**ABSTRACT:** The aim of this study is to research the impact of minimum wage on unemployment, prices, and growth for the Turkish economy. The data used is monthly and covers the period from January 2005 to March 2017. The producer price index represents prices and the industrial production index represents growth. The Autoregressive Distributed Lag (ARDL) model is used to see the effect of the minimum wage on these variables. An error-correction-based Granger causality test is then conducted to see short-run and long-run causalities. The bounds test yields evidence of a long-run relationship between variables. The obtained ARDL results also show that while the minimum wage has a statistically significant effect on unemployment and prices, it does not have a statistically significant effect on production. While there is short-run causality from minimum wage to prices only, the obtained significant error correction terms indicate long-run causality for all of the variables. Consequently, the minimum wage plays a significant role in increasing prices and the number of unemployed people in Turkey.

**KEY WORDS:** Minimum wage, unemployment, time series analyses

**JEL CLASSIFICATION:** J3, E24, C32
1. INTRODUCTION

The effect of the minimum wage on employment, prices, and growth is a topic of debate in economics literature. Some economists maintain that the result of a downward-sloping labour demand curve and an upward-sloping labour supply curve is an increase in the minimum wage that is higher than the market wage, and that the efficiency level will increase unemployment. Such an increase in the minimum wage raises labour costs and compels small-scale firms to dismiss some of their workers. Thus, increasing the minimum wage results in unemployment.

On the other hand, opposite viewpoints claim that an increase in the minimum wage results in an increase in unemployment, as employers respond to minimum wage increases via the employment margin. However, the effects of a minimum wage increase can be offset by reducing fringe benefits and thus reducing the overall labour costs, even if businesses do not reduce the number of workers. This explains why the decrease in the employment level is not as great as indicated by the standard labour supply and demand model (Lee 2004: 658). In addition, since such workers’ marginal propensity to consume is high, a wage increase also increases demand, stimulating the economy and increasing economic growth. This also prevents unemployment, so the effect of an increase in the minimum wage on unemployment may not be as the standard demand and supply model of the labour market indicates.

The effect of the minimum wage is also explained by considering the term length. Akin (2017) indicates that the effect of a minimum wage increase changes depending on whether the term is short or long. In the short run a minimum wage increase can result in increased labour costs, unemployment, and inflation and decreased production and exports. On the other hand, in the mid- and long-term an increase in both the minimum wage and other wages increases total demand and production. An increase in the minimum wage also increases other wages, which increases the wage share in GDP and raises the share of income of workers in the lowest income bracket, thus helping to regulate income distribution.

A wage increase means an increase in costs for producers. If producers do not reflect this increase in prices, cost increases do not form an inflationary pressure. Also, if wage increases parallel an increase in efficiency, costs are not raised and
inflation is not pushed. However, if producers reflect the cost increase in prices a rise in inflation results. Moreover, if the wage increase is above the efficiency level an increase in the minimum wage results in cost inflation and this may cause a price–wage spiral.

Changes in the minimum wage also affect income distribution. If an increase in the minimum wage is covered by the producer’s profit, this can result in better income distribution. On the other hand, if a minimum wage increase is covered by labour, while some workers will benefit from an income increase, others will lose their job. In this case, a minimum wage increase will negatively affect income distribution [Talas (1972) in Özdemir et al. 2012: 2].

Since the theoretical and empirical literature is conflicted regarding the effects of the minimum wage on unemployment, prices, and production, our aim is to test this relationship for Turkey, which recently experienced a significant minimum wage increase. The ARDL Bounds test developed by Peseran, Shin, and Smith (2001) is used to test the long-run relationship, and then ECM-based Granger causality is conducted to see the short- and long-run causalities. The obtained results show that the minimum wage has a statistically significant effect on unemployment and prices. There is short-run causality from minimum wage to prices only, but long-run causality for all three variables.

The rest of the study is organized as follows. The second section explains the theoretical and empirical literature. The third section briefly presents background information about the minimum wage in Turkey. The fourth section reviews the data and methodology used. Lastly, the empirical results and conclusion of the study are presented.

2. THEORETICAL AND EMPIRICAL LITERATURE REVIEW

Before explaining the empirical literature, it is useful to define and give a brief history of the minimum wage.

In ILO member countries, governments set a wage base for wage bargaining. This wage is called the minimum wage and represents the minimum wage required for the employee to live at a minimum level according to the economic and social conditions of the day. In other words, the minimum wage is the base rate
according to the laws and regulations that will meet the minimum requirements of the worker. The minimum wage is the most effective tool to improve the quality of life of the low-skilled, non-competitive, and low-wage workforce (Akgeyik and Yavuz 2006:2).

The minimum wage law resulted from the industrial revolution, when wages were so low that only one person could survive on them, so all family members had to work, including women and children. This historical experience shows that without government intervention, in the market system wages are always subject to being below the minimum wage, as the employer always has more power than the employee. Therefore, determining a minimum wage has become an indispensable necessity (Akın 2017: 141). Namely, because of exploitation of workers, there has been a need for minimum wage law in all societies.

A minimum wage law was first enacted in New Zealand in 1894 and was adopted in Australia in 1896. In the early 1900s, England, France, Germany, and Austria also adopted a minimum wage law, followed by numerous other industrialized and developing countries. A minimum wage ensures that workers receive a fair wage for their work (Neumark and Wascher 2008: 11).

The costs and benefits of the minimum wage are still a subject of debate. In the theory there are two opposing views: one claims that a minimum wage is advantageous for the economy and society, while the other claims that it is damaging.

A minimum wage may increase productivity, since the most efficient workers are chosen and the less efficient workers remain unemployed, thus encouraging less-skilled workers to make more effort. In this way the minimum wage helps to select the most efficient workers, the best equipped employers, and the most advantageous forms of industry. Moreover, it does not deteriorate any of the factors of production (Webb 1912: 979–984).

In Keynesian terms, since low wage groups’ marginal propensity to consume is high, a wage increase will increase total demand, which contributes to economic growth. On the other hand, it will stimulate inflationary pressures. In the long run a demand increase that stems from a wage increase will result in recovery in the economy, increased growth, and increased private sector revenues, which
compensate for the short-run losses. Since the increase in spending increases indirect taxes and the increase in income increases income tax, tax income increases. Thus the medium- to long-term effects of a wage increase are positive (Akin 2017: 146).

The opposing view holds that a minimum wage increase is damaging. A minimum wage policy may have negative effects when this wage is above the equilibrium level because it will decrease employment. Brown et al. (1982) and Card et al. (1993) indicate that these negative effects are weak in the US (Maloney and Mendez 2000: 110). However, other studies show the minimum wage having a strong effect on employment. Since an increase in the minimum wage raises production costs, the competitive power of small and medium-size enterprises especially may decline and cause dismissal of workers. Small and medium enterprises can compete with big technology-based companies only by using their lower-wage advantage. Therefore, when the minimum wage increases, the competitive power of small and medium enterprises decreases. Some businesses pass the wage increase on to prices, increasing inflation. Businesses that cannot pass the wage increase on to prices make a loss and have to dismiss workers, increasing the unemployment level. Therefore, wage increases can affect macroeconomic variables in both the short and long run. As explained, the short-run effects are increased costs, unemployment, and inflation and decreased production and exports. On the other hand, in the long run, increases in both the minimum wage and other wages increase demand, and so production (Akin 2016: 144–145).

Bates Clark (1913), a Neo-classical economist, also claims negative effects of the minimum wage. He argues that in the absence of any new demand for labour, increasing wages will lessen the number of workers employed. Furthermore, Lees Smith (1907) and Taussing (1916) claim that a minimum wage may reduce the employment level of low-skilled workers (Neumark and Wascher 2008: 15).

As in the theoretical literature, in the empirical literature there are many different studies that research the relationship between the minimum wage and macroeconomic variables.

Reynolds and Gregory (1965) and Castillo and Freeman (1992) find for Puerto Rico that an increase in the minimum wage causes unemployment. Reynolds and
Gregory (1965) also conclude that minimum wage increases in Puerto Rico often result in large efficiency improvements in companies with the same capital and labour. Card (1992a) researches the effect of a minimum wage increase on employment for the US in 1989–1990 and concludes that there is no statistically significant relationship. In contrast to Card’s study, Neumark and Wascher’s (1992) study supports the hypothesis that a minimum wage increase reduces employment. They research the employment effects of minimum and subminimum wages using panel data for 50 states for the years 1973–1989 and conclude that an increase in the minimum wage reduces employment among young workers, with employment elasticities with respect to the minimum wage ranging between –0.1 and –0.2 for teenagers and between –0.15 and –0.2 for youths. Card (1992b) finds for California that a minimum wage increase raises teenage employment by about 0.35%.

Katz and Krueger (1992) study the effect of a 1991 minimum wage increase on fast food restaurant’s employment in Texas. Using surveys, they conclude that the minimum wage has a large positive and statistically significant effect on employment. Card and Krueger (1994) also use a survey to examine the effect of a minimum wage increase on fast food restaurant employment in New Jersey and also find a positive relationship between minimum wage increase and employment. Machen and Manning’s (1994) cross-industry study uses UK data and concludes that the minimum wage does not have an adverse impact on employment. They find that employment growth is positively related to minimum wage growth but the coefficient is not statistically significant. Deere, Murphy, and Welch (1995) investigate the effects of a minimum wage increase on low-income and high-income employees, taking into consideration demographic and geographic features. They conclude that the minimum wage increase reduces employment for the low-income group, but find the opposite for the high-income group.

For developing countries, Montenegro and Pages (2004) make a panel data analysis for China and conclude that while an increase in the minimum wage reduces employment for the young and unskilled, it increases employment for women. Lemos et al. (2004) use time series data for Brazil and find that the effect of the minimum wage on employment differs from region to region and for different wage levels. Suryahadi et al. (2003) investigate the effect of the minimum
wage for Indonesia and conclude that while the minimum wage has a negative effect on integrated urban employment, male employment, and female employment, it has a positive effect on unskilled labour and unskilled officer employment. Neumark et al. (2006) research the effect of the minimum wage using Brazilian data and find that an increase in the minimum wage decreases employment for the head of the family but increases employment for other family members.

There are very few studies relating to the Turkish economy. Using Granger causality, Korkmaz and Çoban (2006) research the causal relationship between minimum wage and inflation for Turkey for the period 1969–2006 and find bidirectional causality between the variables. Using an error correction model, Özdemir, Mercan, and Erol (2012) research the effect of the minimum wage on unemployment and find that an increase in the minimum wage increases unemployment in the long run. Dağlıoğlu and Bakır (2015) investigate the relationship between minimum wage and employment on a sectoral basis using panel regression models for 16 sectors, and conclude that there is a positive relationship between minimum wage and employment.

As seen from the literature, the effect of the minimum wage on employment, prices, and growth is disputed. Therefore, our aim is to examine this relationship for a specific country, Turkey, which recently experienced a significant minimum wage increase.

3. MINIMUM WAGE DEVELOPMENTS IN TURKEY

Although in Turkey minimum wage regulations change from time to time, they entered legislation with Labour Law No. 1936 and then were part of Labour Laws No. 1475 and 4857. In 1973, in an agreement with the ILO, this issue became an international commitment. Between 1951 and 1967 local commissions determined the minimum wage, but since 1971 it has been determined by central powers at the country level (Akgeyik and Yavuz 2006:3).

Increasing the minimum wage in Turkey, as in other countries, has always been a topic of debate. Since approximately 40% of employees in Turkey receive the minimum wage, a large proportion of the population is constantly demanding a
minimum wage increase. The highest increase in the minimum wage was 30% in January 2016.

**Graph 1:** Minimum wage (in Turkish Lira) in the sample period

![Minimum wage graph](image)

**Source:** Turkish Statistical Institute

This also caused other wages to increase. Some of the possible results of a minimum wage increase in Turkey are (Akin 2016: 146):

- Small and medium-sized enterprises in particular become less competitive as a result of the wage increase, forcing them to reduce production.
- In countries which Turkey competes with in foreign trade – China, India, Pakistan, Egypt and other African countries, and European countries such as Croatia, Poland, Slovakia, Hungary, the Czech Republic, Estonia, and Romania – wage levels are lower than in Turkey. Raising the minimum wage in Turkey weakens competitiveness, especially in the more labour-intensive sectors. Since Turkey does not have the power to determine export prices, the minimum wage increase is not reflected in prices and no other cost decrease exists to compensate for the wage increase. Depending on the cost increase, companies’ profits may decrease, leading to decreased exports and increased
unemployment. Furthermore, decreased exports may lead to increased foreign trade and current account deficits.

- Some businesses that sell their products in the domestic market pass on this cost burden to prices, increasing inflation. Businesses that cannot pass on this increase to prices make a loss, again increasing unemployment.

Despite these possible effects of the minimum wage on unemployment, prices, and growth, unions insist on increasing the minimum wage further.

4. DATA AND METHODOLOGY

Our empirical estimation is based on monthly observations that cover the period January 2005 to March 2017. The variables used are nominal minimum wage, unemployment rate, industrial production index (representing growth), and producer price index (PPI). All data are obtained from DataStream.

We use the traditional Augmented Dickey Fuller (ADF) (Dickey-Fuller 1979) and Philips Perron (PP) (1988) unit root tests to examine the time series properties of the variables. We also use a breakpoint unit root test, since Perron (1989) indicates that structural change and unit roots are closely related, and conventional unit root tests can be biased toward a false unit root null when the data are trend stationary with a structural break.

In the literature, Phillips and Hansen’s (1990), Johansen and Juselius’ (1990), and Engle and Granger’s (1987) cointegration tests are widely used to determine the cointegration relationship. However, it is impossible to estimate structural breaks with these methods, and these tests require that the order of integration of the variables should be unique. Therefore, we use the ARDL approach, which is more flexible when analysing the effect of the minimum wage on unemployment, growth, and prices. The ARDL approach can be used when the variables are stationary at I(1) or I(0) or are mixed. However, if the series are stationary at the second difference, I(2), this method is inapplicable.

In the analysis all the variables are used in logarithmic form. Figure 1 shows the time series plots of the variables.
As Figure 1 indicates that the industrial production index exhibits strong seasonality, the TRAMO/SEATS method is used to remove the seasonality in the series.

If all the variables guarantee the condition that the series are stationary at I(1) or I(0) or are mixed, the ARDL bounds test approach can be used to analyse cointegration between the variables. This method provides super consistent results for a small sample and does not suffer from an endogeneity problem. An ARDL model that shows long- and short-run relationships for the selected variables can be written as follow:

\[
\begin{align*}
\Delta \text{LPPI}_t &= \beta_{11} + \sum_{i=0}^{n} \alpha_{1i} \Delta \text{LMW}_{t-i} + \sum_{i=0}^{m} \alpha_{2i} \Delta \text{LUNEMP}_{t-i} + \\
&\quad \sum_{i=0}^{p} \alpha_{3i} \Delta \text{LPI}_{t-i} + \sum_{i=1}^{q} \alpha_{4i} \Delta \text{LPPI}_{t-i} + \theta_{11} \text{LMW}_{t-1} + \theta_{21} \text{LUNEMP}_{t-1} + \\
&\quad \theta_{31} \text{LPI}_{t-1} + \theta_{41} \text{LPPI}_{t-1} + \epsilon_t
\end{align*}
\]

\[
\begin{align*}
\Delta \text{LUNEMP}_t &= \beta_{12} + \sum_{i=0}^{s} \alpha_{1i} \Delta \text{LMW}_{t-i} + \sum_{i=0}^{u} \alpha_{2i} \Delta \text{LPI}_{t-i} + \\
&\quad \sum_{i=0}^{v} \alpha_{3i} \Delta \text{LPPI}_{t-i} + \sum_{i=1}^{y} \alpha_{4i} \Delta \text{LUNEMP}_{t-i} + \theta_{12} \text{LMW}_{t-1} + \theta_{22} \text{LPI}_{t-1} + \\
&\quad \theta_{32} \text{LPPI}_{t-1} + \theta_{42} \text{LUNEMP}_{t-1} + \epsilon_t
\end{align*}
\]
\[ \Delta LIPI_{SA_t} = \beta_{11} + \sum_{i=0}^{z} \alpha_{1i} \Delta LW_M_{t-i} + \sum_{i=0}^{a} \alpha_{2i} \Delta LPPI_{t-i} + \sum_{i=0}^{b} \alpha_{3i} \Delta LUNEMP_{t-i} + \sum_{i=1}^{c} \alpha_{4i} \Delta LIPI_{SA_t-i} + \theta_{13} LW_M_{t-1} + \theta_{23} LPPI_{t-1} + \theta_{33} LUNEMP_{t-1} + \theta_{43} LIPI_{SA_{t-1}} + u_{3t} \quad (3) \]

In this notation, \( \beta_{11}, \beta_{12} \) and \( \beta_{13} \) are the drift components and \( u_t \)s are white noise error terms. The term with a summation sign represents the error correction dynamics and the second part of the equation corresponds to the long-run relationship.

In order to test the existence of a long-run relationship between the variables, all the equations are estimated by ordinary least squares (OLS), and then an F-test examines the joint significance of the coefficients of the lagged levels of the variable. Thus, the null hypothesis of no co-integration among the variables in Equation 1 is:

\[ H_0: \theta_{11} = \theta_{21} = \theta_{31} = \theta_{41} = 0 \quad \text{against the alternative hypothesis of cointegration} \quad H_1: \text{at least one of them is different than zero} \]

For Equation 2:

\[ H_0: \theta_{12} = \theta_{22} = \theta_{32} = \theta_{42} = 0 \quad \text{against the alternative hypothesis of cointegration} \quad H_1: \text{at least one of them is different than zero} \]

For Equation 3:

\[ H_0: \theta_{13} = \theta_{23} = \theta_{33} = \theta_{43} = 0 \quad \text{against the alternative hypothesis of cointegration} \quad H_1: \text{at least one of them is different than zero} \]

The ARDL F-statistic should then be compared with the upper bound and lower bound critical values generated by Peseran et al. (2001). If the ARDL F-statistic is lower than the critical bound, it means there is no cointegration relationship between variables. If the ARDL F-statistic is bigger than the critical bound it means there is a cointegration relationship between variables. If the F-statistic remains in between the critical bounds the result changes depending on whether the series are stationary at their level or at first difference. By observing
cointegration between the variables we can examine the short-run and long-run relationships.

By adopting the ARDL approach the short- and long-run dynamic relationships can be estimated. Therefore, equations (1,2,3) can be rewritten as the error correction version of the ARDL model. Since the existence of a cointegration relationship requires at least one-way direction causality between the variables, a Granger causality test should be conducted. However, instead of standard Granger causality it is more suitable to use the ECM-based Granger causality first used by Sargan (1964), developed by Davidson et al. (1978), and popularised after Engle & Granger (1987). Granger (1969) also indicates that the vector error correction method (VECM) is more appropriate to examine the causality between the series if the variables are integrated at I(1).

The VECM is a restricted form of unrestricted VAR (vector autoregression) and this system uses all the series endogenously. Moreover, it allows the predicted variable to explain itself both by its own lags and by the lags of other variables, as well as by the error correction term and the residual term (Shahbaz et al. 2013: 114). With this method, both the short-run and long-run causality can be observed. Therefore, in this study the Granger causality test with the ECM framework is preferred. In addition to the first three models, a fourth model is formed to see the causality in bi-directional form. The model is as follows:

\[
\Delta LPPI_t = \lambda_{11} + \sum_{i=0}^{n} \varphi_{1i} \Delta LMW_{t-i} + \sum_{i=0}^{m} \varphi_{2i} \Delta LUNEMP_{t-i} + \sum_{i=0}^{p} \varphi_{3i} \Delta IPI_{SA_{t-i}} + \sum_{i=1}^{q} \varphi_{4i} \Delta LPPI_{t-i} + \gamma_1 ecm_{t-1} + \epsilon_t 
\] (4)

\[
\Delta LUNEMP_t = \lambda_{12} + \sum_{i=0}^{s} \vartheta_{1i} \Delta LMW_{t-i} + \sum_{i=0}^{u} \vartheta_{2i} \Delta IPI_{SA_{t-i}} + \sum_{i=0}^{v} \vartheta_{3i} \Delta LPPI_{t-i} + \sum_{i=1}^{y} \vartheta_{4i} \Delta LUNEMP_{t-i} + \gamma_2 ecm_{t-1} + \epsilon_2 
\] (5)

\[
\Delta IPI_{SA_{t}} = \lambda_{13} + \sum_{i=0}^{z} \mu_{1i} \Delta LMW_{t-i} + \sum_{i=0}^{q} \mu_{2i} \Delta LPPI_{t-i} + \sum_{i=1}^{c} \mu_{3i} \Delta LUNEMP_{t-i} + \gamma_3 ecm_{t-1} + \epsilon_3 
\] (6)

\[
\Delta LMW_t = \lambda_{14} + \sum_{i=0}^{d} \vartheta_{1i} \Delta IPI_{SA_{t-i}} + \sum_{i=0}^{e} \vartheta_{2i} \Delta LUNEMP_{t-i} + \sum_{i=0}^{f} \vartheta_{3i} \Delta LPPI_{t-i} + \sum_{i=1}^{g} \vartheta_{4i} \Delta LMW_{t-i} + \gamma_4 ecm_{t-1} + \epsilon_4 
\] (7)
In this notation, Δ is the first difference operator. Residuals (e_{1t}, e_{2t}, e_{3t}, e_{4t}) are assumed to be normally distributed and white noise. ecmt_{t−1} is the error correction term which has to be negative and statistically significant and shows the speed of adjustment to long-run equilibrium following a short-run shock. It is derived from the cointegration equation. For the first equation ecmt equals γ_{1}ecmt_{t−1} = LPPI_{t−1} + (β_{11}/θ_{41}) + (θ_{11}/θ_{41})LMW_{t−1} + (θ_{21}/θ_{41})LUNEMP_{t−1} + (θ_{31}/θ_{41})LIPI_{SA}_{t−1}. For the short-run Granger causality the following hypotheses are tested:

\[ H_{01}: \varphi_{1t} = 0 \] implying that ΔLMW does not Granger cause ΔLPPI

\[ H_{02}: \vartheta_{1t} = 0 \] implying that ΔLMW does not Granger cause ΔLUNEMP

\[ H_{03}: \mu_{1t} = 0 \] implying that ΔLMW does not Granger cause ΔLIPI_{SA}

and so on for other variables.

5. EMPRICAL RESULTS

Initially, the series’ stationary level is tested by conducting ADF, PP, and breakpoint unit root tests. Table 1 shows the results.

<table>
<thead>
<tr>
<th>Variables</th>
<th>ADF</th>
<th>PP</th>
<th>Breakpoint Unit Root</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Intercept</td>
<td>Trend and Intercept</td>
<td>Intercept</td>
</tr>
<tr>
<td>LIPI_SA</td>
<td>−1.217642 (0.6659)</td>
<td>−3.195499 (0.0896)</td>
<td>−1.071794 (0.7260)</td>
</tr>
<tr>
<td>∆LIPI_SA</td>
<td>−5.231818 (0.0000)</td>
<td>−12.39019 (0.0000)</td>
<td>−12.3571 (0.0000)</td>
</tr>
<tr>
<td>LMW</td>
<td>0.846486 (0.9945)</td>
<td>−2.345583 (0.4064)</td>
<td>1.993999 (0.9999)</td>
</tr>
<tr>
<td>∆LMW</td>
<td>−13.26172 (0.0000)</td>
<td>−13.30871 (0.0000)</td>
<td>−14.92676 (0.0000)</td>
</tr>
<tr>
<td>LUNEMP</td>
<td>−1.988269 (0.2918)</td>
<td>−2.023381 (0.5832)</td>
<td>−1.633948 (0.4627)</td>
</tr>
<tr>
<td>∆LUNEMP</td>
<td>−3.995006 (0.0019)</td>
<td>−3.988310 (0.0112)</td>
<td>−9.144698 (0.0000)</td>
</tr>
</tbody>
</table>
Results show that all the variables are stationary at first difference. Therefore, conducting the autoregressive distributed lag (ARDL) model established by Peseran et al. (2001) is more suitable for this data set. This methodology provides unbiased results. First, the existence of cointegration among variables is established with the bounds test. The test results are shown in Table 2.

Table 2: Results of ARDL cointegration test

<table>
<thead>
<tr>
<th>Series Model</th>
<th>F-statistic</th>
<th>Significance level</th>
<th>Bound critical values</th>
<th>Decision</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \Delta LPPI_t )</td>
<td>11.00130</td>
<td>1%</td>
<td>4.29</td>
<td>5.61</td>
</tr>
<tr>
<td>( \Delta LUNRATE_t )</td>
<td>11.55253</td>
<td>1%</td>
<td>3.65</td>
<td>4.66</td>
</tr>
<tr>
<td>( \Delta LIPI-SA_t )</td>
<td>6.149843</td>
<td>1%</td>
<td>4.29</td>
<td>5.61</td>
</tr>
</tbody>
</table>

Since the variables are cointegrated in all the models, long-run coefficients are derived. Table 3 shows the models’ long-run coefficients and diagnostic test results.

Table 3: Long-run coefficients

<table>
<thead>
<tr>
<th>Models</th>
<th>( LPPI_{t-1} )</th>
<th>( LUNRATE_{t-1} )</th>
<th>( LIPI-SA_{t-1} )</th>
<th>( LMW_{t-1} )</th>
<th>JB [p-values]</th>
<th>( X^2_{LM} ) [p-values]</th>
<th>( X^2_{white} ) [p-values]</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \Delta LPPI_t )</td>
<td>-</td>
<td>0.1539***</td>
<td>0.7320***</td>
<td>0.4544**</td>
<td>3.2996 [0.1920]</td>
<td>0.4044 [0.8169]</td>
<td>37.688 [0.2251]</td>
</tr>
<tr>
<td>( \Delta LUNRATE_t )</td>
<td>1.7385**</td>
<td>-</td>
<td>-3.3225**</td>
<td>0.6342**</td>
<td>0.1302 [0.9369]</td>
<td>0.3962 [0.8203]</td>
<td>12.320 [0.7216]</td>
</tr>
<tr>
<td>( \Delta LIPI-SA_t )</td>
<td>0.9948***</td>
<td>-0.1978***</td>
<td>-</td>
<td>-0.0975</td>
<td>0.6528 [0.7215]</td>
<td>0.8983 [0.6382]</td>
<td>82.1039 [0.1527]</td>
</tr>
</tbody>
</table>

* Denotes statistically significant at 10%
** Denotes statistically significant at 5%
*** Denotes statistically significant at 1%
Notes: JB indicates Jarqua-Bera normality test, $X^2_{LM}$ indicates Breusch-Godfrey LM test, and $X^2_{white}$ indicates white test. Because of structural breaks a dummy for October 2015 is used in the first model, a dummy for May 2009 is used in the second model, and a dummy for April 2009 in the third model.

As shown, while the minimum wage has a statistically significant effect on prices and unemployment it does not have a statistically significant effect on production. A 1% increase in the minimum wage causes a 0.45% increase in prices. As expected, an increase in the minimum wage raises prices. Moreover, a 1% increase in the minimum wage causes an increase in unemployment of 0.63%. This is also as expected. Therefore, we can conclude that an increase in the minimum wage plays a significant role in the increase in prices and the number of unemployed people.

In terms of diagnostic tests, all the models are normally distributed, since the Jarque-Bera normality test accepts the null hypothesis that says error terms are normally distributed. Moreover, the Breusch-Godfrey LM test shows that the models do not suffer from a serial correlation problem. The white test also shows that the models do not have a heteroscedasticity problem.

The presence of cointegration between minimum wage, unemployment, prices, and industrial production requires at least one direction of Granger causality. The advantage of using the ECM-based Granger causality test is that it gives both short- and long-run causal relationships. Table 4 shows the short- and long-run Granger causality results.

Table 4 shows that in the short run the minimum wage only has statistically significant causality on prices. On the other hand, since one period lagged error correction term is statistically significant at the 1% level in all the equations, minimum wage, unemployment, prices, and production have bi-directional Granger causality in the long run. Moreover, the significance of the $\text{ecm}_{t-1}$ coefficients supports the bounds test results which show that the variables are moving together in the long run. Therefore, we can conclude that the obtained significant error correction terms indicate long-run causality for all variables.
Table 4: Results of causality test

<table>
<thead>
<tr>
<th>Dependent Variable</th>
<th>F-statistics</th>
<th>[p-values]</th>
<th>ecm_{t-1} [p-values]</th>
</tr>
</thead>
<tbody>
<tr>
<td>ΔLPPI_t</td>
<td>-</td>
<td>3.951084***</td>
<td>[0.0007]</td>
</tr>
<tr>
<td>ΔLUNRATE_t</td>
<td>2.471341**</td>
<td>2.698353***</td>
<td>[0.0096]</td>
</tr>
<tr>
<td>ΔLIPI_SA_t</td>
<td>1.467263</td>
<td>6.358800***</td>
<td>[0.0023]</td>
</tr>
<tr>
<td>ΔLMW_t</td>
<td>2.930235**</td>
<td>0.012802</td>
<td>[0.0360]</td>
</tr>
</tbody>
</table>

* indicates statistically significant at 10% level.
** indicates statistically significant at 5% level.
*** indicates statistically significant at 1% level

Notes: The selected lag lengths for ΔLPPI_t=[5,8,8,7], ΔLUNRATE_t=[4,4,0,4], ΔLIPI_SA_t=[3,4,1,2], and ΔLMW_t=[1,0,3,0]

6. CONCLUSION

The effect of the minimum wage on macroeconomic variables, especially employment, is a much-debated topic in the economics literature. Neoclassical theory says that an increase in the minimum wage forces sectors to raise wages and reduces employment. On the other hand, Keynesians claim that since low wage groups’ marginal propensity to consume is high, an increase in the minimum wage will increase total demand and contribute to economic growth and thus employment. While some economists in the empirical literature (Reynolds and Gregory 1965; Brown, Gilroy and Kohen 1982; Castillo and Freeman 1992; Deere, Murphy and Welch 1995; Neumark and Wascher 1992, 2008) find that an increase in the minimum wage has a negative effect on employment, other economists (Card 1992a, 1992b; Katz and Krueger 1992; Card and Krueger 1994; Bernstein and Schmitt 1998) find that a minimum wage increase either does not have a statistically significant effect on employment or has a positive effect.

Because the theory is inconclusive, our aim is to investigate the effect of the minimum wage on employment, prices, and growth for Turkey, which recently experienced a significant minimum wage increase. The data used is monthly and covers the period from January 2005 to March 2017. The conducted unit root
tests showed that the series are stationary at their first difference. Therefore, we used the ARDL approach, which provides super consistent results in a small sample and does not suffer from endogeneity. After finding cointegration among variables we obtained long-run coefficients. The results show that while the minimum wage has a statistically significant effect on prices and unemployment, it does not have a statistically significant effect on production. A 1% increase in the minimum wage causes a 0.45% increase in prices and a 0.63% increase in unemployment. Since the presence of cointegration between the variables requires at least one-way Granger causality, ECM-based Granger causality analyses was used to provide both short- and long-run causality results. The obtained results show that in the short run the minimum wage only has a statistically significant causality on prices. On the other hand, since in all the equations one period lagged error correction term is statistically significant at the 1% level, minimum wage, unemployment, prices, and production have bi-directional Granger causality in the long run. Consequently, looking at the test results it can be concluded that an increase in the minimum wage plays a significant role in increasing prices and unemployment in Turkey. Therefore, when deciding minimum wage increases, politicians should not succumb to popular demand and should carefully consider the results of their actions.

REFERENCES


Received: March 14, 2019
Accepted: August 29, 2019