Iran. Econ. Rev. Vol. 21, No.1, 2017. pp. 169-180

Pass-Through Effects of Global Food Prices on Consumer Prices in Iran

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Received: 2016/09/17 Accepted: 2016/11/13

Abstract

The objective of this study is using the Markov Switching Vector Autoregressive method and regime dependent impulse response functions to measure the pass-through of world food prices to consumer price index in Iran from 1990 to 2013. With respect to information criteria and the log-likelihood ratio statistic, MSIA(2)-VAR(1) model has a better fit to data than other models. The magnitude of the pass-through in first and second regimes is respectively 0.43 and 0.94 after two years. We show that the magnitude of the pass-through from world food prices to consumer price index resulting from recent world food price shocks has been higher than before.

Keywords: Domestic Prices, Pass-Through, World Food Prices.
JEL Classification: Q02, C34.

1. Introduction

Food commodity prices substantially increased in global markets during 2007-08 and 2010-11. In the second quarter of 2008, world food price indices were three times higher than in the beginning of the 2000s. Food prices spiked again starting in August 2010 (Von Braun & Tadesse, 2012). Severe weather events, increased demand for basic

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food commodities as biofuel input, rising energy prices, larger meat consumption in emerging countries, exchange rate volatilities and low stock level expectations are the reasons historically as has been revealed in the literature (Yang et al., 2015). Spikes in the global prices of main food commodities represent a direct threat to global development. Sharp increases in food prices rise poverty considerably, reduce the level and quality of nutrition, and decline the consumption other non-food services such as education and healthcare, all of which adversely influence future growth of global economy (Ivanic & Martin, 2014).

It is expected that the rise in food commodity prices on international markets pass-through to domestic consumer and producer prices. Gabrijelcic et al. (2012) discussed that the magnitude and speed of pass-through depends on the margins in the food processing and distribution sectors. Some studies have investigated the pass-through of world food prices to the domestic prices and most of these empirical works have reported a significant effect of world food prices on local consumer prices and inflation. Liu & Tsang (2008) determined that a 10% increase in international commodity prices would result in a 0.24% increase in China’s inflation 3 months later. Zhang & Law (2010) have argued that food price inflation did not generated significant effects on non-food price inflation in China. Jongwanich & Park (2011) concluded that the magnitude of the pass-through has been limited from global food and oil price shocks to inflation in developing Asian countries, and government policies such as subsidies and price controls have played a role in reducing or delaying the pass-through effect.

Jalil & Zea (2011) studied how international food price shocks have impacted local inflation processes in Brazil, Chile, Colombia, Mexico, and Peru in the past decade. The results indicate that international food inflation shocks take from one to six quarters to pass-through to domestic headline inflation, depending on the country. Ivanic et al. (2012) assessed the effect of price changes for 38 agricultural products on poverty using detailed data on patterns of production and consumption in 28 countries. They found an average poverty increase of 1.1% in low income countries and 0.7% in middle income countries.

Gelos & Ustyugova (2012) analyzed inflationary effects of commodity price shocks in 31 advanced and 61 emerging and developing economies. The results showed that domestic inflation significantly responds to global price shocks in economies with higher food shares in consumer price index (CPI) basket, fuel intensities, and pre-existing inflation. Belke & Dreger (2013) investigated the effects of global oil and food price shocks to the consumer prices in Middle East-North African (MENA) countries. According to the results oil and food price shocks increase the domestic prices in the long run. Sivarajasingham & Balamurali (2014) measured the pass-through of global food price inflation in Srilanka. They showed that the global food price pass-through have statistically significant effect on food price and headline inflation in the long and short run.

Economists and policymakers believe that high inflation has negative effects on the economy. There are several channels through which economy has been influenced by inflation. Bonato & Jbili (2009) argued that inflation prevents an efficient allocation of resources by obscuring relative prices; imposes welfare costs on society by increasing distortions, inhibits financial development, making intermediation more costly, and negatively affects the distribution of income and wealth as low income households do not hold assets to hedge against inflation.

Inflation is not a new phenomenon in Iran. The Iranian economy has been faced with inflation in last decades. The consumer price index with annual average increase of 17.33%, increased from 0.5 to 361.7 from 1973 to 2012. Due to subsidy reform, international sanctions and world price volatilities in food and oil markets, the inflation rate in Iran increased from 10.74% in 2001 to 26.62% in 2012. Inflation is a major challenge for policymakers in Iran, and reduction of it to a single digit level was one of the important objectives of Iranian economy in the recent years. Therefore, analysis of high inflation determinants in Iran is useful for policymakers and contributes to a better design of policies. This study investigates how world food prices are passed on to consumer price index in Iran. In order to explore this issue, MS-VAR approach is used.

2. Data and Method

The data set used in this paper comprises GDP, consumer price index, world food price index, exchange rate and interest rate. In this research, we use quarterly data covering the period from 1990:Q2-2013:Q1, published by Iran's Central Bank, Statistical Center of Iran and FAO. This period is chosen due to the availability of the data of variables.

Many previous researches focused on the pass-through of world food prices to domestic prices using linear approaches. In this paper, MS-VAR model is used to estimate the world food price pass-through to consumer prices in Iran. After the estimation of MS-VAR model, regime dependent impulse response functions are used to calculate the magnitude of the world food prices pass-through into domestic prices in Iran. We employed MS-VAR model to capture nonlinearities in the process of price pass-through from world markets into domestic markets. Markov switching (MS) technique was proposed by Hamilton (1990) and later extended to multivariate models such as MS-VAR and MS-VEC by Krolzig (1997, 1999). The MS-VAR model can be written as:

$$y_t = \alpha(s_t) + \sum_{i=1}^{p} A_i(s_t) y_{t-i} + \varepsilon_t$$

Where y_t is a vector that includes endogenous variables, α is the intercept vector, A_i are parameter matrices, s_t presents the regime in time t, and ε_t is an error term. In this case, intercept and endogenous variables depend on regime s_t. This type of MS-VAR models allows for changes in intercept (α) and in the endogenous variables parameters (A_i) in each regime.

In regime-switching models the parameters of a time series model depend upon a stochastic and unobservable variable s_t which represents the state. The stochastic process generating the regimes is a Markov chain defined by the transition probabilities (Krolzig, 2001):

$$p_{ij} = \Pr(s_{t+1} = j | s_t = i), \quad \sum_{j=1}^{M} p_{ij} = 1 \ \forall\, i, j \quad \in \{1, \ldots, M\}$$

In this model, there are M possible states for the y_t and the matrix of transition probabilities is:

$$P = \begin{bmatrix} p_{11} & p_{12} & \cdots & p_{1M} \\ p_{21} & p_{22} & \cdots & p_{2M} \\ \vdots & \vdots & \vdots & \vdots \\ p_{M1} & p_{M2} & \cdots & p_{MM} \end{bmatrix}$$

where p_{ij} shows the probability of changing from regime i to the regime j.

A model where the intercept (I), autoregressive coefficients (A) and variance (H) are regime dependent is denoted by MSIAH(m)-VAR(p), where m indicates the number of regimes and p reflects the order of the VAR.

In this study, the vector of endogenous variables includes the natural logarithm of the world food price index (LWFPI), natural logarithm of the consumer price index (LCPI), natural logarithm of the exchange rate (LER), natural logarithm of GDP (LGDP) and natural logarithm of the interest rate (LR).

In order to calculate the pass-through extent, we need to extract regime dependent impulse response functions after the MS-VAR estimation. The magnitude of pass-through is obtained by dividing the cumulative response of the domestic prices to world food prices shock after *i* quarters by the cumulative response of the world food prices to world food prices shock after *i* quarters. Therefore, the pass-through magnitude is defined as the below equation (Duma 2008; Jalil & Zea 2011):

$$PT_{t,t+j} = CPI_{t,t+j}/WFPI_{t,t+j}$$

where $CPI_{t,t+j}$ is the cumulative change in the consumer price index and $WFPI_{t,t+j}$ is the cumulative change in the world food price index between quarters t and $t+j$.

3. Result and Discussion

At the first stage of time series data analysis, the unit root of the variables should be assessed. For this purpose, we employ the DF-GLS unit root test. We examine null hypotheses for the variables with a time trend and intercept. Table 1 shows the unit root test results. According to the results, the null hypothesis of the unit root (non-stationary) cannot be rejected at the significance level of 5% for any of the variables. Then, the unit root hypothesis was examined for the first difference of variables. Results for each of the first differenced series clearly indicate rejection of the null hypothesis. Therefore, the first differences of the variables are stationary and the integration order of series is one, I(1).

Table 1: Results for Unit Root Test

Variables	DF-GLS		integration order
	level	first difference	
LCPI	-1.15	-3.61**	I(1)
LWFPI	-1.68	-7.41***	I(1)
LER	-1.07	-4.34***	I(1)
LGDP	-1.71	-4.50***	I(1)
LR	-1.74	-9.62***	I(1)

Note: *** and ** imply statistical significance at the 1% and 5% levels, respectively

In next step, we specify the lag length of the VAR model. For this purpose, Schwarz information criterion (SIC) and the Hannan-Quinn (HQ) information criterion are used which support a first order VAR model. Thus, a first order VAR was used in the estimation procedure of MS-VAR model. With respect to the information criteria and the log-likelihood ratio statistic, the MS model was estimated with two regimes that allowed changes in the intercept (I) and autoregressive parameters (A). Therefore, the chosen kind of MS-VAR specification is the MSIA(2)-VAR(1).

Table 2 reports the estimation results of the MSIA(2)-VAR(1) model using the Expectation-Maximization (EM) algorithm, which is a technique for the Maximum-Likelihood (ML) estimation. According to the results, the log-likelihood ratio test for linearity rejects the hypothesis of the linear VAR model against the alternative of the MSIA specification at the significance level of 1%. Hence, the MSIA model seems to fit the data better than a linear model.

The results show that the autoregressive coefficients have same sign in the regimes but most of them vary in value. The constant term has negative sign in the first and second regimes. The first lag of consumer price index (LCPI(-1)) has significant positive relationship with the LCPI at two regimes. The results clearly indicate that LWFPI has a positive impact on the LCPI in the regime 1 and 2. The LER and LR have positive and negative effects on LCPI, respectively.

The results indicate that the chance of remaining in the first regime in the next quarter is 96% and this probability for the second regime is 0.97. In other words, the second regime is more persistent than the first regime. The probability of going from one regime to another regime is greater than zero thereby shifts are not permanent. In other words, the Markov process is irreducible.

Table 2: Results of the MSIA(2)-VAR(1) Estimation

	Regime 1		Regime 2	
Variables	**coefficient**	**t statistics**	**coefficient**	**t statistics**
Intercept	-1.89	-0.60	-1.85	-0.37
LCPI(-1)	0.90	7.25	0.87	7.83
LWFPI(-1)	0.10	1.30	0.09	0.70
LER(-1)	0.09	1.12	0.16	2.17
LGDP(-1)	0.11	0.45	0.07	0.15
LR(-1)	-0.05	-0.28	-0.05	-0.52
SE	0.048		0.048	
Number of obs.	52		39	
duration	26.33		36.74	
regime classification	1990:3 - 2002:4 2008:4 - 2009:1		2003:1 - 2008:3 2009:2 - 2013:1	
Average quarterly inflation (%)	5		4.4	
Average Annual inflation (%)	21.5		17.2	
P_{11} 0.96	P_{22} 0.97	AIC	-16.24	HQ -15.39
Log-likelihood	816.29	LR linearity test	75.72 (0.00)	

Regime 1, which covers the periods 1990:3-2002:4 and 2008:4-2009:1 can be identified as the high inflation state with a 5% average quarterly inflation and 21.5% average annual inflation. Regime 2, which covers the periods 2003:1-2008:3 and 2009:2-2013:1, can be characterized as the lower inflation regime with 4.4% average quarterly inflation and 17.2% average annual inflation. Regime 2 perfectly matches the global food price spikes during the 2007-08 and 2010-11 periods. While the first regime has a larger number of observations with a shorter duration (26.3 quarters), the second regime has fewer observations and longer duration (36.7 quarters). Figure 1 shows the regime classifications using smoothed and filtered probabilities.

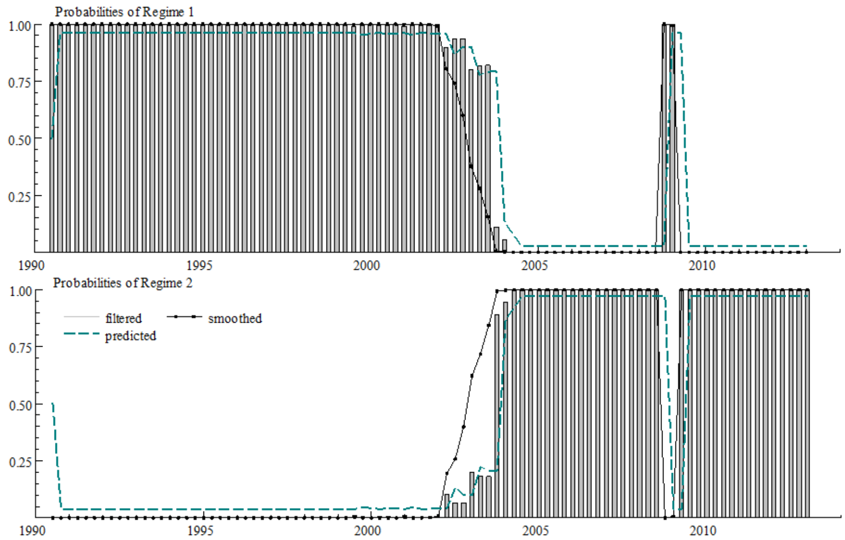

Figure 1: Smoothed and Filtered Probabilities of MSIA(2)-VAR(1) Model

We can calculate regime dependent impulse response functions based on MSIA(2)-VAR(1) estimation results. In this section, we report the estimated response of LCPI to LWFPI shock in each regime. Impulse response functions were used in order to interpret the passing through of world food price shock to consumer price index in Iran.

Figure 2 shows the consumer price index response to a unit shock in the world food price at classified regimes. As would be expected, consumer price index has a positive reaction to world food price shock in both regimes. The results show that the response of consumer price index to world food price shock in second regime is higher than the first regime response. In first regime, the highest effect from the shock was experienced ten quarters after the shock took place. After this point, the response of the consumer price index to world food price shock decreases. In second regime, the domestic prices response to the world food prices shock has an ascending trend.

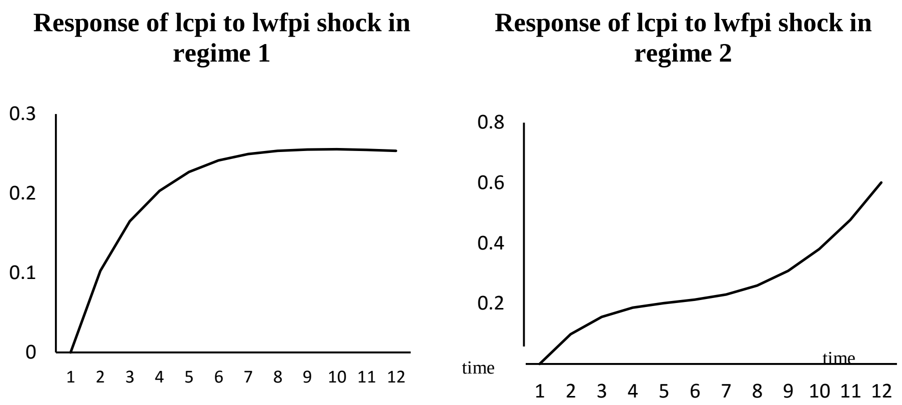

Figure 2: Response of Consumer Price Index to World Food Price Shock

The magnitude of pass-through is computed as the ratio of the cumulative response of the consumer price index to the shock of world food price and the cumulative response of the world food price to its own shock. The pass-through magnitudes were calculated using the regime dependent impulse response functions, which are presented in table 3. In first regime, the world food price pass-through after one quarter is 0.06 and then increases to 0.12 at the end of second quarter. The magnitude of the pass-through in first regime rises to 0.43 after two years. In second regime, the magnitude of initial shock pass-through after a quarter is 0.05. Then, the pass-through magnitude grows to 0.11 and 0.23 after two and four quarters, respectively. Finally, it increases to 0.94 after two years.

Table 3: World Food Price Pass-Through to Consumer Price Index

Quarter	**1**	**2**	**4**	**8**
Regime 1	0.06	0.12	0.23	0.43
Regime 2	0.05	0.11	0.23	0.94

With respect to the findings, there is a considerable difference in terms of the pass-through magnitude between the regimes. The pass-through magnitude in first regime is much lower than second regime. To explain this behavior, we observed that the second regime closely matches the sharp increases in world food prices, subsidy reform policy implementation, currency depreciation and hyperinflation in Iran. Due to reduction of subsidies particularly for food and energy in December

2010, global prices transmitted faster and larger to domestic inflation in Iran. Furthermore, Iran imports a substantial amount of food commodities including grains, oilseeds and meat. Based on annual review of the Iranian central bank, the total value of agricultural products imports of Iran were 13.5 billion US dollars in 2013. Therefore, the rapid currency depreciation and dependence on imports have led to a significant increase in the pass-through magnitude of global food prices to domestic prices in the last decade in Iran.

4. Conclusion

The sudden increases in food prices during last decade have attracted attention from researchers and policymakers and inflationary effects of these shocks have been the subject of much research. This paper presents measures of the pass-through of world food prices to the consumer price index in Iran. Using the Markov switching vector autoregressive method, regime dependent impulse response functions and quarterly data set, the pass-through of world food prices to consumer prices was analyzed between 1990 and 2013.

The results reveal that the MS-VAR model provides a suitable framework for modeling world food price pass-through. The findings are in line with the results of prior studies. The results provide evidence of the differences in magnitude of pass-through among classified regimes. We show that the magnitude of the pass-through from world food prices to the consumer price index in the first and second inflation regimes is respectively 0.43 and 0.94 after eight quarters. During the period of recent world food price shocks, which was characterized with second regime, the magnitude of the pass-through has been larger. Higher dependence on agricultural inputs and food imports, as well as subsidy reform were major determinants of the global food price pass-through to domestic prices in Iran. The Iranian government can reduce the global prices pass-through to domestic inflation by adopting inflation targeting policies and appreciation of the domestic currency.

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