}$$

where $\hat{\phi}$ is an ordinate of a pdf of a $K$-variate normal distribution with the mean set to a $K$-dimensional vector of zeros and the covariance matrix equal to $\left[I_r \otimes \nu^{-1} P_\tau\right]_{k\in\mathcal{K}:k\in\mathcal{K}}$.

# B Tables

Table 1: The burden of the Syrian and the Egyptian tariff, 1930

| **Syrian exports to Egypt** | | | **Syrian imports from Egypt** | | |
|---|---|---|---|---|---|
| Commodity | Value SYR £ | Egyptian tariff [%] | Commodity | Value SYR £ | Syrian tariff [%] |
| Ovine animals | 811,997 | 7 | Rice | 1,238,325 | 15 |
| Butter | 402,096 | 12.4 | Asphalt | 160,176 | 11 |
| Fruit paste | 369,828 | 23.7 | Raw hides | 64,131 | exempt |
| Olive oil | 166,657 | 18.8 | Box cartons | 58,465 | 10 |
| Dried legumes | 148,309 | 62 | Leaf tobacco | 39,666 | 31 |
| Cotton cloth | 124,204 | 16 | Sole leather | 27,742 | 15 |
| Oranges | 103,265 | 65.1 | Cotton cloth | 26,516 | 20 |
| Wheat | 98,281 | 14 | Cigarette paper | 24,138 | 35 |
| Dried apricots | 92,216 | 12 | Jute sacks | 22,479 | exempt |
| grapes | 72,306 | 9.4 | Beer | 21,375 | 25 |

*Source*: Burns (1933).

Table 2: Cointegration rank testing

| $r$ | barley | flour | olive oil | rice | soap | sugar | wheat |
|---|---|---|---|---|---|---|---|
| | Bayes factors: $\ln B_{r.0}$ | | | | | | |
| 1 | -1.066 | -1.838 | -7.786 | -2.600 | -3.247 | -0.093 | -1.836 |
| 2 | -2.268 | -2.812 | -9.258 | -3.994 | -6.307 | -0.810 | -3.054 |
| 3 | -5.288 | -8.905 | -11.528 | -10.216 | - | -8.239 | -5.993 |
| 4 | -13.601 | - | - | - | - | - | -14.779 |
| | Posterior probabilities: $\Pr[M_r\|data]$ | | | | | | |
| 0 | 0.688 | 0.820 | 0.9995 | 0.915 | 0.961 | 0.424 | 0.827 |
| 1 | 0.237 | 0.130 | 0.0004 | 0.068 | 0.037 | 0.387 | 0.132 |
| 2 | 0.071 | 0.049 | 0.0001 | 0.017 | 0.002 | 0.189 | 0.039 |
| 3 | 0.003 | 0.000 | 0.0000 | 0.000 | - | 0.000 | 0.002 |
| 4 | 0.000 | - | - | - | - | - | 0.000 |

*Note*: The top panel of the table presents the logarithm of Bayes factors. This is equal to the logarithm of the ratio of the posterior probabilities of two models: $\ln(\Pr[M_r|data]/\Pr[M_0|data])$, where $M_i$ is the SVEC model with four lags ($p = 4$) and the cointegration rank $i$. The bottom panel presents the posterior probabilities of models with cointegrating rank $r$ implied by the Bayes factors. These probabilities are computed by: $\Pr[M_r|data] = B_{r.0}\left(\sum_{i=0}^{N} B_{i.0}\right)^{-1}$.

Table 3: Average monthly bilateral trade between Egypt and Syria, selected commodities, 1923, 1924, 1930 [kg]

| **Commodity** | **Trade direction** | **1923** | **1924** | **1930** |
|---|---|---|---|---|
| Barley | from Syria to Egypt | 98,023 | 54,250 | |
| Oil | from Syria to Egypt | 4,512 | 3,812 | 7,686 |
| Wheat | from Syria to Egypt | 28,550 | 81,616 | 198,307 |
| Soap | from Syria to Egypt | 6,278 | | |
| Rice | from Egypt to Syria | 67,500 | 67,500 | 96,717 |

*Source*: Bulletin Économique(1923, 1924, 1930); Zilkha (1937)

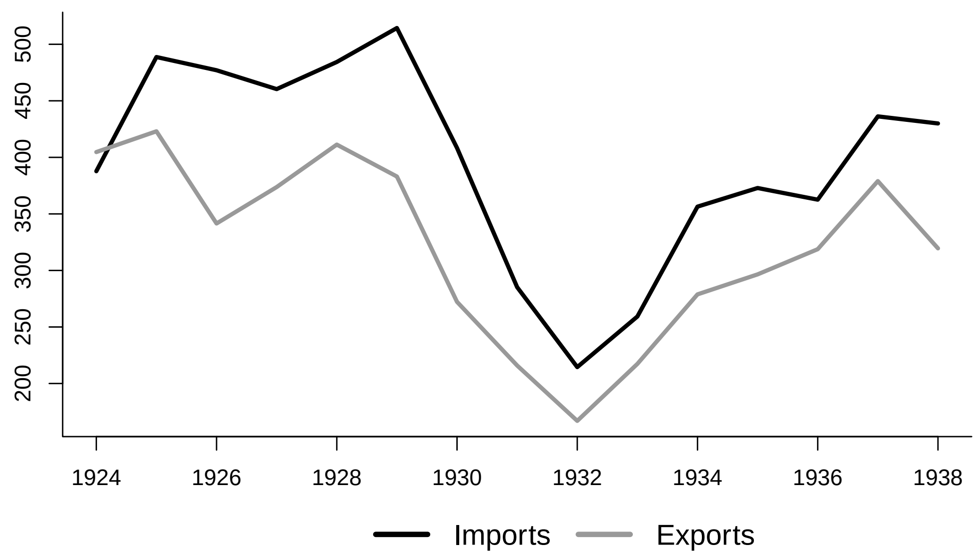
Table 4: Restrictions imposed on $A_0$ matrices.

| Barley | | | | Flour | | | | Olive oil | | | | Rice | | | | $\Delta y_t$ |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| * | * | * | 0 | * | * | * | - | * | * | - | * | * | 0 | 0 | - | $\Delta y_{Cairo.t}$ |
| 0 | * | * | * | 0 | * | * | - | 0 | * | - | * | * | * | 0 | - | $\Delta y_{Alexandria.t}$ |
| 0 | 0 | * | * | 0 | 0 | * | - | 0 | 0 | - | * | * | * | * | - | $\Delta y_{Beirut.t}$ |
| 0 | 0 | 0 | * | - | - | - | - | - | - | - | - | - | - | - | - | $\Delta y_{Aleppo.t}$ |
| **Soap** | | | | **Sugar** | | | | **Wheat** | | | | | | | | $\Delta y_t$ |
| * | - | * | - | * | * | * | - | * | * | * | 0 | | | | | $\Delta y_{Cairo.t}$ |
| - | - | - | - | 0 | * | * | - | 0 | * | * | * | | | | | $\Delta y_{Alexandria.t}$ |
| 0 | - | * | - | 0 | 0 | * | - | 0 | 0 | * | * | | | | | $\Delta y_{Beirut.t}$ |
| - | - | - | - | - | - | - | - | 0 | 0 | 0 | * | | | | | $\Delta y_{Aleppo.t}$ |

*Legend*: * – free parameter to be estimated, 0 – exclusion restriction, - – the effect is not included in the system

## C Figures

Figure 1: Middle Eastern total imports and exports in millions US $, 1924-38.

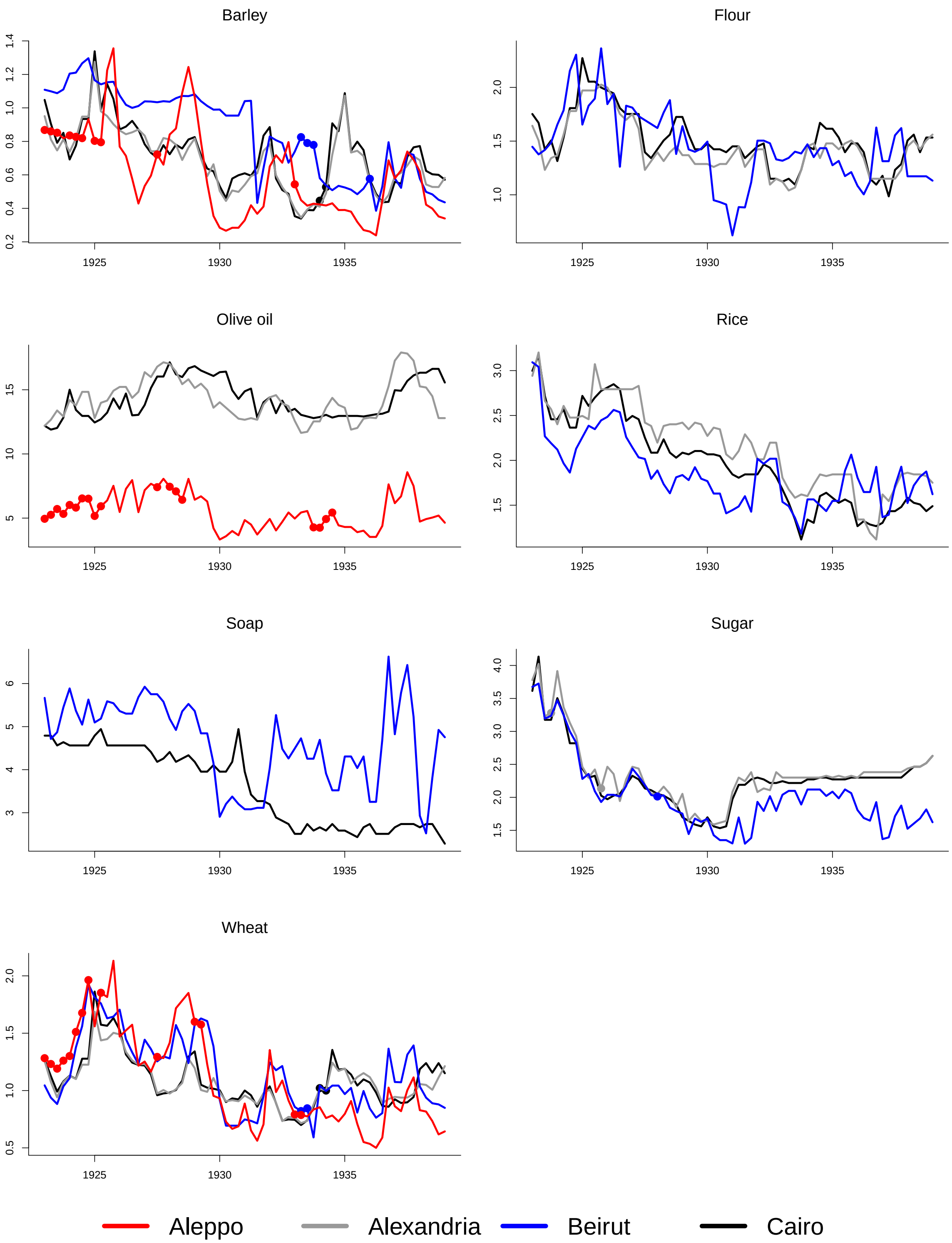



*Notes*: Includes trade of Syria, Egypt, Turkey, Iraq and Palestine.
*Sources*: Iraq: League of Nations (1924-41). Egypt: Annual bulletin of foreign trade (1924-41). Jewish Agency for Palestine (1945). Turkey: Başbakanlik Devlet İstatistik Göstergeler (1992). Syria: Saade (1942); Méouchy et al., (2004). Palestine: (Himadeh, 1939, Anglo-American Committee of Inquiry on Jewish Problems in Palestine and Europe (1946)).

Figure 2: Trade between Syria and Egypt, 1921-1941 in millions of Syrian £.

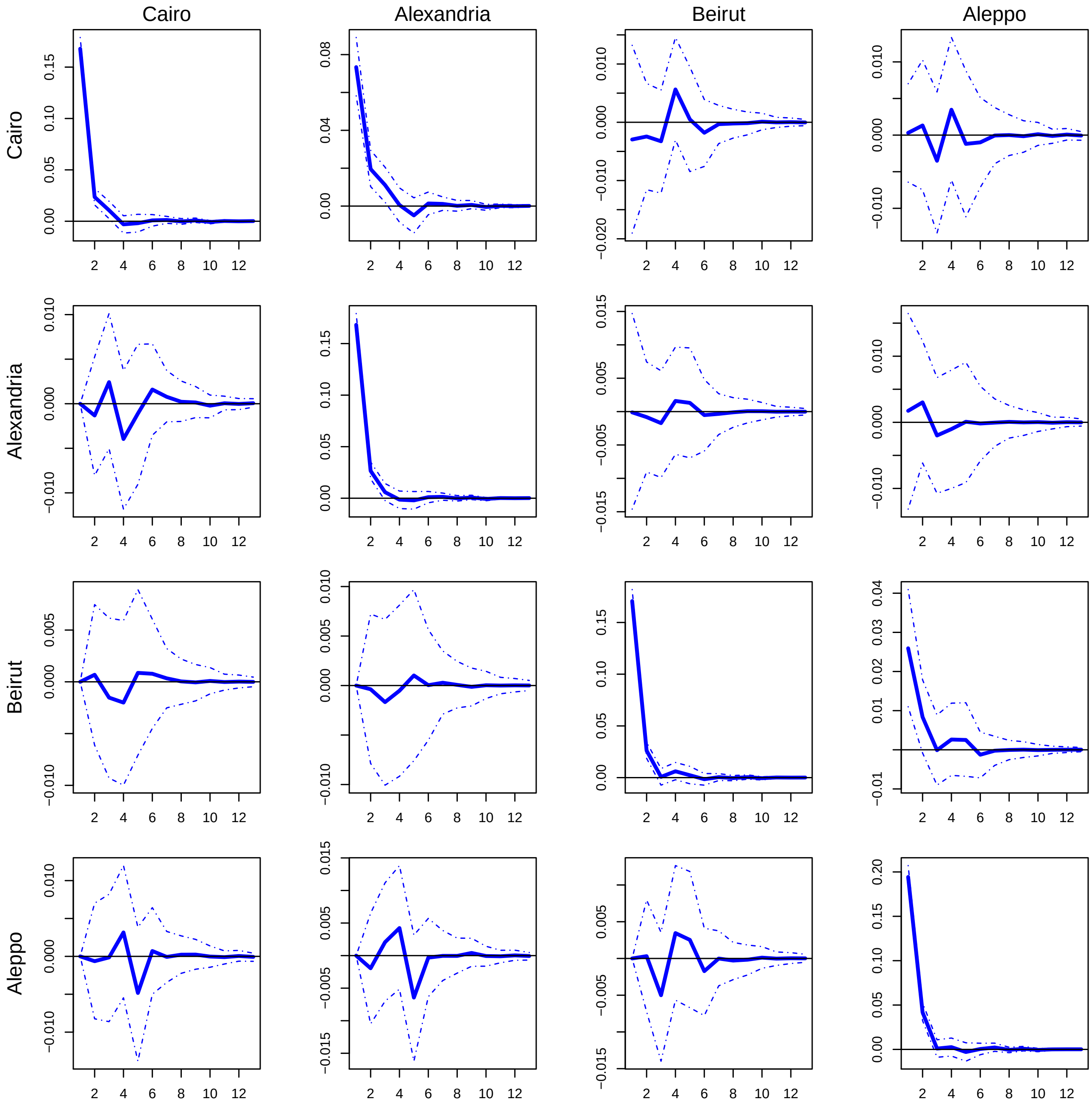

*Sources*: 1921–4: Haut Commissariat de la République Française en Syrie et au Liban (1927); 1925–33: Himadeh (1936); 1934: Bulletin Économique Trimestriel (1934); 1935–1941: Saade (1942).

Figure 3: Prices of commodities in Egypt and Syria, 1923-1939 [GB£/kg].



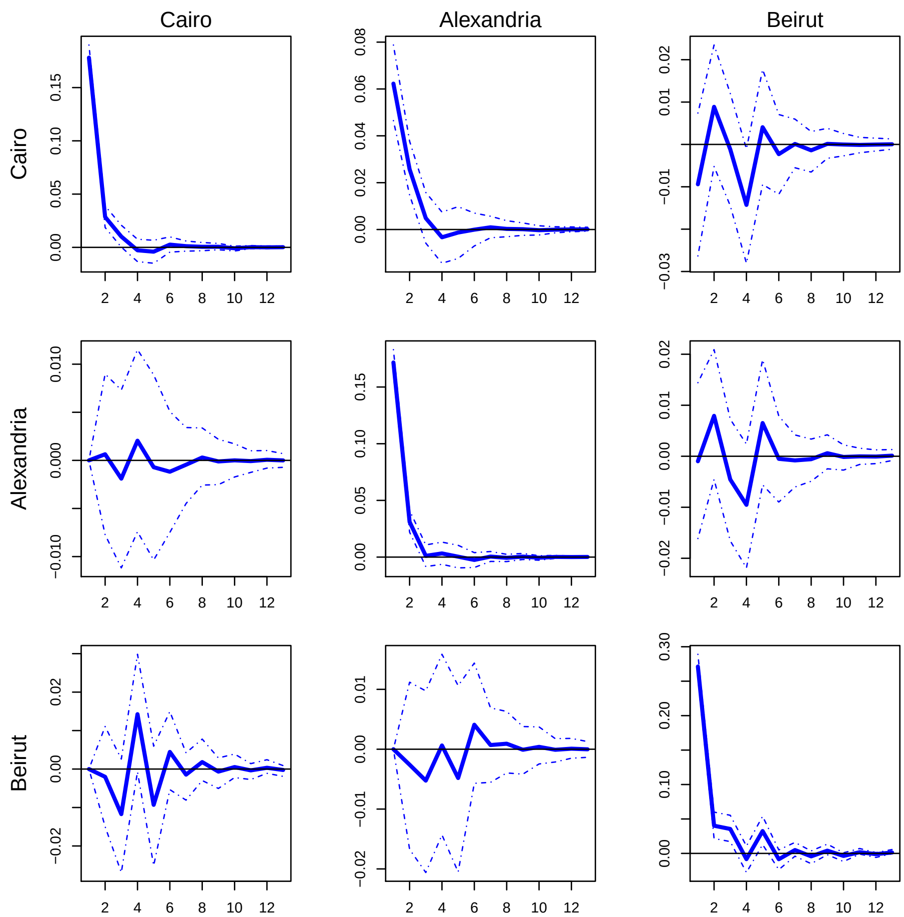

*Sources*: See data section. Quarterly time series converted to GB£per kilogram. The points denote interpolated missing observations.

Figure 4: Barley: impulse response functions analysis.



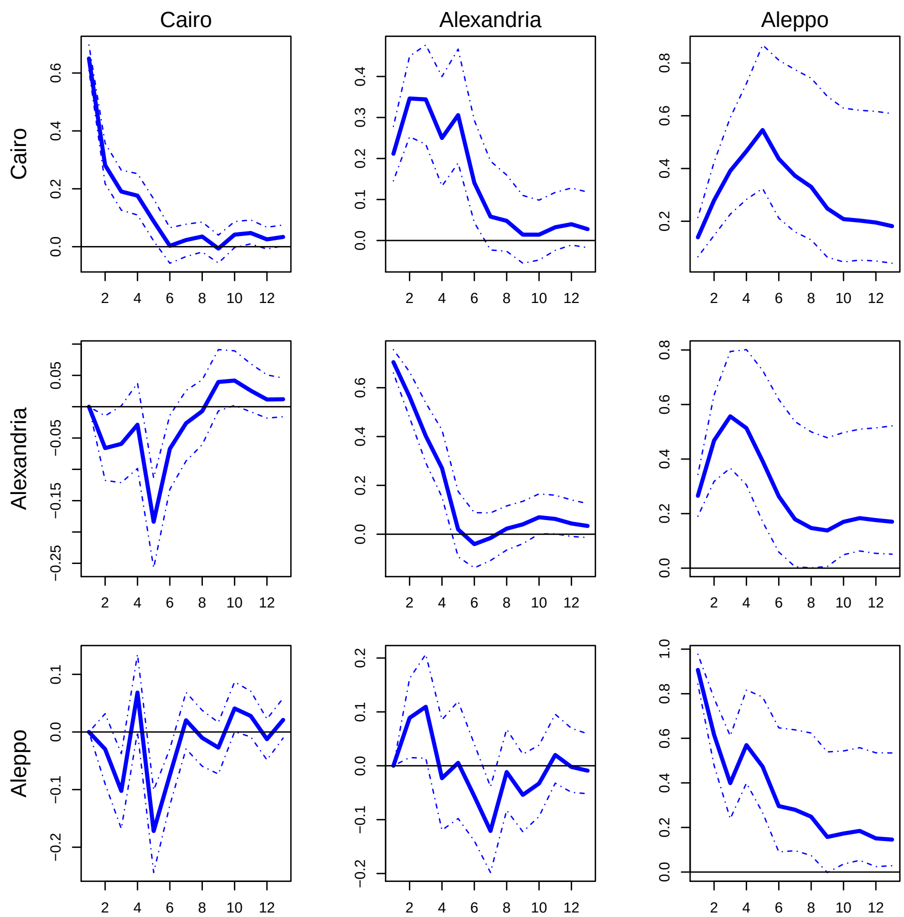

*Note:* The graphs present the impulse response functions (IRF) of a price shock of a magnitude of one standard deviation from the city listed at the top of the columns to the prices in the cities listed on the left hand side of each row. The dotted lines represent the bounds of 50% highest posterior probability intervals for the IRFs at each horizon. The horizontal axes represent time horizon after the occurrence of the price shocks.

Figure 5: Flour: impulse response functions analysis.



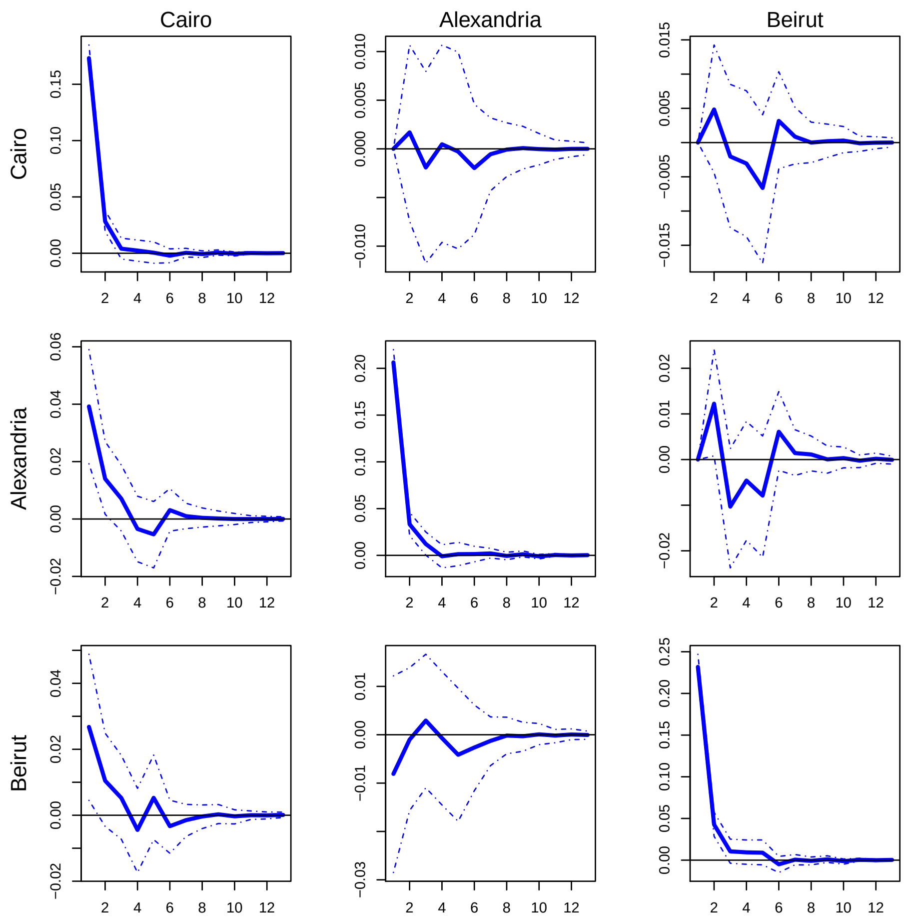

See the note to Figure 4.

Figure 6: Olive oil: impulse response functions analysis.



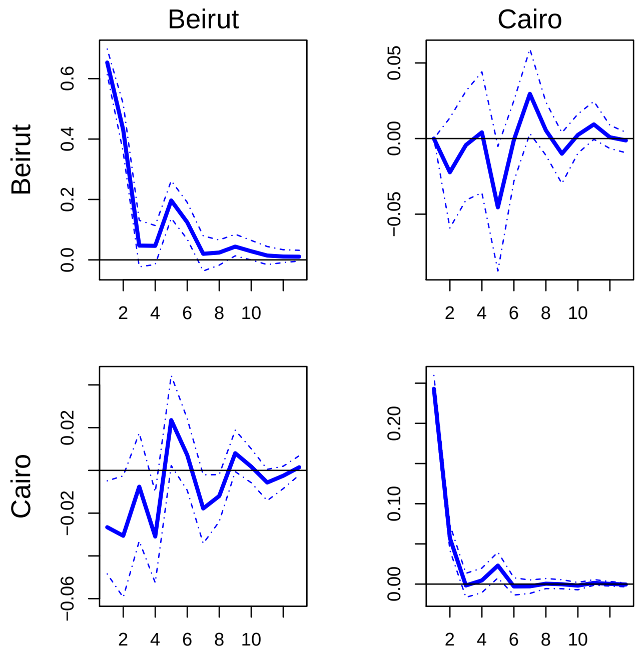

See the note to Figure 4.

Figure 7: Rice: impulse response functions analysis.



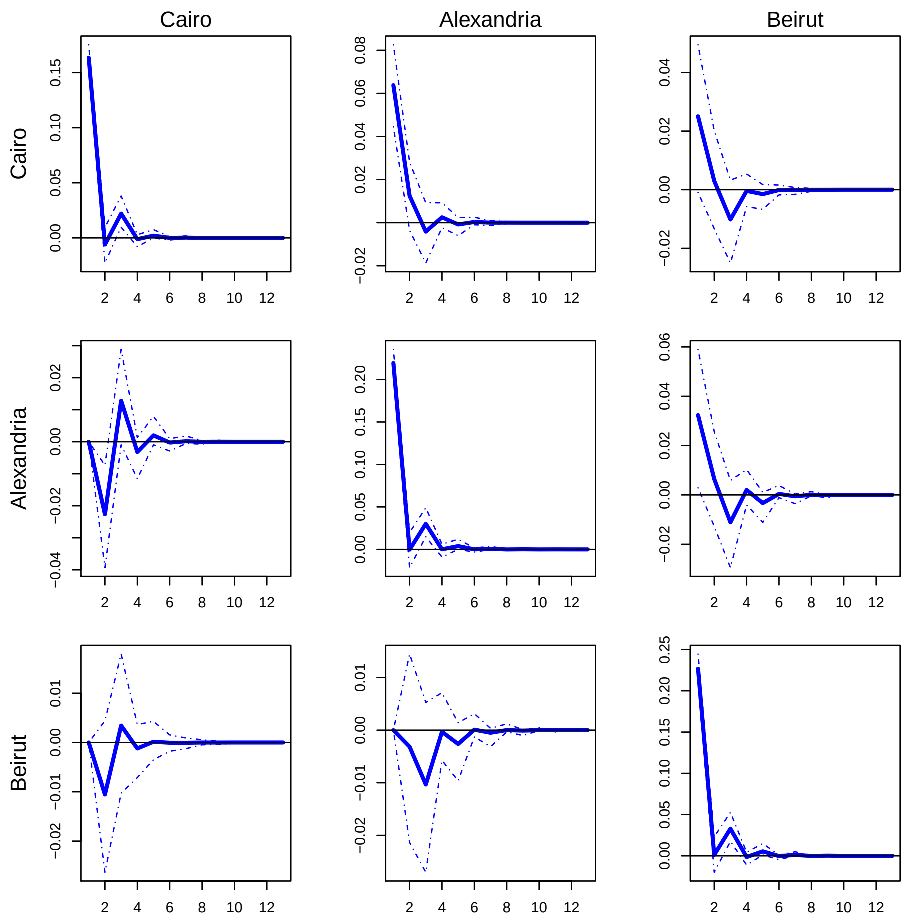

See the note to Figure 4.

Figure 8: Soap: impulse response functions analysis.



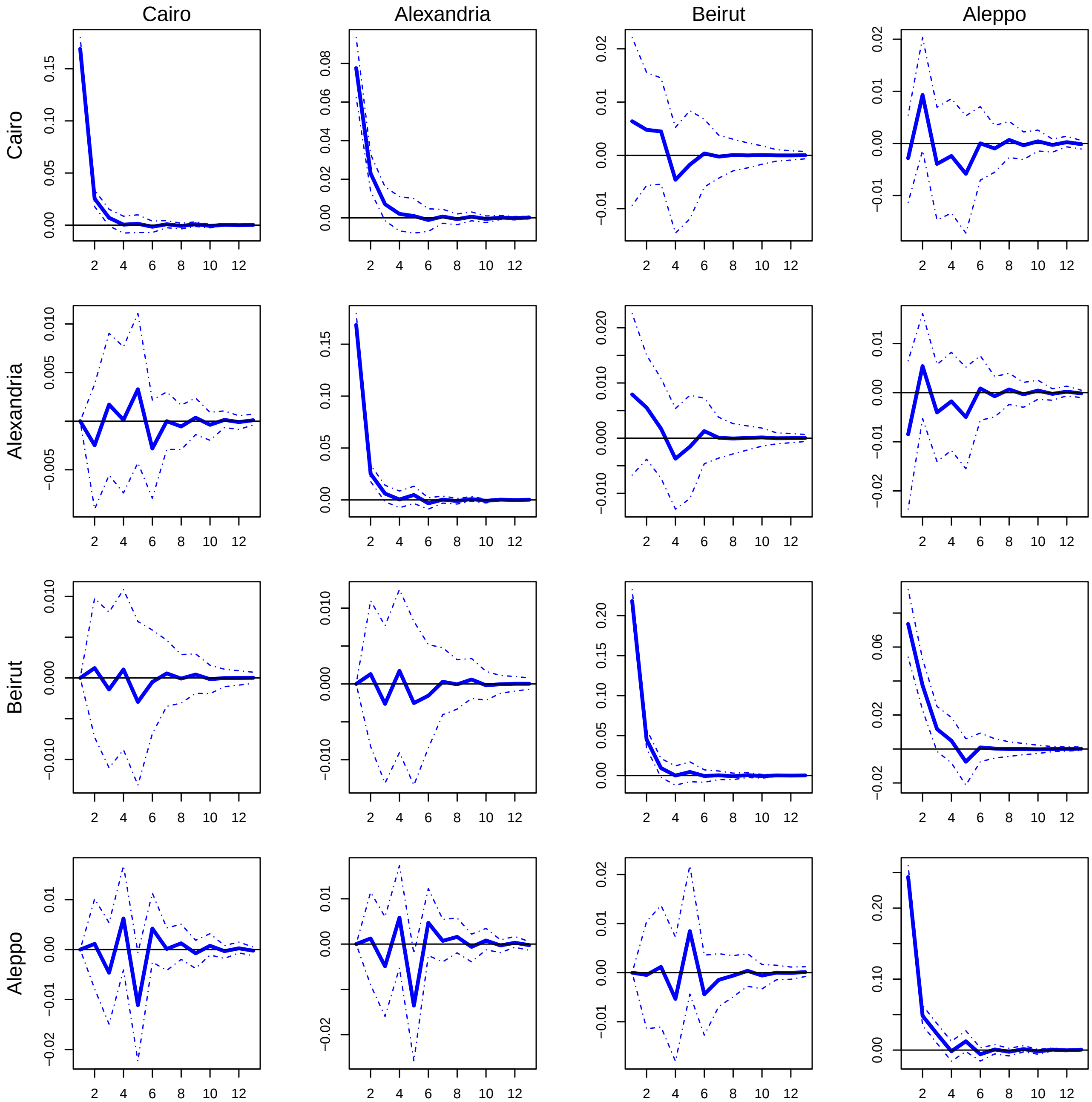

See the note to Figure 4.

Figure 9: Sugar: impulse response functions analysis.




See the note to Figure 4.

Figure 10: Wheat: impulse response functions analysis.




See the note to Figure 4.

# D Additional Materials

## D.1 Tables

Table 5: Posterior moments of shrinkage parameters

| | barley | flour | olive oil | rice | soap | sugar | wheat |
|---|---|---|---|---|---|---|---|
| | Overall shrinkage parameter $\nu$ | | | | | | |
| mean | 9.21 | 6.79 | 15.19 | 7.49 | 7.67 | 4.19 | 10.25 |
| standard deviation | 3.25 | 2.73 | 3.55 | 2.91 | 2.67 | 2.34 | 3.30 |
| | Cointegration shrinkage parameter $\tau$ | | | | | | |
| mean | 0.27 | 0.27 | 0.24 | 0.27 | 0.23 | 0.32 | 0.27 |
| standard deviation | 0.10 | 0.10 | 0.08 | 0.11 | 0.07 | 0.14 | 0.10 |

*Note*: The results are reported for models with lag order, $p = 4$, and cointegrating rank, $r = 1$.

Table 6: Cointegration rank testing for different lag orders

| $r$ | $p = 1$ | $p = 2$ | $p = 3$ | $p = 4$ |
|---|---|---|---|---|
| | Barley | | | |
| 1 | -0.681 | -0.789 | -0.919 | -1.066 |
| 2 | -1.858 | -1.996 | -2.112 | -2.268 |
| 3 | -5.038 | -5.050 | -5.178 | -5.288 |
| 4 | -13.263 | -13.372 | -13.466 | -13.601 |
| | Flour | | | |
| 1 | -1.179 | -1.494 | -1.666 | -1.838 |
| 2 | -1.999 | -2.368 | -2.621 | -2.812 |
| 3 | -8.331 | -8.647 | -8.767 | -8.905 |
| | Olive oil | | | |
| 1 | -6.676 | -7.159 | -7.502 | -7.786 |
| 2 | -6.749 | -8.194 | -8.320 | -9.258 |
| 3 | -8.857 | -10.626 | -10.464 | -11.528 |
| | Rice | | | |
| 1 | -0.909 | -2.188 | -2.479 | -2.600 |
| 2 | -1.752 | -3.393 | -3.849 | -3.994 |
| 3 | -8.675 | -9.783 | -10.100 | -10.216 |
| | Soap | | | |
| 1 | -2.730 | -3.135 | -3.251 | -3.247 |
| 2 | -4.898 | -6.220 | -6.097 | -6.307 |
| | Sugar | | | |
| 1 | 2.180 | -1.560 | -1.197 | -0.093 |
| 2 | 1.304 | -2.429 | -1.976 | -0.810 |
| 3 | -7.562 | -9.367 | -9.038 | -8.239 |
| | Wheat | | | |
| 1 | -1.188 | -1.533 | -1.734 | -1.836 |
| 2 | -2.183 | -2.607 | -2.886 | -3.054 |
| 3 | -5.292 | -5.591 | -5.912 | -5.993 |
| 4 | -14.170 | -14.445 | -14.659 | -14.779 |

*Note*: The table presents the logarithm of Bayes factors equal to the logarithm of the ratio of the posterior probabilities of two models: $\ln(\Pr[M_r|data]/\Pr[M_0|data])$, where $M_i$ is the SVEC model with four lags ($p = 4$) and the cointegration rank $i$.

## D.2 Specification *a priori* of matrices $\tilde{H}$

Barley and wheat:

$$\tilde{H}_{r=1} = \begin{bmatrix} 1 \\ -1 \\ 0 \\ 0 \end{bmatrix}; \qquad \tilde{H}_{r=2} = \begin{bmatrix} 1 & 0 \\ -1 & 0 \\ 0 & 1 \\ 0 & -1 \end{bmatrix}; \qquad \tilde{H}_{r=3} = \begin{bmatrix} 1 & 0 & 0 \\ -1 & 0 & 1 \\ 0 & 1 & -1 \\ 0 & -1 & 0 \end{bmatrix}$$

Flour and sugar:

$$\tilde{H}_{r=1} = \begin{bmatrix} 1 \\ -1 \\ 0 \end{bmatrix}; \qquad \tilde{H}_{r=2} = \begin{bmatrix} 1 & 1 \\ -1 & 0 \\ 0 & -1 \end{bmatrix}$$

Olive oil:

$$\tilde{H}_{r=1} = \begin{bmatrix} 1 \\ -1 \\ 0 \end{bmatrix}; \qquad \tilde{H}_{r=2} = \begin{bmatrix} 1 & 1 \\ -1 & 0 \\ 0 & -1 \end{bmatrix}$$

Rice:

$$\tilde{H}_{r=1} = \begin{bmatrix} 1 \\ -1 \\ 0 \end{bmatrix}; \qquad \tilde{H}_{r=2} = \begin{bmatrix} 1 & 1 \\ -1 & 0 \\ 0 & -1 \end{bmatrix}$$

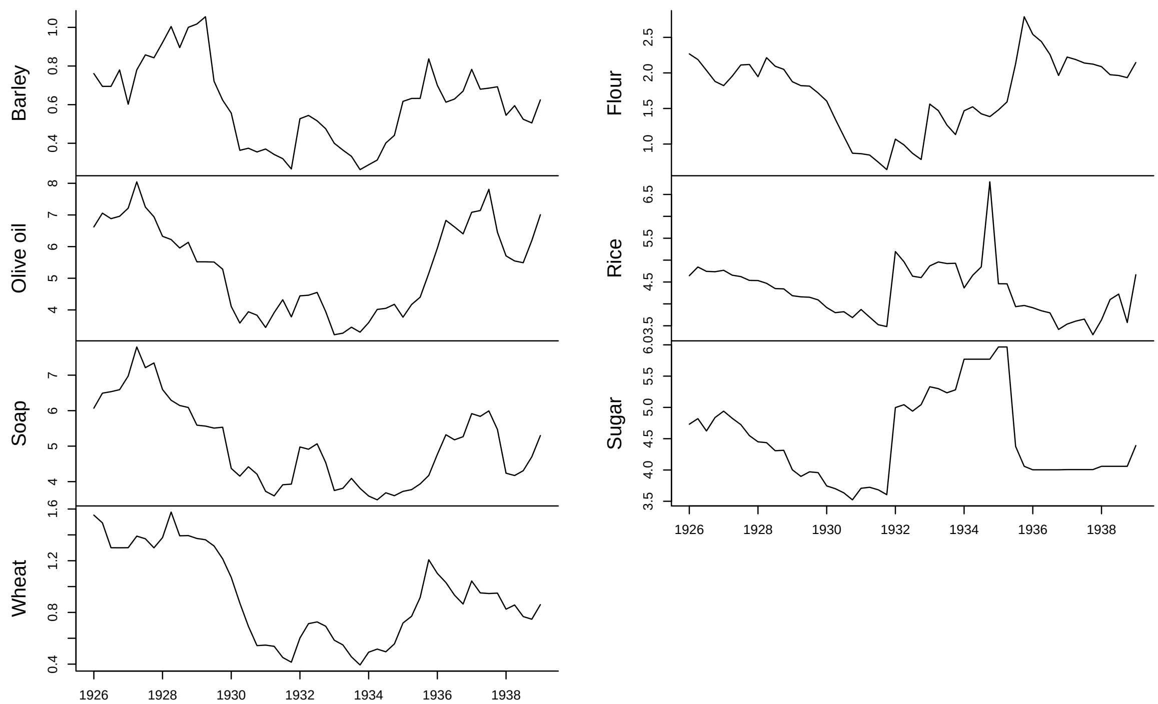
Soap:

$$\tilde{H}_{r=1} = \begin{bmatrix} 1 \\ -1 \end{bmatrix}$$

## D.3 Results for the systems including series for Istanbul

### D.3.1 Istanbul data

Figure 11: Commodity prices for Istanbul, 1926-39 [GB£/kg].



### D.3.2 Cointegration rank testing

Table 7: Cointegration rank testing for different lag orders

| $r$ | $p = 1$ | $p = 2$ | $p = 3$ | $p = 4$ |
|---|---|---|---|---|
| | Barley | | | |
| 1 | -0.137 | -0.265 | -0.517 | -0.672 |
| 2 | -0.707 | -0.783 | -0.974 | -1.194 |
| 3 | -3.329 | -3.207 | -3.485 | -3.687 |
| 4 | -8.736 | -8.658 | -8.867 | -8.941 |
| 5 | -19.420 | -19.491 | -19.488 | -19.706 |
| | Flour | | | |
| 1 | -1.348 | -1.829 | -2.071 | -2.192 |
| 2 | -2.597 | -3.174 | -3.475 | -3.627 |
| 3 | -6.307 | -6.660 | -6.961 | -7.106 |
| 4 | -15.269 | -15.664 | -15.862 | -15.940 |
| | Olive oil | | | |
| 1 | -7.953 | -8.233 | -8.755 | -8.850 |
| 2 | -8.179 | -8.117 | -9.172 | -9.767 |
| 3 | -9.670 | -9.305 | -10.917 | -11.373 |
| 4 | -17.531 | -15.832 | -16.846 | -18.676 |
| | Rice | | | |
| 1 | -2.372 | -2.827 | -3.255 | -3.330 |
| 2 | -3.482 | -4.178 | -4.716 | -4.857 |
| 3 | -7.348 | -7.865 | -8.467 | -8.636 |
| 4 | -16.442 | -16.810 | -17.347 | -17.470 |
| | Soap | | | |
| 1 | -3.757 | -3.986 | -4.196 | -4.086 |
| 2 | -5.174 | -5.515 | -5.661 | -5.944 |
| 3 | -10.194 | -10.931 | -11.205 | -11.440 |
| | Sugar | | | |
| 1 | -3.214 | -3.553 | -3.765 | -3.829 |
| 2 | -4.291 | -4.716 | -4.944 | -5.031 |
| 3 | -8.128 | -8.524 | -8.695 | -8.688 |
| 4 | -17.480 | -17.784 | -17.857 | -17.883 |
| | Wheat | | | |
| 1 | -0.840 | -1.121 | -1.422 | -1.560 |
| 2 | -1.314 | -1.635 | -1.908 | -2.205 |
| 3 | -4.059 | -4.102 | -4.463 | -4.696 |
| 4 | -9.531 | -9.605 | -9.972 | -10.044 |
| 5 | -20.876 | -21.040 | -21.222 | -21.416 |

*Note*: See the note to Table 2.

### D.3.3 Impulse response functions analysis

Figure 12: Barley: impulse response functions analysis for data including Istanbul.

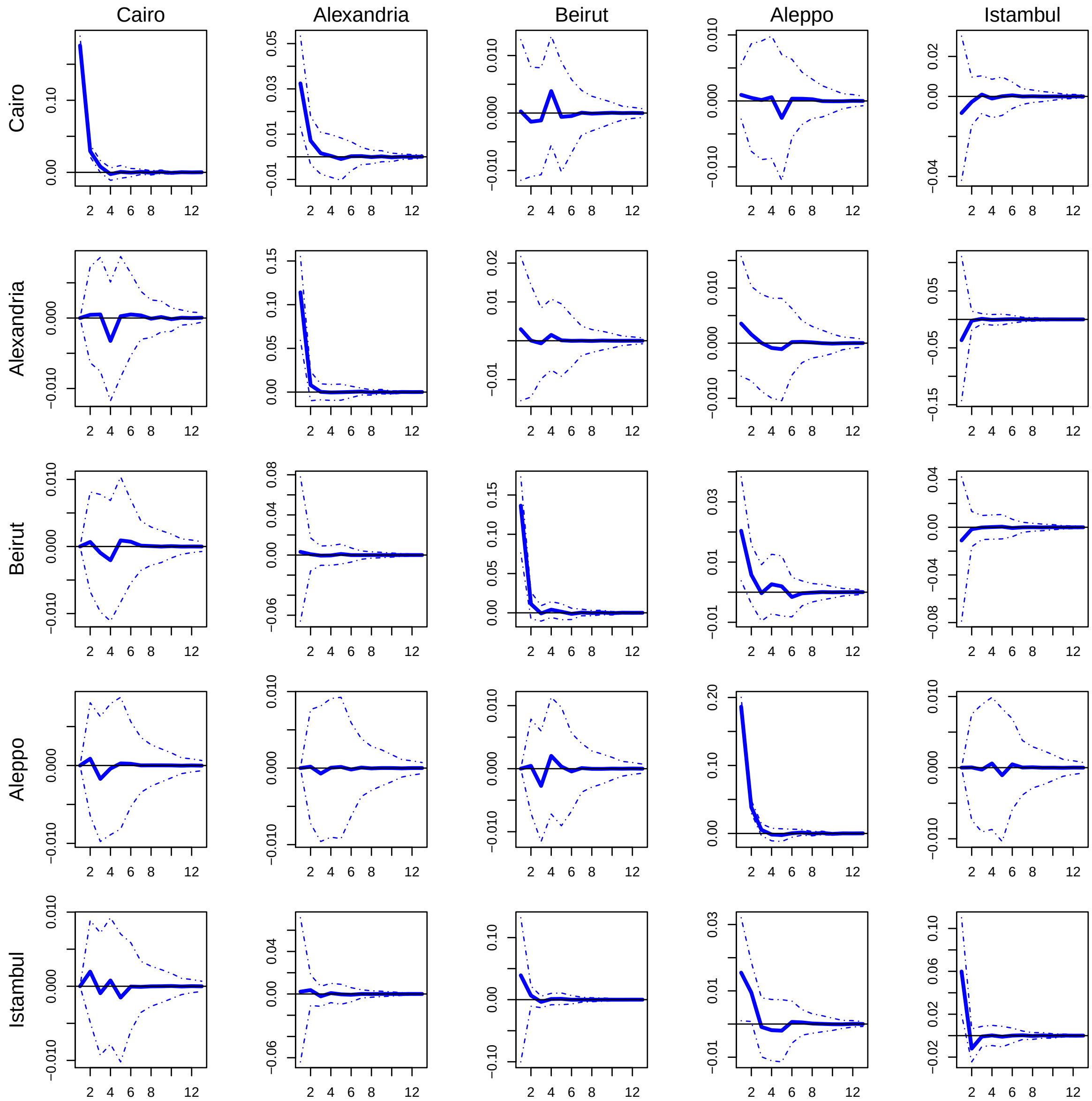


Figure 13: Flour: impulse response functions analysis for data including Istanbul.

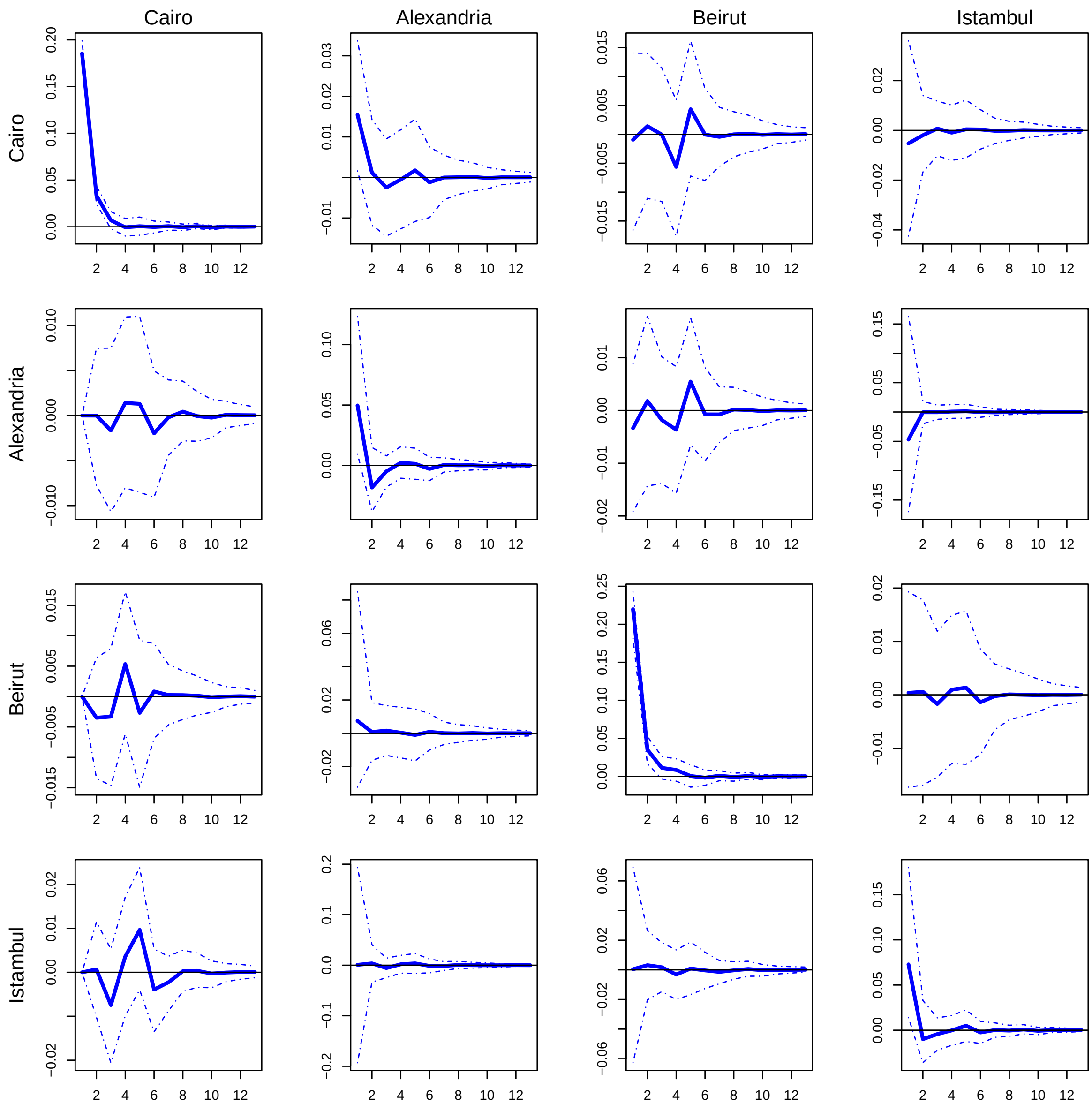


Figure 14: Olive oil: impulse response functions analysis for data including Istanbul.

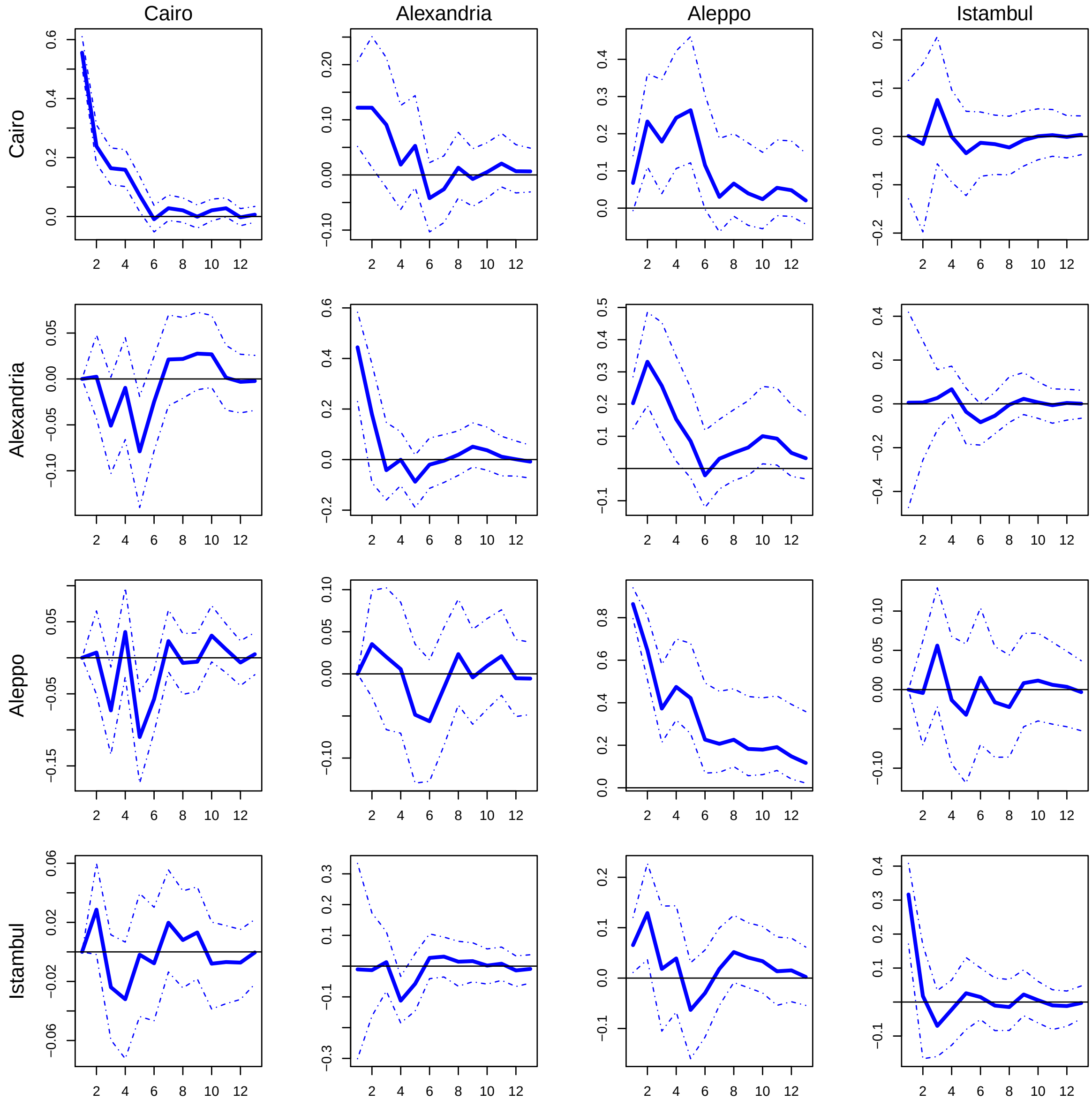


Figure 15: Rice: impulse response functions analysis for data including Istanbul.

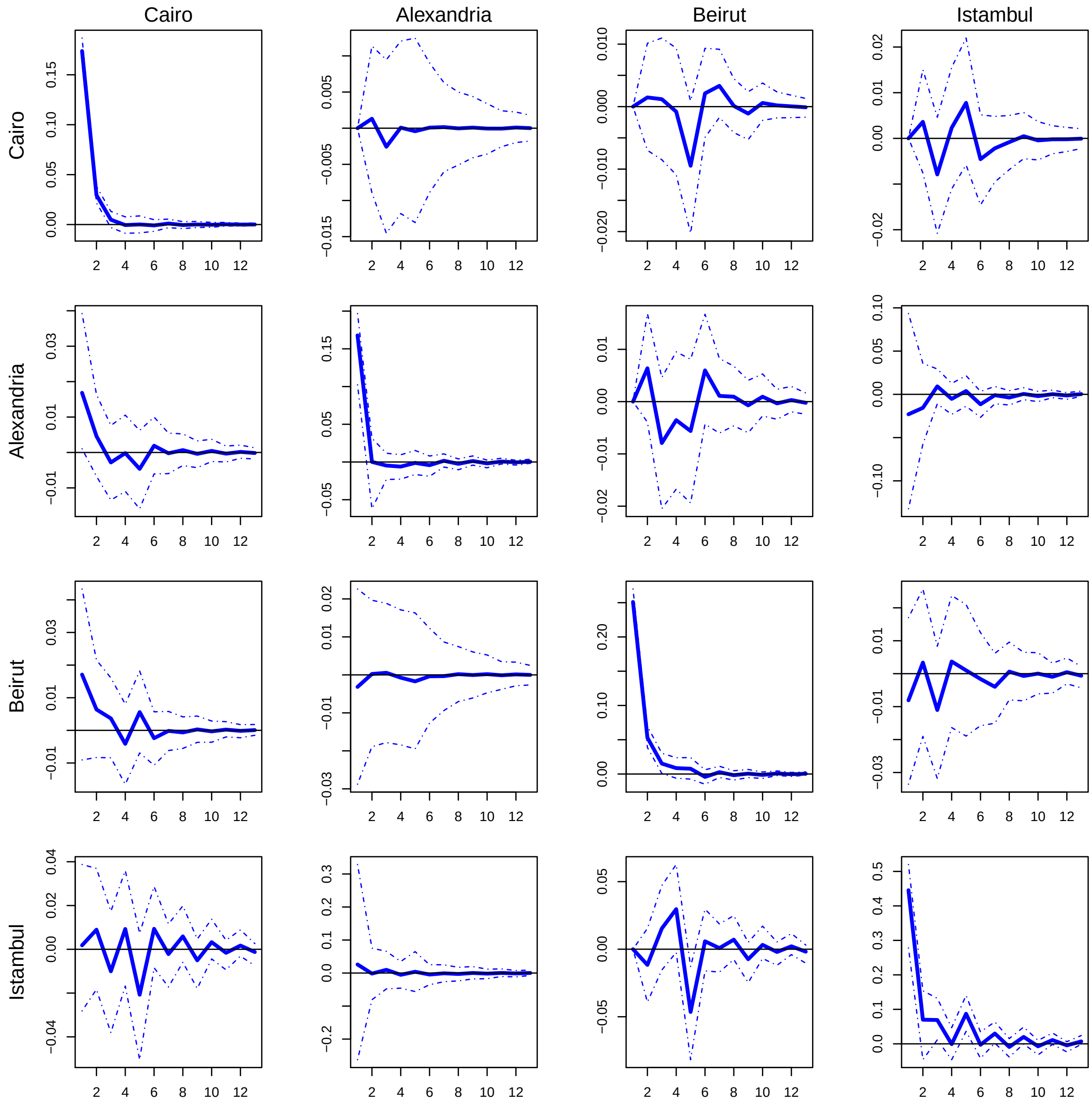


Figure 16: Soap: impulse response functions analysis for data including Istanbul.

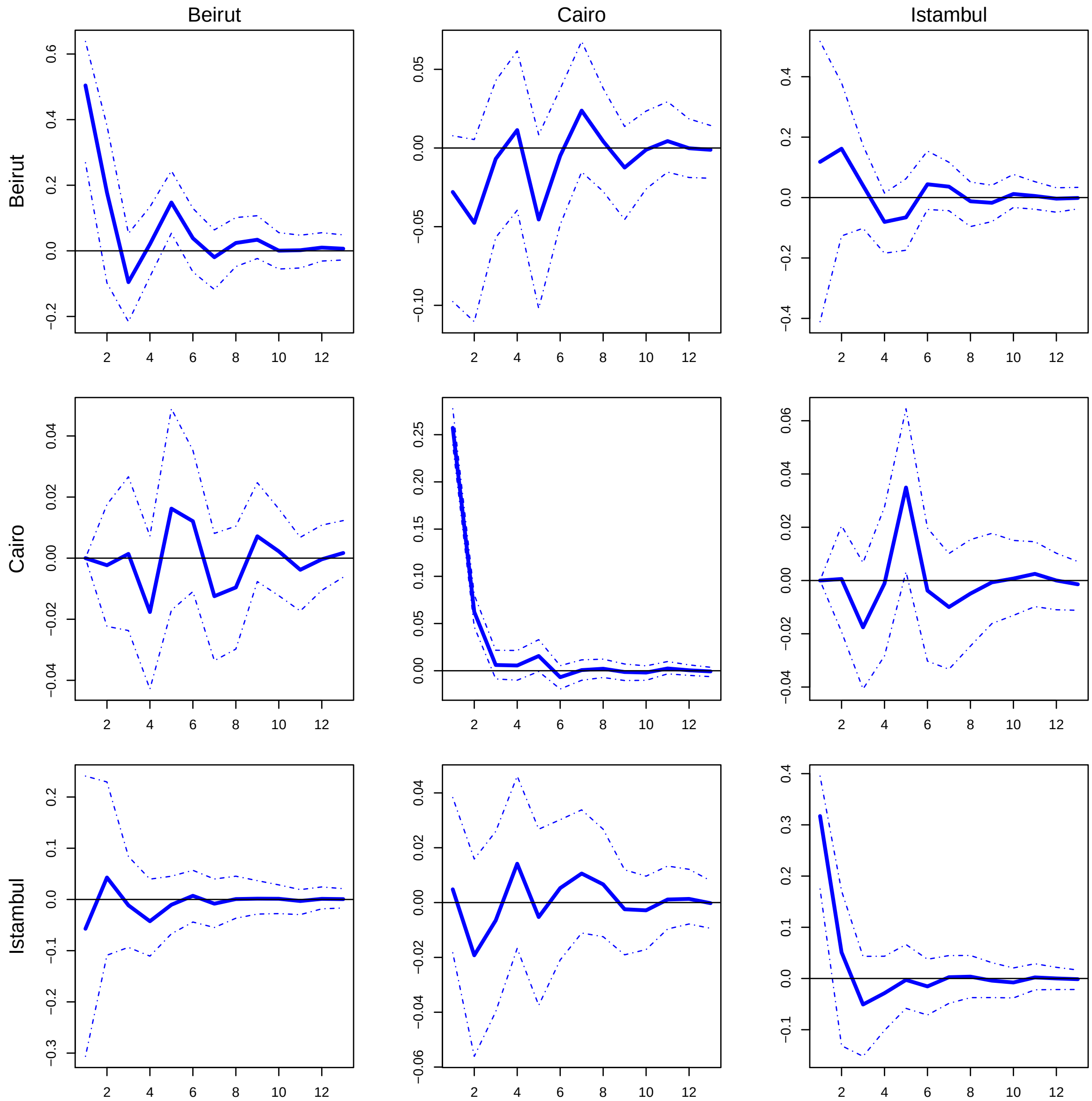


Figure 17: Sugar: impulse response functions analysis for data including Istanbul.

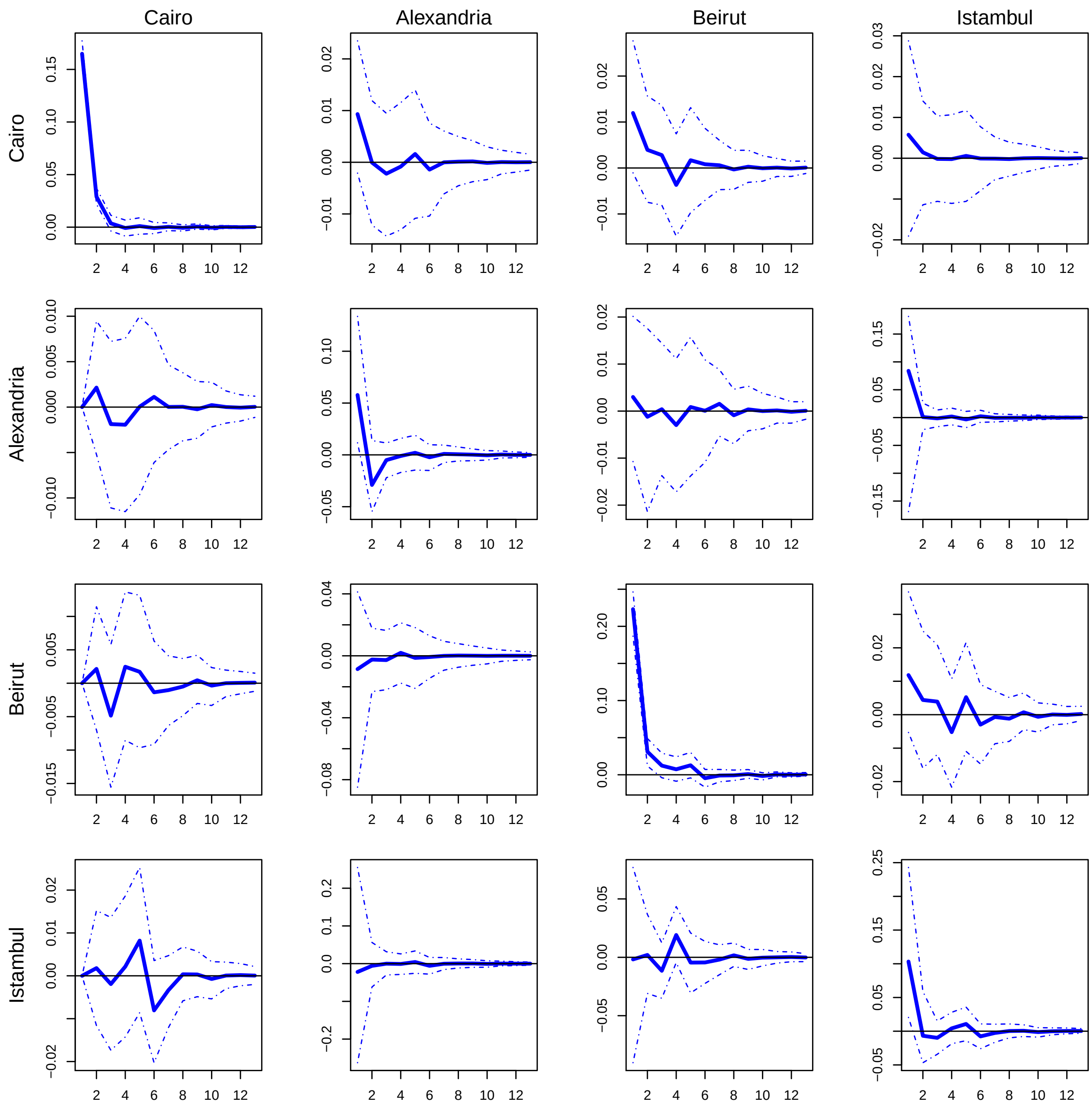


Figure 18: Wheat: impulse response functions analysis for data including Istanbul.