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# Spatial Integration and Asymmetric Price Transmission in Selected Iranian Chicken Markets

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# Spatial Integration and Asymmetric Price Transmission in Selected Iranian Chicken Markets

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*Abstract*— **This study evaluates pattern of price adjustments in selected spatially separated chicken markets in Iran using weekly price data from 1998:17 to 2006:41 including 441 observations in total. The results of Tsay's test suggest that threshold behavior characterize spatial price linkages among the selected markets that imply on using the threshold models. We use the multi-dimensional (two and three regime) threshold cointegration of TAR and M-TAR models. Our results confirm the different speed of adjustment in response to positive and negative shocks in every case. We also utilize impulse response function to investigate dynamic patterns of adjustments in response to shocks.**

*Keywords*— **Spatial Integration, Price Transmission, Threshold Autoregression, Chicken**

## I. Introduction

Market integration usually consider the extent to which price shocks are transmitted from one spatially market to another [1]. In such situation, markets are connected by a process of arbitrage and this will be reflected in the price series of commodities. The extent of co-movement between prices in different markets is considered as a measure of market integration [2]. Markets that are poorly integrated may convey inaccurate price information, leading to the process of arbitrage be a risky activity and so product movements are inefficient.

Shortcoming of earlier methods analyzing market integration and recognition of the significant role of transaction costs, particularly in developing countries, led to application of new methods such as threshold autoregressive models which clearly identify the influences of transaction costs on spatial price linkages. The TAR models recognize thresholds caused by transaction costs result in non–linear pattern of price adjustment. As Serra et al [3] describe, this non-linear pattern is caused by the price differences between two markets being below or above the transaction cost, which will make spatial arbitrage unprofitable and profitable, respectively. The non–linear pattern of price adjustment is represented through a combination of different regimes, which larger shocks (shocks above some threshold) bring about a different response than do smaller shocks, and so threshold effects will be occurred. As Abdulay [4] argues the limitation of this approach is the assumption of constant transaction costs (in proportional terms), implying a fixed neutral band over the period being studied. However, transaction costs and the neutral bands which result may not be constant in the long run and may even be nonstationary [5].

There are a few studies on price transmission in Iranian markets. Specifically only three studies zoom in on agricultural markets which used new methods of price transmission. On the other hand, all of these studies analyzed asymmetric vertical price transmission, and our study is the first that evaluates spatially price transmission in the agricultural markets. The objective of this analysis is to evaluate spatially price linkages among several chicken markets in Iran.

## II. Materials and methods

Consider the Engle and Granger [6] cointegration relationship that defines dynamic long run equilibrium relationship between the price in a given local market ( $P_t^1$ ) and the price in the central market ( $P_t^2$ ):

$$P_t^1 = \alpha + \beta P_t^2 + \mu_t \qquad [1]$$

Where, $\mu_t$ is a random error term with constant variance that can be contemporaneously correlated. With subject to presence of transaction costs asymmetric price transmission is overcoming specifically in the developing countries and so, the usual cointegration method would be misspecified under this condition. Enders and Granger [7] considered alternative specification of error-correction model, called the threshold auto regressive (TAR) model, that can be written as:

$$\Delta\mu_t = I_t \rho_1 \mu_t + (1 - I_t)\rho_2 \mu_{t-1} + \varepsilon_t \qquad [2]$$

where $I_t$ is the Heaviside indicator function such that

$$I_t = \begin{cases} 1 & if \quad \mu_{t-1} \geq \tau \\ 0 & if \quad \mu_{t-1} \prec \tau \end{cases} \qquad [3]$$

and $\tau$=the value of the threshold and $\{\varepsilon_t\}$is a sequence of zero-mean, constant-variance random variables, such that $\varepsilon_t$

is independent of $\mu_j$, j<t. In general, the value of $\tau$ is unknown and needs to be estimated along with the values $\rho_1$ and $\rho_2$. Enders and Granger [7] augmented equation 2 by the lagged changes in the $\{\mu_t\}$ sequence such that it becomes the p*th* – order process:

$$\mu_t = \rho_1 I_t \mu_t + \rho_2 (1 - I_t)\mu_{t-1} + \sum_{i=1}^{p=1} \gamma_i \Delta\mu_{t-1} + \varepsilon_t \qquad [4]$$

The various model–selection criteria (such as AIC or BIC), determine the appropriate lag length in the above equation. In the equation 3, the Heaviside indicator depends on the level of $\mu_{t-1}$. Enders and Granger [7] suggested an alternative such that the threshold depends on the previous period's change in $\mu_{t-1}$. In this situation, alternation Heaviside Indicator can be written as:

$$I_t = \begin{cases} 1 & if \quad \Delta\mu_{t-1} \geq \tau \\ 0 & if \quad \Delta\mu_{t-1} < \tau \end{cases} \qquad [5]$$

The constructed models by the 1, 4, 5 equations called momentum threshold autoregressive (M–TAR) model in which the $\{\mu_t\}$ series exhibits more momentum in one direction than the other. As the TAR model, in the M-TAR model the Heaviside indicator can be used in a dynamic model augmented by lagged changes in $\Delta\mu_t$, too.

The TAR model can captures asymmetrically "deep" movements in a series, while the momentum model can capture the possibility of asymmetrically "sharp" movements in the series [7]. The test statistics for the null hypothesis of symmetric adjustment ($\rho1=\rho2$) and the null hypothesis of no cointegration ($\rho1=\rho2=0$) are called F and F*, respectively. The appropriate critical values for F* statistics are presented in Enders and Siklos [8].

Tsay [9] provided a test for the null hypothesis of AR against alternative hypothesis of TAR. It is a simple nonparametric test for nonlinear adjustment implied by thresholds in an autoregressive series. If the nonlinear adjustment confirmed, then the value of threshold is estimated by using the grid search, which define alternative regimes. Chan [10] suggested the grid search that minimizes the sum of squared error criterion. Balk and Fomby [11] noted that the simple framework of threshold cointegration can be augmented to the multi-threshold models which imply on multi-dimensional parametric regimes. In this study, we follow two- thresholds ($\tau_1$, $\tau_2$) case that imply on three-regime. Two dimensional TAR model would be written as:

$$\Delta\mu_t = \begin{cases} \rho^1 \mu_{t-1} + \varepsilon_t^1 & if \quad -\infty < \mu_{t-1} \leq \tau_1 \\ \rho^2 \mu_{t-1} + \varepsilon_t^2 & if \quad \tau_1 < \mu_{t-1} \leq \tau_2 \\ \rho^3 \mu_{t-1} + \varepsilon_t^3 & if \quad \tau_2 < \mu_{t-1} \leq +\infty \end{cases} \qquad [7]$$

Where $\varepsilon_t$ are the residuals with the zero mean. In practice, the above model augmented with the lagged $\Delta\mu_{t-1}$. The two-dimensional grid search is used to estimate the two $\tau_1$ and $\tau_2$ thresholds. The significance of the differences in parameters across regimes is performed. The critical values have been prepared from Enders and Siklos [8]. If the three-regime TAR is found to not be significant against the AR model, a two-regime TAR is estimated and tested against a standard AR model.

## III. Results and discussion

Our application is to weekly chicken prices from 1998:17 to 2006:41 at Tehran, Qom and Ghazvin markets. Because Tehran was the largest in terms of production volume, taken as the central market and the remaining two markets (Qom and Ghazvin) were taken as the local markets. Data come from the "Iran state livestock affairs logistics (IranSLAL)". The empirical analysis is based upon logarithmic transformations of prices. Our evaluations are of a pair-wise nature, which the linkages between each of local markets prices and the central market prices are considered. The hypothesis that the price series concerned are nonstationary is tested using the augmented Dickey-Fuller (ADF) and Philips-Perron (PP) tests. The results of the ADF and PP tests are presented in table 1. Both tests confirmed a single unit root in each series and so, all of series are I(1).

Table 1. Unit root test for the series

| | Levels | | First-differences | |
|---|---|---|---|---|
| | ADF | PP | ADF | PP |
| Tehran | -1.50 | -2.33 | -7.13* | -18.13* |
| Qom | -1.78 | -2.45 | -7.09* | -19.44* |
| Ghazvin | -1.64 | -2.51 | -6.75* | -17.38* |

Note: * indicate statistically significant in the 1 percent level. Numbers in parentheses indicate the number of lag lengths.

The market integration relation in equation 1 is estimated and tested for cointegration. The estimated models are

$$PQ=-0.452+1.049PT+\hat{\mu}_{t_1} \qquad R2=0.967$$
$$(0.0829)\ (0.0093)$$
$$PG=-0.121+1.014PT+\hat{\mu}_{t_2} \qquad R2=0.986$$
$$(0.0502)\ (0.0056)$$

Where, PT, PQ and PG refer to the prices of the Tehran, Qom and Ghazvin markets, respectively. Numbers in parentheses indicate the value of standard errors. Verification of nonstaionary of price series implies that the estimated standard errors are not consistent, although the estimates of the parameters are consistent. Thus, the formal hypothesis testing cannot be carried out [4]. However, the parameters are all close to one in numerical values that provide support for spatial integration in the chicken markets. Cointegration testing results are presented in table 2. Evidence of long run equilibria among the pairs of series is strong. Given that the Engle-Granger test has lower power than the Johansen test, the latter was also applied to

examine the long run relationship between the prices. This test could not reject the null hypothesis of one cointegration vector between central and local market prices in both cases. The AIC and SBC selected the appropriate order of VAR in the Johansen cointegration test at 3 in both price relationships.

Table 2. Cointegration testing results

| Markets | Test | Test statistic |
|---|---|---|
| Qom-Tehran | ADF test of Engle-Granger | -6.34* |
| | Max eigen value test: r=0 | 63.839* |
| | Trace test: r=0 | 71.217* |
| Ghazvin-Tehran | ADF test of Engle-Granger | -5.69* |
| | Max eigen value test: r=0 | 73.545* |
| | Trace test: r=0 | 81.134* |

Note: * indicate statistical significant in α=0.01 level.

Tsay's test was conducted for the OLS residuals. Test results are presented in table 3. Follow as Obstfeld and Taylor [12], the test run with both increasing and decreasing ordering in the arranged autoregression. In every case, Tsay's test rejects the null hypothesis of no thresholds in the conventional statistical levels and thus suggests that threshold behavior characterizes spatial price linkages among the regional markets.

Table 3. Tsay's test results

| Markets | Tsay's statistic (increasing order) | Tsay's statistic (decreasing order) |
|---|---|---|
| Qom-Tehran | 9.936 (0.000) | 14.772 (0.000) |
| Ghazvin-Tehran | 3.893 (0.021) | 2.419 (0.090) |

Note: numbers in parentheses indicate probability values (P-value).

The three-regime TAR and three-regime M-TAR models were estimated for the central and local markets. The lower and upper thresholds in both TAR and M-TAR models were estimated using the Chan [10] method. The results from three-regime TAR model are given in the upper section of table 4. Based on the both LR and F statistics, there are not any statistically significant differences between estimated parameters. Hence, a three-regime TAR was not statistically significant against an AR model for both pairs of prices. However, based on the F-statistic (F*) estimated parameters are statistically different from zero (ρ1= ρ2= ρ3=0). To estimate the lower and upper thresholds in the three-regime M-TAR model, the same analysis in Chan's methodology was performed. However, instead of $\mu_{t-1}$, the previous period's change in $\mu_{t-1}$ ($\Delta\mu_{t-1}$) was replaced. The results from three-regime M-TAR model are given in the lower section of table 4. Such as the results of three-regime TAR model, both LR and F tests imply on the no statistically significance differences between parameters. Moreover, the value of F*-statistic confirm the statistically difference of parameters from zero. In general, based upon the results, the two-regime TAR and M-TAR models should be estimated.

The results from two-regime TAR and M-TAR models as well as the simple AR model are presented in table 5. The results from AR model are given in the first section of table. The AIC and SBC criteria were selected 3 lags length for both price relationships in an AR model. The estimated coefficient for Qom and Tehran markets is upper than Ghazvin and Tehran markets, indicating a higher speed of adjustment toward the equilibrium for the former markets. The half-lives (HL) given by Ln(0.5)/Ln(1+ρ), represent the period of time (in weeks) required for one-half of a deviation from equilibrium to be eliminated. The estimated HL in a AR model for Qom-Tehran and Ghazvin-Tehran markets are 1.57 and 1.74 weeks, respectively.

As shown in table 5, the null hypothesis of ρ1= ρ2 in all cases was rejected strongly. The critical values for the null hypothesis of ρ1= ρ2=0 are prepared in Enders and Siklos [14]. From the F* statistic the null hypothesis of ρ1= ρ2=0 was not confirmed for none of cases, strongly.

Moreover, the point estimates of ρ1 and ρ2 for Qom-Tehran markets (as an example) in the TAR model are -0.504 and -0.277, respectively. These values suggest that approximately 50 percent of a positive deviation and 28 percent of a negative deviation are eliminated within a week. The same results can be found from the order of related half-lives. So that 1 and 2.11 weeks needed to one-half of a positive and negative deviation from equilibrium to be eliminated. Similarly, in the M-TAR model, the reported point estimates of ρ1 and ρ2 for Qom-Tehran imply that approximately 44 percent of positive and 21 percent of negative deviations are eliminated within a week. So that, 1.2 and 2.92 weeks required for one-half of a positive and negative deviation from equilibrium to be eliminated. Underlying values for the Ghazvin-Tehran markets are 0.89 and 2.24 weeks, respectively.

Since both the TAR and M-TAR models suggest asymmetric adjustment mechanism for the series, it would be interesting to ascertain whether adjustment follows a TAR or M-TAR process. We use the AIC and SBC test values to select the model with the best overall fit. As is evident in table 5, in the Qom-Tehran markets the TAR model and in the Ghazvin-Tehran markets the M-TAR model yields the lowest AIC and SBC. Therefore, for the Qom-Tehran and Ghazvin-Tehran markets, the TAR and M-TAR models are preferable, respectively.

### *A. Threshold Error-Correction Model*

When the price series are integrated and cointegrated, their short run dynamics can be examined with an error correction model [6]. Enders and Granger [7] argued that if the cointegration satisfied, then estimate of symmetric error correction model would be incorrect, because such a model

Table 4. Three-regime TAR and M-TAR models: Estimations and test results

| Markets | τ1 | τ2 | ρ1 | ρ2 | ρ3 | LR | F | F* |
|---|---|---|---|---|---|---|---|---|
| Three-Regime TAR Model | | | | | | | | |
| Qom-Tehran | -0.0762 | 0.0108 | 0.1425 (0.5067) | -0.2952 (0.5022) | -0.3661 (0.5108) | 0.721 [0.697] | 0.355 [0.701] | 4.443 [0.004] |
| Ghazvin-Tehran | -0.0425 | 0.0038 | -0.1794 (0.0033) | -0.3600* (0.0856) | -0.1632 (0.1191) | 4.554 [0.1026] | 2.268 [0.1048] | 24.448 [0.000] |
| Three-Regime M-TAR Model | | | | | | | | |
| Qom-Tehran | -0.0017 | 0.0024 | -0.3097** (0.0654) | -0.2446* (0.1242) | -0.3511** (0.0703) | 0.696 [0.706] | 0.344 [0.709] | 14.924 [0.000] |
| Ghazvin-Tehran | -0.0055 | 0.0035 | -0.4892** (0.692) | -0.5392** (0.1259) | -0.5119** (0.0777) | 0.149 [0.928] | 0.074 [0.929] | 30.88 [0.000] |

Note: * and ** indicate statistically significant in the 5 and 1 percent levels, respectively; numbers in parentheses and brackets indicate standard errors and probability values (P-value), respectively.

Table 5. AR, two-regime TAR and M-TAR models: Estimations and test results

Autoregressive model (AR)

| Markets | ρ | HL | Appropriate lag length | LM statistic |
|---|---|---|---|---|
| Qom-Tehran | -0.329 (0.0514) | 1.74 | 3 | 0.242 |
| Ghazvin-Tehran | -0.356 (0.0625) | 1.57 | 3 | 0.955 |

Threshold autoregressive model (TAR)

| Markets | $\rho_1$ | $HL_1$ | $\rho_2$ | $HL_2$ | Appropriate lag length | LM statistic | τ | F | $F^*$ | AIC | SBC |
|---|---|---|---|---|---|---|---|---|---|---|---|
| Qom-Tehran | $-0.504^{***}$ (0.0952) | 1 | $-0.277^{***}$ (0.0545) | 2.11 | 2 | 1.287 | 0.03055 | $5.07^{**}$ [0.02] | $22.91^{***}$ [0.000] | -6.575 | -6.537 |
| Ghazvin-Tehran | $-0.279^{***}$ (0.0746) | 2.11 | $-0.450^{***}$ (0.0797) | 1.16 | 3 | 1.230 | 0.01080 | $3.52^{*}$ [0.06] | $18.15^{***}$ [0.000] | -7.469 | -7.422 |

Momentum threshold autoregressive model (M-TAR)

| Markets | $\rho_1$ | $HL_1$ | $\rho_2$ | $HL_2$ | Appropriate lag length | LM statistic | τ | F | $F^*$ | AIC | SBC |
|---|---|---|---|---|---|---|---|---|---|---|---|
| Qom-Tehran | $-0.438^{***}$ (0.0520) | 1.20 | $-0.211^{**}$ (0.0932) | 2.92 | 1 | 0.256 | -0.0380 | $5.023^{**}$ [0.02] | $69.44^{***}$ [0.000] | -6.551 | -6.523 |
| Ghazvin-Tehran | $-0.539^{***}$ (0.0588) | 0.89 | $-0.266^{***}$ (0.0834) | 2.24 | 1 | 0.349 | -0.0270 | $8.57^{***}$ [0.003] | $43.23^{***}$ [0.000] | -7.468 | -7.440 |

Note: $^{*}$, $^{**}$ and $^{***}$ indicate statistically significant in the 10, 5 and 1 percent levels, respectively; numbers in parentheses and brackets indicate standard errors and probability values (P-value), respectively; The critical values of F* in the TAR model (case of two- variables and 4 lags) are 6.35, 7.41 and 9.88 at the significant levels of 10, 5 and 1 percent, respectively; The underlying values for the M-TAR model are 5.52, 6.56 and 8.91; the respective 95% critical value for LM test is 3.84 in comparison with the estimated values for both price linkages reject the null hypothesis of autocorrelation.

would not reveal differential adjustments to positive and negative deviations. Hence, we employed asymmetric error correction models as follow:

$$\Delta P_t^1 = \sum_{s=1}^{k} \alpha_s \Delta P_{t-s}^{\ 1} + \sum_{s=0}^{k} \beta_s \Delta P_{t-s}^{\ 2} + \gamma_1 Z_plus_{t-1} + \gamma_2 Z_\min us_{t-1} \quad [8]$$

Where k is the lag length, the Z-plust-1 and Z-minust-1 are the error correction terms from the threshold cointegration regressions and show the adjustment to the positive and negative shocks, respectively. The error-correction terms for the Qom-Tehran markets are as follow:

$$Z_plus_{t-1}=I_t\,(P_{t-1}^{\ Q}+0.452-1.049P_{t-1}^{\ T})$$
$$Z_minus_{t-1}=(1-I_t)\,(P_{t-1}^{\ Q}+0.452-1.049P_{t-1}^{\ T})$$

$I_t$ is the Heaviside indicator function with the consistent threshold. The value of consistent threshold in this case is τ=0.03055. The error correction terms for the Ghazvin-Tehran markets are as follow:

$$Z_plus_{t-1}=I_t\,(P_{t-1}^{\ G}+0.121-1.014P_{t-1}^{\ T})$$
$$Z_minus_{t-1}=(1-I_t)\,(P_{t-1}^{\ G}+0.121-1.014P_{t-1}^{\ T})$$

Where $I_t$ is the momentum Heaviside indicator function with the consistent threshold that its value in this case is τ=0.027. The results of the estimated coefficients are presented in table 6. The AIC and SBC were used to determine the lag length. The results show that in both Qom-Tehran and Ghazvin-Tehran markets, the coefficients on the positive and negative error correction terms are significant at conventional levels, implying that the prices in local markets respond to both positive and negative discrepancies in the long run price equilibrium. However,

the threshold TAR model suggest much faster adjustments in response to negative deviations from equilibrium (such as increase in Tehran chicken prices) compared with positive deviations (such as decrease in Tehran chicken prices). Vice versa, in the Ghazvin prices the positive shocks tend to adjust to equilibrium relatively quickly whereas the negative shocks tend to persist. The point estimates for Qom-Tehran (as an example) imply that Qom prices adjust so as to eliminate about 32 percent of a unit negative change, but 26 percent of a positive change in the deviation from the equilibrium relationship created by changes in Tehran prices. The hypothesis of complete market segmentation, which implies that none of the Tehran prices significantly influences local market prices, can also be rejected for both Qom and Ghazvin markets.

Table 6. Asymmetric error correction model results

| | $\Delta P_Q$ | $\Delta P_G$ | | $\Delta P_Q$ | $\Delta P_G$ |
|---|---|---|---|---|---|
| Constant | -0.001 | -0.0006 | ΔPT(t-1) | 0.311** | 0.253** |
| | (0.0021) | (0.001) | | (0.062) | (0.068) |
| ΔPQ(t-1) | -0.237** | | ΔPT(t-2) | 0.195** | 0.187** |
| | (0.056) | | | (0.055) | (0.062) |
| ΔPQ(t-2) | -0.146** | | ΔPT(t-3) | | 0.156** |
| | (0.050) | | | | (0.052) |
| ΔPG(t-1) | | -0.193** | Z_plust-1 | -0.259** | - |
| | | (0.065) | | (0.066) | 0.428** |
| ΔPG(t-2) | | -0.150** | Z_minust- | -0.320** | -0.17* |
| | | (0.060) | 1 | (0.072) | (0.090) |
| ΔPG(t-3) | | -0.154** | LM | 0.496 | 1.119 |
| | | (0.052) | | | |
| ΔPT | 0.797** | 0.931** | LR | 312.3 | 580.8 |
| | (0.042) | (0.027) | | [0.000] | [0.000] |

Note: * and ** indicate statistically significant in the 5 and 1 percent levels, respectively; numbers in parentheses and brackets indicate standard errors and probability values (P-value), respectively; LM statistic confirm no autocorrelation in both error correction models; LR statistic confirm statistical significance of Tehran price coefficients joint with the error correction terms in both local markets.

Since the impulse response functions (IRFs) provide richer inferences regarding the dynamics of price adjustments than standard regression analyses [13] in this section IRFs are developed which estimate the impacts of positive and negative shocks to the price spread between the central and local markets.

Based on the previous results, a unit increase in the Tehran prices lead to 0.797 unit increase in the Qom prices and so 0.203 unit decreases in price spread that 0.3201 unit of this decrease eliminate within a week. This decrease in price spread eliminate within 10 weeks. While, the price spread between Tehran and Qom returns to its equilibrium level within 14 weeks after experiencing a positive shock. Thus, there is asymmetric response to positive and negative shocks for Qom prices.

Whereas for Tehran and Ghazvin with positive shocks reverting faster to the equilibrium level than negative shocks. Such that, the price spread between Tehran and Ghazvin returns to its equilibrium level within 18 and 10 weeks after experiencing negative and positive shocks, respectively. The related figures are available from authors.

## IV. Conclusions

Because of the shortcoming of prior approaches the threshold cointegration methods has been considered for price transmission, recently. This study evaluated pattern of chicken price adjustments between Tehran as a central market and Qom and Ghazvin as local markets. The results of Tsay's test imply on asymmetric price transmission indicate using threshold models. At first, we estimated the three regime TAR and M-TAR models where, the results indicate on the no significance difference between parameters. This imply on using the two-regime models. Based on the AIC and SBC criteria the two regime TAR and M-TAR models determined as the best fit to Qom-Tehran and Ghazvin-Tehran markets, respectively. We also estimate the simple AR model, but we could not get a positive result of much faster adjustments in response to deviation from equilibrium in the threshold models than the simple AR. Moreover, as expected, Qom-Tehran markets suggest much faster adjustments in response to negative shocks (such as increase in Tehran prices) than positive shocks (such as decreas in Tehran prices). Such that based on the IRFs results for Qom-Tehran, 14 and 10 weeks require for adjusting the positive and negative shocks, respectively. However, in the Ghazvin-Tehran markets the results are not as expected. Such that the speed of adjustment is much faster to positive shocks (10 weeks) than negative shocks (18 weeks). It may be explained by small-sample period. Since Hansen [14] and Enders and Falk [15] showed that OLS estimates of the speed of adjustment terms have poor small-sample properties. However, because of unavailable observations before the 1998, we could not use the larger sample in our analysis. Moreover, the more little share of Qom market in chicken productions than to Ghazvin and Tehran that lead to the Qom market be as a price taker from Tehran market, directly.

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