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Abstract

The time-series approach used in the minimum wage literature essentially aims to estimate a treatment effect of increasing the minimum wage. In this paper, we employ a novel approach based on aggregate time-series data that allows us to determine if minimum wage changes have significant effects on employment. This involves the use of tests for structural breaks as a device for identifying discontinuities in the data which potentially represent treatment effects. In an application based on Australian data, the tentative conclusion is that the introduction of minimum wage legislation in Australia in 1997 and subsequent minimum wage increases appear not to have had any significant negative employment effects for teenagers.

1. Introduction

Until recently, an evaluation of the impact of minimum wage legislation on the labour market employment has traditionally been based on time-series studies. A widely cited paper is the review by Brown, Gilroy and Kohen (1982), who in their survey of time-series studies up to 1981, found a reduction of between one and three percent in teenage employment as a result of a 10 percent increase in the US federal minimum wage. This estimate used to be regarded as the "consensus" estimate and cited in debates surrounding minimum wages around the developed world.

Following the increasing availability of individual, firm, industry and state level data sets, however, time series approaches appear to have abruptly fallen out of favour. This alternative micro-level approach to analyzing the impact of minimum wages has been termed the "new minimum wage research." This body of research approaches the issue from new directions and can be broadly grouped into two categories: panel data studies that employ state-specific data over time, and case studies that focus on the effects of minimum wage changes in specific states. Neumark and Wascher (2007) provide an extensive review of this recent literature. According to this new line of research, there is a wide range of existing estimates and a lack of consensus about the overall effects on low-wage employment of an increase in the minimum wage. Some researchers (e.g., Card and Krueger, 1995) have even found in that in certain industries, employment may actually have increased in response to an increase in the minimum wage.

This paper contributes to the literature on minimum wages by reconsidering the use of a time-series approach to determine the relationship between minimum wage legislation and employment. In contrast to the cross-sectional and short panel data sets that are often now used in the new minimum wage research, time-series analysis is appealing because it is able to provide feedback on longer run impacts of minimum wage changes. As the objective of the time-series approach used in the minimum wage literature is essentially in estimating a treatment effect, we employ a novel approach based on aggregate time-series data that allows us to determine if minimum wage changes have significant effects on employment. This involves the use of tests for structural breaks as a device for identifying discontinuities in the data which potentially represent treatment effects. As shown in Piehl et al. (2003), such tests for unknown structural breaks provide a useful framework for estimating the parameters of interest in an evaluation framework. Such tests can be used to test for the existence of a break in the data series and pinpoint the timing of the break. These results can then be used to calculate the magnitude of the impact. Although Piehl et al. (2003) focus on testing for a single structural break as a result of the implementation of a single program, we extend their approach to the case where there are multiple treatments (i.e., multiple discrete increases in the minimum wage) by adopting tests for a single and multiple structural breaks. Allowing for a multiple sequence of treatments is of practical relevance in many contexts because it sometimes takes several exposures to treatment before there are any discernible effects.

In this paper, we apply this technique by analyzing quarterly time-series data on teenage employment in Australia for the period 1992 Q1 to 2008 Q1. For the case of Australia, 1997 is an important year because it was the year that the Australian Industrial Relations Commission (AIRC) established a federal minimum wage, with "[t]he main reason for so deciding is to give effect to the statutory requirement to have regard, when adjusting the safety net, to the needs of the low paid."[1] We focus on eleven federal minimum wage increases from when it was introduced in April 1997 to June 2007 in the states of Victoria, the Australian Capital Territory and the Northern Territory. We focus on these three states because during the period 1997 to 2007, only these states had all employees under federal industrial jurisdiction and subject to a binding federal minimum wage.

In the context of minimum wage legislation, the use of tests for *unknown* structural breaks is useful even though the dates when minimum wage changes came into effect are known for three reasons. First, there could be anticipation effects that complicate the identification of the precise timing of an effect. Second, states might not have necessarily immediately enforced changes in the federal minimum wage, which would give rise to lagged effects. Third, even if states immediately implemented such minimum wage changes, employers might not react immediately. For example, an increase in the minimum wage would probably first affect the hiring of new workers and not necessarily the firing of existing workers. This would give rise to effects that are spread out over time, which also make attributing the effects of minimum wage changes to a precise point in time difficult.

[1] *Safety Net Review – Wages – April 1997*, section 8.2.4.

For these reasons, under the traditional time-series approach, any dummy variables that are included to represent minimum wage changes may not enter at the right time for evaluating their effects, giving rise to inaccurate estimated effects.

There exist few studies of the effects of minimum wages in Australia. Leigh (2003, 2004) used a difference-in-difference approach to estimate the elasticity of labour demand with respect to the Western Australian minimum wage and found an elasticity of -0.29. Based on a survey of small and medium sized businesses, Harding and Harding (2004) estimated that the short run elasticity of labour demand with respect to the minimum wage was -0.21. Given that the operation of minimum wages in Australia is complex, it is not surprising that the findings in both these studies have been subject to much criticism and debate. For example, in his critique of Leigh's (2003) paper, Watson (2004) notes that despite Leigh's attempt to use a quasi-experimental design, proper natural experiments (on the relationship between minimum wages and employment) still remain to be done in Australia. Similarly, the *Safety Net Review – Wages – May 2004* highlights several weaknesses of the *Minimum Wages Report* by Harding and Harding (2004). These include: (1) the report is based on an extrapolation of the responses of just 37 firms who reported an adverse economic impact from the May 2003 safety net adjustment; (ii) there appear to be significant differences between a number of industry sector estimates extrapolated from the report questionnaire and those from established ABS surveys; and (iii) the response rate of the report survey was 20 to 22 percent.

Although the international literature on minimum wages is voluminous, differing economic conditions and contexts affecting minimum wages in various countries imply that those results might not be directly relevant to Australia. This is because Australia's minimum wage system prescribes not one minimum wage but a series of minimum wages at higher levels through the wages distribution; Australia's minimum wage is higher in relative terms; and because Australia's minimum wage is relatively high and likely to cover a higher proportion of employees than other countries.[2]

The remainder of this paper is organized as follows. Section 2 provides a brief overview of recent developments in time-series studies of minimum wages. Section 3

[2] For example, in May 2002, it was estimated that 23.2 percent of the workforce were covered by the minimum wage system (ABS, 'Employee Earnings and Hours, Australia, May 2002', Catalogue No.6306.0, p. 44, Table 23).

introduces the Australian data that we use for our empirical analysis. In section 4, we check for stationarity properties of the time-series data. Section 5 discusses the econometric model, the structural break tests we employ and their results. Section 6 discusses the results of a robustness test based on the traditional time-series approach, and a robustness check using a longer time-series. Finally, section 7 concludes.

2. Time-Series Approaches to Estimating Minimum Wage Effects

The early time-series studies reviewed in Brown, Gilroy and Kohen (1982) attempted to estimate the effect of minimum wages on the labour force of youth based on single equation models of the type:

$$EP = f(MW, D, X) \tag{1}$$

where *EP* is the teenage employment to population ratio, *MW* is a measure of the minimum wage, *D* a business-cycle variable, and *X* represents exogenous explanatory factors that control for labour supply effects. The relationship is assumed to be linear with all variables expressed in logarithms. Typically, the minimum wage variable used in the Kaitz index – this is defined as the minimum wage relative to the average wage weighted by the coverage of the minimum wage.

Subsequent studies by Solon (1985) and Wellington (1991) highlighted that there was substantial residual autocorrelation in many of these early studies, and suggested that one should include interactions between the quarterly seasonal dummies and a linear and quadratic trend along with modelling the error as a first-order autoregressive or AR(1) process:

$$EP = f(MW, D, X, T, T^2, S) \tag{2}$$

where *T* is the time trend and *S* represents seasonal dummies.

More recently, Park and Ratti (1998) and Williams and Mills (2001) point out that previous time-series studies of minimum wages did not account adequately for serial correlation and non-stationarity in the data, which result in inconsistent estimates of the effects of minimum wages on employment. Park and Ratti (1998) suggest applying an autoregressive conditional heteroscedasticity (ARCH) model on the transformed data (in order to achieve stationarity). As endogeneity of some of the *X* variables is also possible, Williams and Mills (2001) suggest using vector autoregressions. An alternative approach to circumvent the econometric issues based on the specification in Equation (2) was provided by Bazen and Marimoutou (2002). They suggest using a more flexible approach to the specification of various components of the basic time-series model, in which the trend, cyclical and seasonal components are treated as stochastic rather than deterministic.

The decline in the popularity of time-series approaches to analyzing the effects of minimum wages is partly due the influential book by Card and Krueger (1995), who express scepticism that variations of Equations (1) and (2) can be used to accurately measure the effects of changes in the minimum wage. Their criticism is based on the fact that it is difficult to choose the correct set of *X* variables, and that there are concerns over using the Kaitz index of the minimum wage as the main variable of interest. For example, a change in the coefficient to the Kaitz index could be due to a change in coverage or a change in average wages and not purely due to a change in the minimum wage. In addition, another factor that has led to the demise of time-series studies of minimum wages in the US is that given the recent proliferation of state minimum wages that are above the federal level, identification in time-series studies has become more problematic.[3]

Instead of using variations of Equations (1) or (2) that has the shortcoming of using the Kaitz index, this paper adopts the evaluation approach used in Piehl et al. (2003).[4] The idea involves modelling the dependent variable of interest using a parsimonious time-series model, and using a structural break test to determine if the timing of changes in policy coincides with statistically significant discontinuities in the data series of the dependent variable. In our context, as changes in minimum wage legislation involved several discrete changes, we extend the approach used in Piehl et al. (2003) to the case where there are sequential multiple treatments by adopting tests for a single break (Quandt, 1960) and multiple structural breaks (Bai, 1997).

[3] Note that this scenario in turn helps provide identification using the panel data and case study approaches in the US, explaining the proliferation of such "new" approaches in the literature.
[4] In any event, due to the lack of reliable coverage data in Australia, it is not possible to estimate time-series regressions based on the Kaitz index.

From an evaluation perspective, an important advantage of the structural break approach is that it can even be used in the case when there are no obvious or appropriate comparison groups, as is often the case in practice. In certain situations, even if comparison groups are available, such an approach might still be useful.

First, a difficulty is sometimes encountered in assigning a correct starting date of the intervention for the comparison group in order to facilitate calculation of average treatment effects. For example, in an evaluation setting where the treatment group is enrolled into a particular program over time and the comparison group consists of non-participants who might be enrolled into treatment at a later date, one proposed approach in the literature is to randomly draw start dates for the comparison group (e.g., Lechner, 1999). But this solution is not completely satisfactory – see Fredriksson and Johansson (2003) for a critique of this approach. As an alternative, a structural break approach would focus on specific cohorts of individuals entering a particular treatment and examining if their outcomes experience a structural break some time after the exposure to the treatment (i.e., allowing for an initial period of locking-in effects).

Second, for the case of analyzing a sequence of multiple treatments (e.g., see Lechner and Miquel, 2005), a quasi-experimental approach would require a very strong set of identifying assumptions. For example, Lechner (2006) states that if the assumptions underlying matching in a static context can be characterized as being data hungry, then the assumptions underlying matching in a causal sequences of interventions can be characterized as being starving for data because past intermediate outcomes will also need to be taken into account. On the other hand, a structural break approach based on a single break and/or multiple breaks can allow past intermediate outcomes to be taken into account using more parsimonious reduced-form models.

However, not having a comparison group to represent a plausible counterfactual clearly also results in certain limitations. The implications of not having a comparison group is that even when a break is identified, this does not constitute conclusive evidence that the break is solely due to the implementation of the program as many other factors could have occurred simultaneously. However, institutional knowledge can be useful in this case to aid in determining if such breaks are solely due to the effects of one policy change, or plausibly due to other exogenous shocks. In other words, if a large effect is found that coincides with the dates surrounding changes in minimum wage legislation, a competing explanation would need to be able to account for the sudden change in the employment to population ratio. On the other hand, if no breaks are found during the period when minimum wage changes took place, then this would be evidence in favour of there being no program effects. Although it is possible that other exogenous shocks might cancel out whatever positive or negative effects minimum wages might have on employment, it is probably quite unlikely that such coincidental cancelling out of effects occurs when an examination is made of a series of minimum wage changes.

3. Data

The Australian data used in this paper are time-series data for the period 1992 Q1 to 2008 Q1 and come from the Labour Force Survey conducted by the Australian Bureau of Statistics (ABS), a regular monthly survey of Australians aged 15 and over.[5]

In April 1997, the AIRC introduced a federal minimum wage of A$359.40 per week, with appropriate adjustments for junior, part-time and casual employees.[6] Given the standard 38 hour work week in Australia, this was equivalent to A$9.46 per hour. The setting of minimum wages by the AIRC was influenced heavily by the concept of a ‘living wage’ that can be traced back to the *Harvester* decision of 1907, where Justice Higgins expounded on the notion of a ‘fair and reasonable wage.’ For the period analyzed in this paper, employees in the states of the Australian Capital Territory and the Northern Territory were under complete federal jurisdiction and had their minimum wages set by the AIRC, while Victoria had a large majority of its employees subject to federal jurisdiction (for the purposes of analysis in this paper, we assume that all Victorian employees were also covered by federal awards). In the other five states in Australia, whether federal minimum wages applied to an employee depended on the employee’s industry and whether the

[5] The detailed data used in this paper were obtained from the ABS data cubes (Cat No. 6291.0.55.001). Monthly data from the Labour Force Survey were aggregated to quarterly data to be consistent with the majority of time-series studies examining the effects of minimum wages, which are based on quarterly data. Although data from 1978 Q2 are available, we use the shorter time series after the 1990-1991 recession because we are primarily interested in a possible break date around the time of the introduction of the minimum wage legislation in 1997, and subsequent break dates after further changes to minimum wage levels.

[6] Workers aged under 21 years were generally paid between 50 to 90 percent of the minimum, with the rate varying by occupation and industry.

employing company had operations in multiple states. These states, however, had their own state industrial tribunals which generally adopted the federal minimum wage changes after a brief lag.

It is worth keeping in mind that examining any effects of the introduction of the Australian federal minimum wage in 1997 is made relative to the system that was in place before that – one that comprised of a complex web of award wages for different occupational categories. The comparison is not relative to a labour market where there are no wage floors.

Data from the Labour Force Survey collected by the ABS has been used in the past to analyze the effects of minimum wages, but not from a time-series perspective. Leigh's (2003, 2004) difference-in-difference strategy was to compare employment in Western Australia with employment in other states before and after a rise in the Western Australian statutory minimum in order to estimate the elasticity of labour demand with respect to the minimum wage. One problem with using a difference-in-difference approach in the Australian context is that no state really represents a plausible counterfactual, as all states were either subject to state-level or federal-level minimum wage increases. Although we use the data from the same source as Leigh (augmented with more recent data), our strategy is completely different and is based on a structural break test in the aggregate employment to population ratio data series for the three states fully under federal jurisdiction. The exact months of these eleven federal minimum wage changes we examine and the corresponding percentage increases in the minimum wage (nominal and real) are given in Table 1.

The total nominal increase in minimum wages between 1997 and 2007 was 45.3%. In real terms, this was equivalent to a 10.5% increase. These are the eleven time points around which one might expect discontinuities in the employment to population ratio if changes in the federal minimum wage legislation have any impacts on employment via their employees covered by federal awards.

For the empirical work in this paper, we do not use seasonally adjusted employment to population ratios like the adjustment made by Leigh (2003, 2004) who used a simple rolling average formula to adjust for trends based over the past three years. Instead, we use the raw non-adjusted data for our analysis and account for seasonality by including appropriate controls in our time-series model.

Table 1: Australian Federal Adult Minimum Wages

Year	Date Came into Effect	Federal Minimum Wage (in nominal dollars)	Hourly Equivalent (in nominal dollars)	Nominal Percentage Increase from Previous Year[7]	Federal Minimum Wage (in 1997 dollars)	Real Percentage Increase from Previous Year
1997	22 Apr 1997	359.40	9.46	2.86	359.40	2.61
1998	29 Apr 1998	373.40	9.83	3.90	370.24	3.02
1999	29 Apr 1999	385.40	10.14	3.21	376.62	1.72
2000	1 May 2000	400.40	10.54	3.89	374.52	-0.56
2001	2 May 2001	413.40	10.88	3.25	370.45	-1.09
2002	9 May 2002	431.40	11.35	4.35	375.31	1.31
2003	6 May 2003	448.40	11.80	3.94	379.58	1.14
2004	5 May 2004	467.40	12.30	4.24	386.60	1.85
2005	7 June 2005	484.40	12.75	3.64	390.25	0.94
2006	1 Dec 2006	511.76	13.47	5.65	398.20	2.04
2007	1 Oct 2007	522.12	13.74	2.02	397.00	-0.30

Notes: From 2006 onwards, the Australian Fair Pay Commission took over the role of the Australian Industrial Relations Commission in setting minimum wage rates. Hourly equivalents are calculated based on a 38 hour work week. Real wages are deflated using the CPI for all of Australia.

3.1 Descriptives

Figure 1 shows time-series data over the period 1978 Q2 to 2008 Q1 for the employment to population ratios for the states of Victoria, the Australian Capital Territory and the Northern Territory for 15 to 19 year olds. The vertical line in Q2 1997 depicts the introduction of the federal minimum wage legislation in Australia. Similarly, Figure 2 graphs the part-time employment to population ratio for 15 to 19 year olds in the same three states.

Given that small employment changes occur from one month to the next almost all the time, Keenan (1995) as described the effort of isolating minimum wage effects as 'looking for a needle in a haystack.' Although 'eye-balling' the descriptive evidence does not suggest that there were any significant effects of increases in the minimum wage on the employment of these young workers, this can be difficult to see graphically given possible serial correlation, seasonality effects and time trends. Furthermore, no control variables are included. In the next few sections, we formalize the analysis using an econometric model that controls for such factors.

[7] The percentage in 1997 reflects a A$10 per week increase from the C14 classification rate in the Metal Industry Award, which the AIRC at the time of introducing the minimum wage viewed as an equivalent of the minimum wage.

The two outcomes we focus on are: (i) the teenage (ages 15-19) full-time equivalent employment to population ratio (where full-time equivalent employment involves aggregating full-time and part-time employment figures and counting each part-time employee as 20/40 of a full-time employee); and (ii) the teenage (ages 15-19) part-time employment to population ratio. The former is labelled as *fte* while the latter is denoted as *ptr*.

Following the lead of many papers in this literature, we choose to focus on teenage outcomes because it is likely that changes in minimum wages will have the most effect on this subgroup of the population. In addition, the following variables that have been commonly used in past studies are used in our empirical analysis. To proxy for overall labour demand and business cycle effects, we use the unemployment rate for males aged 25 to 54 (denoted as *unemp*). To proxy for labour supply, we use the population of teenagers aged 15 to 19 as a proportion of the total working force population (referred as *tpop*).

Figure 1: Full-Time Equivalent Employment to Population Ratios for Teenagers

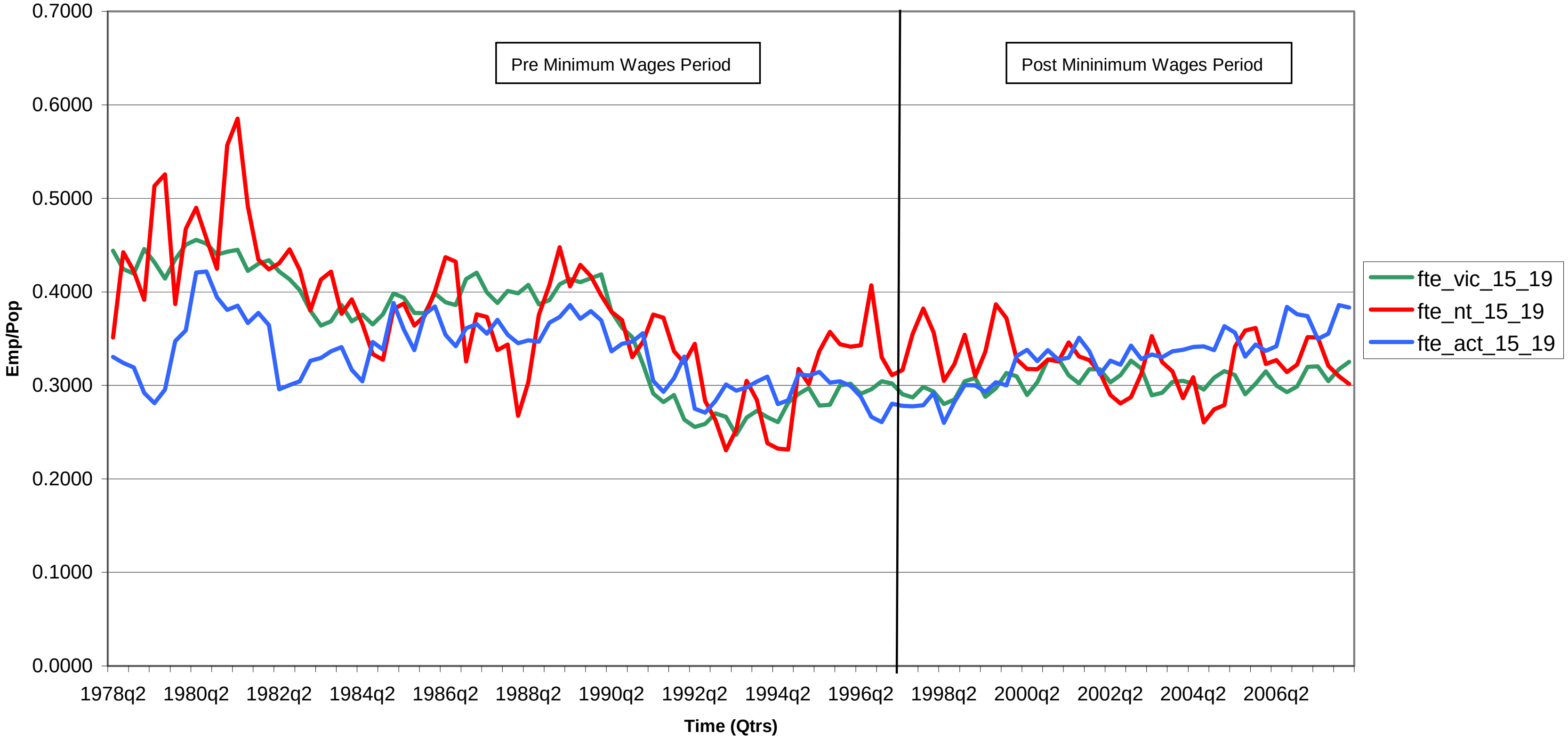
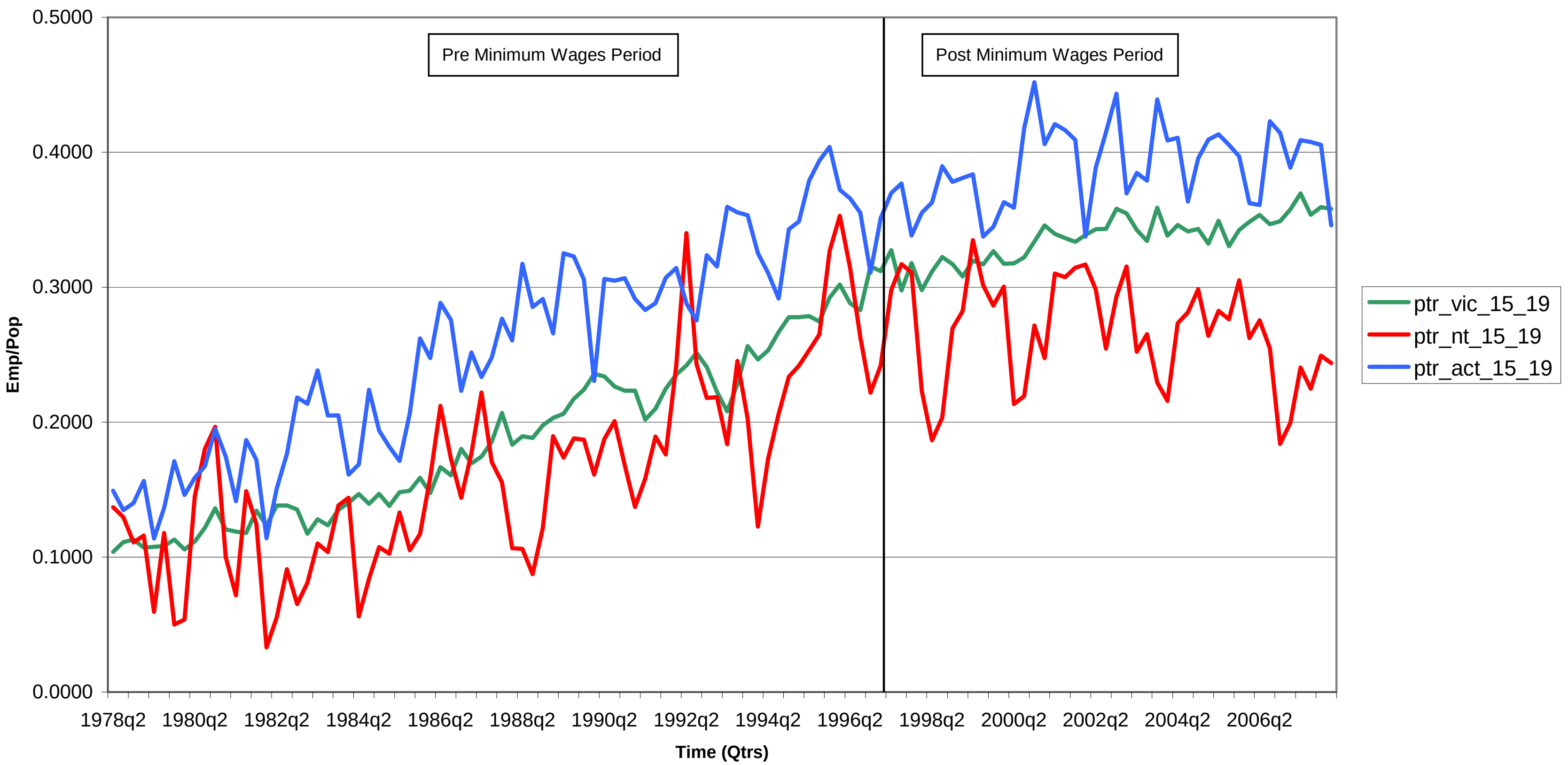


Figure 2: Part-Time Employment to Population Ratios for Teenagers



3.2 Who gets affected by changes in the minimum wage?

The impact of minimum wage increases is challenging to estimate because typically only a small proportion of the workforce is subject to the minimum wage (including the teenage workforce). As a result, increasing the minimum wage usually has a very small impact on average wages and on total employment. In other words, any estimates of the effects of minimum wages based on total employment would include wages of employees who are not affected by increases in the minimum wage. As a result, any estimated effects would be smaller than it would be if impacts on those directly affected could be isolated.

Defining 'minimum wage workers' as those earning between 100 percent and 120 percent of the federal minimum wage, and 'subminimum wage workers' as those whose hourly wages are below the federal minimum wage, Leigh (2007) estimates that over the period 1994-2002, there were approximately 10-12 percent of the labour force in each group. These figures are important because they are related to the use of the employment to population ratio as our dependent variable and the so-called 'fallacy of the inflated denominator' (Brown, 1988, p. 144). Given that these estimates are for the entire working population and that it is likely relatively more teenagers earn wages closer to the minimum wage, these likely represent lower bound estimates of the proportion of teenagers for whom minimum wages "bite." One possible adjustment would be to re-weight the employment impacts of a minimum wage change by the inverse of the proportion of employees who are actually affected by an increase in that minimum wage (e.g., see James, Wooden and Dawkins, 2001). Such adjustments can be helpful in making the results of minimum wage studies comparable to those that focus on wage elasticities. This will avoid understating the impact of minimum wages on the employment of those whose wages will be affected by such an increase.

Studying the effects of changes in minimum wages in Australia is complicated by the fact that it is not only the wage floor that moves, but also the whole pay scale for employees under federal jurisdiction. Put another way, when minimum wage changes take place, employees covered by award agreements who are paid above the minimum wage also get an increase in wages because changes are made to the entire pay and classification scale that includes a number of other skill levels. For example, precisely one year following the introduction of minimum wages, the April 1998 safety net decision raised the federal minimum wage by A$14 per week and increased wages by the same amount for award wages up to $550 per week. There was also an increase of A$12 in rates between A$550 and A$700 per week, and A$10 per week above A$700.[8] As a result, the eleven policy changes examined – federal changes to minimum wages in Australia over the period 1997 to 2007 – might be more properly viewed as a wider encompassing change to the structure of classification rates. Such changes are often referred to as 'safety net adjustments' by the Australian government. It is important to appreciate that a significant proportion of the Australian workforce relies on such safety net adjustments for increases in pay.

Given the complex range of factors affecting employment, it is a challenge to draw specific conclusions on the impact of safety net adjustments on employment. But given limited design options for an econometric study due to the lack of a comparison group, the structural break approach lets the data speak out and can potentially identify any large impacts due to the structural policy change to wage structures introduced by the Australian government in 1997.

4. Stationarity of the Data

As is typical of any time series analysis, an important first step of the modelling exercise is to determine the stationarity property of the series. An assessment of the unit root property of the data series is accomplished by employing the Augmented Dickey Fuller test (1984) (ADF) and the Zivot and Andrews (1992) (ZA) test that allows an endogenously determined breakpoint in the intercept, and in both the intercept and trend. As argued by Perron (1989), failing to account for a structural break in the conventional unit root test may lead to a loss of power and wrongly infer the presence of a unit root when in fact the series is stationary around a one time structural break. Given that minimum wage changes could effect the employment to population ratio, it is deemed essential to allow for a possible regime shift in the *EP* series comprising *fte* and *ptr*.

In its general form with breaks in both the intercept and the trend function, the test involves running the following regression for all potential breakpoints, T_B $(1 < T_B < T)$,

[8] See *Safety Net Review – Wages – April 1998*.

$$\Delta y_t = \mu + \beta t + \theta_1 DU_t + \gamma_1 DT_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + \varepsilon_t \qquad (3)$$

where DU_t and DT_t are break dummy variables that are defined as

$$DU_t = \begin{cases} 1 & if \quad t > T_B \\ 0 & \text{otherwise} \end{cases}$$

and

$$DT_t = \begin{cases} t - T_B & if \quad t > T_B \\ 0 & \text{otherwise} \end{cases}$$

where T is the sample size and k is the number of lags determined for each possible breakpoint by the Bayesian Information Criteria. Equation (3) is sequentially estimated and T_B is chosen so as to minimize the one-sided t-statistics of the unit root null hypothesis with no break (i.e. $H_0 : \alpha = 0$).

It is common to exclude the end-points of the sample when implementing the ZA unit root tests. This is due to the fact that the asymptotic distribution of their test statistics diverges to infinity when the end points are included. We report the results for 'trimming region' of the sample as suggested by Zivot and Andrews (1992) that is (0.15T, 0.85T). We also consider other trimming factors like 10% and 5%. Although not reported here, the results are largely consistent with those reported in Table 2.

Critical values at conventional levels of significance for the unit root tests are obtained from Zivot and Andrews (1992). For consistency with the reporting of the ADF tests results, only the results of the ZA test for a structural break in the intercept, and a break in both the intercept and trend are reported.[9]

[9] We also performed the test for a break in the trend and the results, although not reported here, are consistent with the finding in Table 1. Results for this set of test are available from the authors upon request.

Table 2: Unit Root Tests Results for Structural Break Models

	fte	*ptr*	*tpop*	*unemp*
Victoria				
ADF test				
(a)	-2.636 [5]	-1.738 [3]	-1.986 [2]	-0.344 [3]
(b)	-2.831 [5]	-1.359 [3]	-2.686 [2]	-4.773* [0]
ZA test				
(a)	-5.140 [9]	-2.361 [3]		
	{2003:1}	{1995:1}		
(b)	-4.652 [9]	-4.709 [3]		
	{2000:3}	{1996:4}		
Northern Territory				
ADF test				
(a)	-3.551* [0]	-4.131* [1]	-0.275*** [0]	-4.332*[0]
(b)	-3.632** [0]	-4.306* [1]	-1.309 [0]	-4.554*[0]
ZA test				
(a)	-5.505* [0]	-5.592* [1]		
	{1994:4}	{1995:2}		
(b)	-5.951* [0]	-5.865* [1]		
	{1994:4}	{1994:4}		
Australian Capital Territory				
ADF test				
(a)	-1.789 [0]	-3.553* [0]	-3.522** [2]	0.471 [3]
(b)	-3.988** [0]	-4.435* [0]	-3.513** [1]	-3.604**[0]
ZA test				
(a)	-5.458* [0]	-4.961 [4]		
	{1996:2}	{2005:4}		
(b)	-5.533** [0]	-5.780* [4]		
	{1996:2}	{2000:3}		

Notes: ADF = Augmented Dickey-Fuller Test. ZA = Zivot and Andrews test. (a) and (b) denote tests are conducted with intercept, and with both trend and intercept respectively. Figures in [.] represent the AIC-selected lag length. *,** and *** denote significance at 1%, 5% and 10% levels. The ZA test has a null hypothesis of a unit root with no break, and an alternative hypothesis of stationarity with a single break. The figure in {} under ZA test denotes break date. The 1%, 5% and 10% critical values for the ZA test with break(s) in intercept (intercept and trend) are -5.34 (-5.57), -4.80 (-5.08), and -4.58 (-4.82) respectively. The sample period for the unit root tests is for 1992 Q1 to 2008 Q1.

It can be seen from Table 2 that the ADF test results for Victoria indicate that all series, apart from *unemp,* are non-stationary. Although the ADF results for the auxiliary regressions with (a) intercept and (b) with intercept and trend yield different results about the stationarity property of the *ptr* and *unemp* series, we are inclined to accept the results of the latter as the plots of the data reveal the existence of a trend. In the case of Northern

Territory, the unit root test results suggest that *fte*, *ptr* and *unemp* are stationary. Finally, in the case of ACT we find that all series are stationary.

The ZA tests results for *fte* and *ptr* in Victoria fail to identify any structural break in either the intercept or both in the intercept and trend.[10] The tests fail to reject the null of a unit root with no break in the underlying process. The results of the ADF and ZA tests for *ptr* both point to the same result, suggesting non-stationarity of the series. For the Northern Territory, the ZA tests reject the null of a unit root with no break in both *fte* and *ptr*, and they identify the presence of a break in the intercept and trend function in 1994 Q4. The break date, however, falls outside of the minimum wage changes that occurred between 1997 Q2 and 2007 Q4. As for *fte* and *ptr* in the ACT, the ZA tests reject the null in favour of stationarity with breaks. The identified break dates are 1996 Q2 for *fte* and 2000 Q3 for *ptr*. The latter break date could emanate from changes in the minimum wage rates that took place around that time.

On the basis of our unit root test results, we take the first difference of the series whenever we find that the series is non-stationary. Because all series are expressed in logarithms, the first difference of a series can be interpreted as the growth rate of the variable concerned.

5. Methods

In practice, to apply the unknown structural break point technique in a program evaluation setting, one first needs to define the regression relationship of interest. In other words, we need to have the correct specification of the regression model under the null hypothesis of no break. Even though pre/post (or before/after) analyses of time series data appear to be intuitive, a scientifically valid evaluation requires more than testing the difference in a simple time series. It is important that the regression model is correctly specified, eliminating any possible trend effects that a simple pre/post comparison would pick up and erroneously identify as a treatment effect. For example, if a linear trend belongs in the model, then we would need to include it in the analysis in order to have the correct inference on the break in mean. However, if there is no trend in the true relationship

[10] We do not conduct the ZA test for *tpop* and *unemp* because we are not interested in whether there are breaks in their series.

and we include a trend variable, this will obscure inference on the break in mean as inclusion of a trend could absorb some of the change in mean.

Equation (4) is the base time-series regression model we use to model the employment to population ratio and to check for possible structural breaks, with data spanning from 1992 Q1 to 2008 Q1. All data are expressed in natural logarithms.

$$EP = f(D, X, T, T^2, S, ST, ST^2) \qquad (4)$$

In modelling *EP* we performed the Ljung-Box (1978) test to ensure that the residuals and squared residuals from the regressions are free from serial correlation.[11] In cases where serial correlation in the residuals is identified, we include an appropriate number of lagged dependent variables to purge the problem. Optimality of the number of lagged autoregressive variable is further confirmed using the Akaike Information Criteria (AIC). We find that for Victoria no lagged dependent variable is required in the model specification for both *ptr* and *fte*. In the case of the Northern Territory, two lagged dependent variables are included as regressors. Finally, for the ACT, one (two) lagged dependent variable(s) is (are) included for the *fte* (*ptr*) specification.

5.1 Structural break tests results

In the context of evaluation, a finding of a structural break that coincides with the implementation of changes in minimum wages can be interpreted as evidence supporting an effect of minimum wages. On the other hand, a finding of no structural break or a structural break at an alternative date would be evidence against there being an effect of minimum wages.

Chow (1960) proposed an *F*-test for a one-time structural change in one or more estimated regression coefficients when the date of the break is known. In the case of the model in equation (4), the null hypothesis is

$$\begin{aligned} EP_t = {} & \mu + \alpha_2 Q2 + \alpha_3 Q3 + \alpha_4 Q4 + \beta_1 \cdot T + \beta_2 \cdot T^2 + \gamma_2 Q2 \cdot T + \gamma_3 Q3 \cdot T + \gamma_4 Q4 \cdot T + \\ & \delta_2 Q2 \cdot T^2 + \delta_3 Q3 \cdot T^2 + \delta_4 Q4 \cdot T^2 + \lambda_1 tpop + \lambda_2 unemp + \varepsilon_t \end{aligned}$$

[11] Because we employ quarterly data, it is reasonable to consider up to order four for tests of serial correlation.

and the alternative hypothesis is

$$EP_t = \mu^{/} + \alpha_2^{/} Q2 + \alpha_3^{/} Q3 + \alpha_4^{/} Q4 + \beta_1^{/} \cdot T + \beta_2^{/} \cdot T^2 + \gamma_2^{/} Q2 \cdot T + \gamma_3^{/} Q3 \cdot T + \gamma_4^{/} Q4 \cdot T + \\ \delta_2^{/} Q2 \cdot T^2 + \delta_3^{/} Q3 \cdot T^2 + \delta_4^{/} Q4 \cdot T^2 + \lambda_1^{/} tpop + \lambda_2^{/} unemp + \nu_t$$

where the parameters marked with a prime ($^{/}$) are different from their corresponding ones without a prime. The Chow test statistic for a particular break date involves splitting the sample at that break date and estimating the model parameters separately on each sub-sample, as well as for the whole sample. The respective residual sum of squares (RSS) are computed and used to calculate the Wald statistic as follows

$$W = \frac{RSS - (RSS_1 + RSS_2)/k}{(RSS_1 + RSS_2)/T - 2k} \tag{5}$$

where RSS is the residual sum of squares for the whole sample, and the subscripts 1 and 2 denote the first and second sub-samples. T is the number of observations and k is the number of regressors in the sub-sample regression. Thus the test is one of how much the RSS for the whole sample is bigger than the sum of the RSS for the two sub-samples. If the coefficients do not change much between the samples, the RSS will not rise much upon imposing the constancy parameter restriction across the two sub-samples. However, in practice the date of the break is often not known *a priori* thus one would need to endogenously search for this structural change. The Chow (1960) test can be easily augmented to search for a break over all possible break dates. The test involves splitting the sample into two sub-periods over all possible break dates (τ) and estimating the parameters for each subperiod. A Wald statistic is then employed to test the equality of the two sets of parameters. In the presence of an unknown break date, the unidentified nuisance parameter implies that the *W*-test does not have a standard distribution. Andrews (1993) considers the distribution of this test statistic when the researcher searches over all possible values of τ . He proposed the test statistic

$$SupW = \max_{\tau} W$$

where $\pi \cdot T \leq \tau \leq (1-\pi)\cdot T$ and π is referred to the "trim factor." Andrews (1993) shows that this statistic converges to a non-standard distribution under very general conditions and provide tabulated asymptotic critical values. Like for the ZA test, one decision that needs to be made when applying these structural break tests is the choice of the "trimming" value. When one searches over all possible locations for a break in some parameters, it is important to specify how far into the sample one starts looking for a break and how close to the end of the sample one stops looking. The reason for not looking from the first observation to the last is that there must be a sufficient number of observations on either side of the point under consideration to estimate the regression relationship both before and after the break point.

The starting date of the sample for the structural break test is governed by the number of regressors that need to be estimated. Allowing for quarterly dummies, trend, squared trend and interactive terms between them as well as the lags of the dependent variable to correct for serial correlation in the residuals, a total of 20 observations are required from the start of the sample to the first candidate breakpoint. This implies that 1992 Q1 is a reasonable starting date if the test were to detect possible break(s) at the onset of the minimum wage in 1997 Q2.[12] Unfortunately, the need to estimate a significant number of regressors has the effect of reducing the ability to identify possible break dates towards the end of our sample period. As a result, only break dates to the end of 2003 can be identified, implying that the effects of minimum wage changes from 2004 to 2007 cannot be accounted for.[13]

Instead of reporting the *SupW* test statistic, we plot the sequence of the computed Chow statistics as a function of candidate break dates. Visual inspection of the plot would not only provide inference about the presence of a possible break in the underlying process but would also track the general trend of the *W* test statistic over the possible break dates. Figures 3 to 5 show plots of the *W* test statistic for all three states.

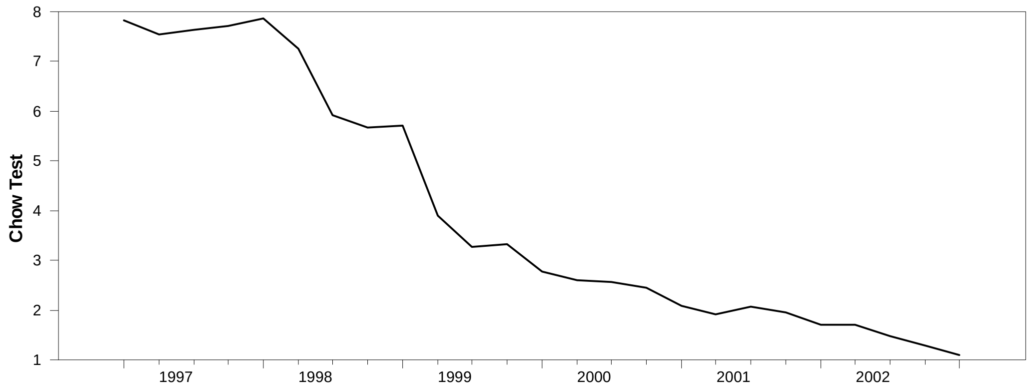
[12] We consider other starting dates for the sample involving a year and two years prior to 1992 Q1. The results are qualitatively unchanged.

[13] In Section 6, we use the model given in equation (2) to model the effects of minimum wages in the traditional way as a robustness test.

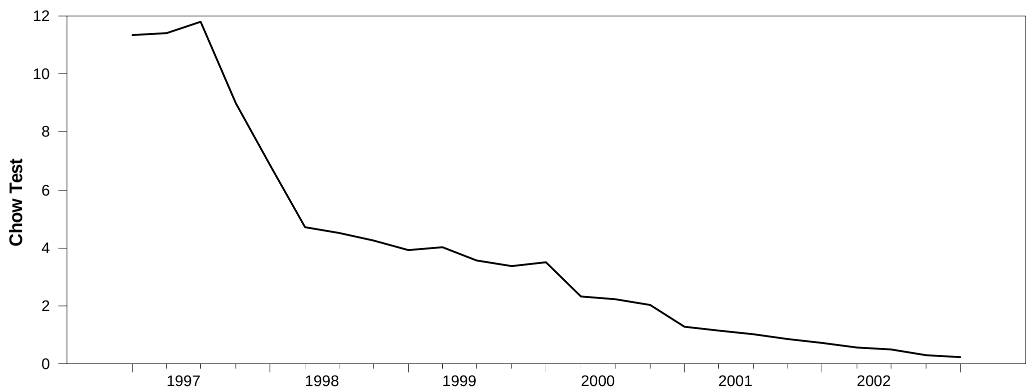
In Figure 3, the *W* test statistic for both *fte* and *ptr* is significantly lower than the 5 per cent asymptotic critical value of 35.95 implying that the parameters constancy null hypothesis is not rejected in any of the candidate break dates. Put differently, we fail to find any evidence of a break in the underlying process of the series and there is no evidence to suggest that the series of minimum wage changes in the period 1997 Q2 to 2002 Q3 has an impact on employment dynamics. The same conclusion can be reached for the other two states (Figures 4 and 5). Our results are subject to an important caveat. Due to the 'trimming factor' constraint in testing for a possible break date in periods subsequent to 2003, it may be that minimum wage changes occurring in the period 2003-2008 could affect employment. Notwithstanding such a caveat, the low value and the observed downward trend in the plot of *W* test statistic are indicative that wage changes occurring in the latter part of the sample are unlikely to exert a significant influence on employment in all three states. In light of the evidence that there is no single break in the employment series, we do not proceed to test for possible multiple structural breaks using Bai's (1997) sequential multiple structural breaks test.

Figure 3: Structural Break Tests for Victoria

Full-time equivalent employment



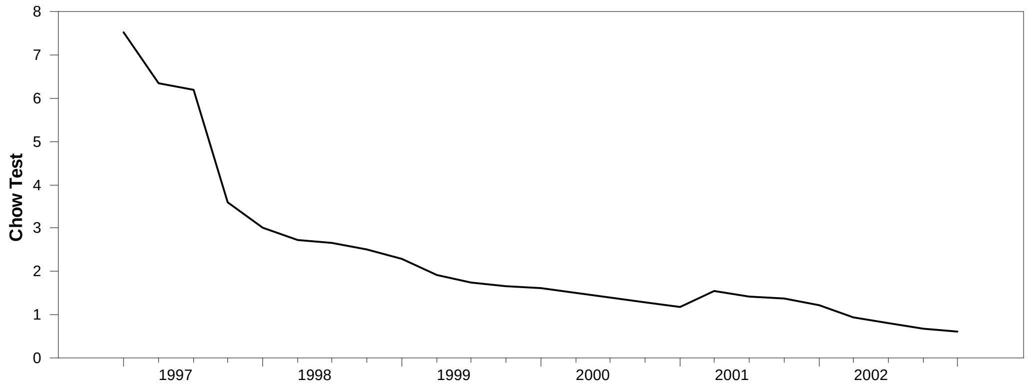
Part-time employment



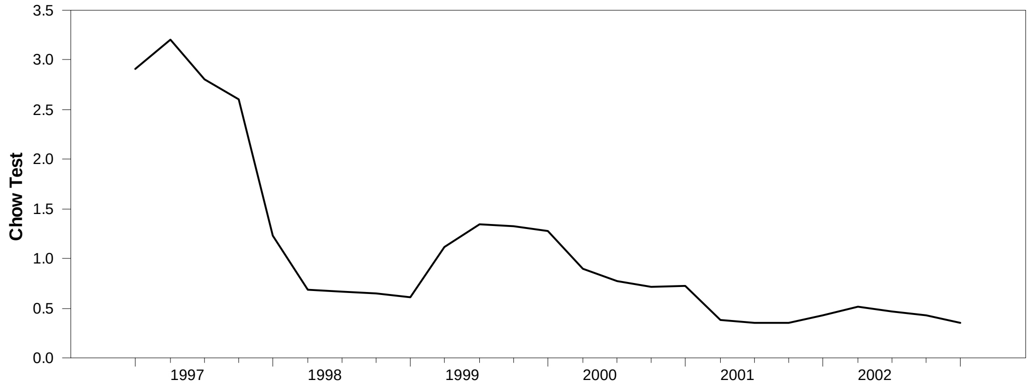
Notes: Sample data from 1992 Q1 to 2008 Q1. 5% critical values from Andrews (1993) are 32.65 for both full-time and part-time employment (based on π = 0.2 and k = 14).

Figure 4: Structural Break Tests for Northern Territory

Full-time equivalent employment



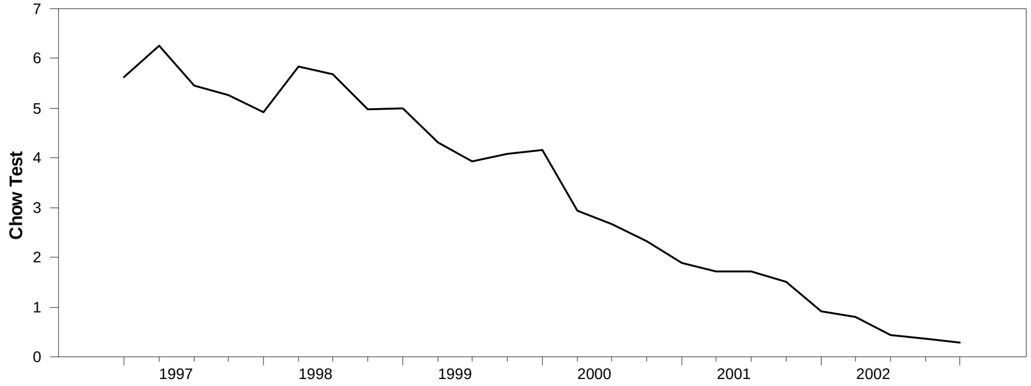
Part-time employment



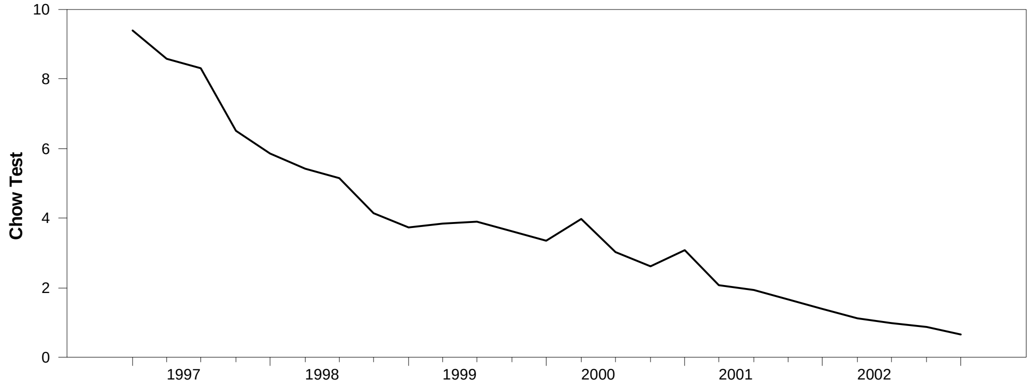
Notes: Sample data from 1992 Q1 to 2008 Q1. 5% critical values from Andrews (1993) are 35.95 for both full-time and part-time employment (based on $\pi = 0.2$ and $k = 16$).

Figure 5: Structural Break Tests for Australian Capital Territory

Full-time equivalent employment



Part-time employment



Notes: Sample data from 1992 Q1 to 2008 Q1. 5% critical values from Andrews (1993) are 34.41 for full-time employment (based on π =0.2 and k = 15) and 35.95 for part-time employment (based on π = 0.2 and k = 16).

6. Robustness Checks

In order to verify the results presented in the previous section, we use our time-series data to estimate a variation of the traditional time-series model used to examine the effects of minimum wages. The general specification of this model is provided in Equation (2). Given our interest in testing the statistical significance of changes in the minimum wage rate on employment, we employ a shorter sample spanning the period 1996 Q1 to 2008 Q1.[14] We also test for the stationarity property of the series for this shorter sample. The results that are provided in the Appendix (Table A1) are by and large similar to those reported in Table 2 for the longer sample. To be consistent with the model specification in the structural break test, we allow for lagged dependent variables in cases where there is evidence of serial correlation in the regression residuals. The estimated models are identical to the specifications employed in the structural break test, with the exception that we include two additional regressors involving the minimum wage and adult average weekly ordinary time earnings. As information on coverage is not available in Australia to enable us to compute the Kaitz index, we use an alternative specification to capture effects of the minimum wage. Following the suggestion of Card and Krueger (1995), we include the real minimum wage (*m*) and the real adult wage (*aw*) as separate independent variables, and interpret the coefficient on the real minimum wage as the effect of minimum wages.

The results for the model estimated are displayed in Table 3. This shorter time series is used because data on minimum wages from 1978 are not available (and data from the C14 award rate is used as a proxy for the minimum wage in 1996). The diagnostic tests indicate that our model is free from any problem of serial correlation or ARCH effects. Consistent with the results of the structural break tests, the coefficient on the minimum wage variable is never significant, implying that changes in minimum wages appear to have no negative employment effects.

[14] We perform a robustness check of our regression results using the sample period for the structural break test. The results are qualitatively unchanged. These results are available from the authors upon request.

Table 3: Robustness Check Using the Traditional Time-Series Approach

	Full-time			Part-time		
	Victoria	NT	ACT	Victoria	NT	ACT
Intercept	0.202	0.472	0.273	-2.137*	-0.723	5.657*
	(0.388)	(0.909)	(1.760)	(0.470)	(1.655)	(1.703)
Y(-1)	-	0.815*	0.374*	-	0.763*	0.506*
		(0.099)	(0.120)		(0.124)	(0.113)
Y(-2)	-	-0.439*	-	-	-0.348*	-0.427*
		(0.142)			(0.108)	(-0.138)
Q2	0.059	-1.684	-1.719**	0.368	0.996	-3.445**
	(0.424)	(1.287)	(0.823)	(0.726)	(2.845)	(1.590)
Q3	-0.737***	1.929**	-1.235***	-0.398	-1.409	-0.413
	(0.420)	(1.049)	(0.656)	(0.526)	(1.952)	(1.128)
Q4	-0.558	-2.261	-0.964	0.119	-3.894***	-3.668*
	(0.348)	(1.734)	(0.640)	(0.524)	(2.256)	(1.041)
T	-0.007	-0.022	-0.002	0.016***	6.72×10^{-4}	0.002
	(0.007)	(0.019)	(0.009)	(0.009)	(3.36×10^{-2})	(0.021)
T^2	4.49×10^{-5}	1.12×10^{-4}	3.01×10^{-5}	5.89×10^{-5}	1.13×10^{-5}	5.04×10^{-6}
	(3.48×10^{-5})	(9.64×10^{-5})	(4.56×10^{-5})	(4.43×10^{-5})	(1.71×10^{-4})	(1.11×10^{-4})
Q2T	-0.001	0.035	0.036**	-0.007	-0.026	0.072**
	(0.009)	(0.026)	(0.017)	(0.015)	(0.061)	(0.034)
Q3T	0.018**	-0.035***	0.025***	0.008	0.031	0.006
	(0.009)	(0.020)	(0.014)	(0.011)	(0.041)	(0.023)
Q4T	0.0143***	0.046	0.018	-0.002	0.085***	0.071*
	(0.007)	(0.034)	(0.013)	(0.011)	(0.047)	(0.022)
$Q2T^2$	5.08×10^{-6}	-1.73×10^{-4}	1.83×10^{-4}**	4.31×10^{-5}	1.59×10^{-4}	3.58×10^{-4}**
	(4.90×10^{-5})	(1.38×10^{-4})	(8.94×10^{-5})	(7.55×10^{-5})	(3.18×10^{-4})	(1.78×10^{-4})
$Q3T^2$	1.06×10^{-4}**	1.60×10^{-4}	1.22×10^{-4}***	4.17×10^{-5}	-1.62×10^{-4}	1.91×10^{-5}
	(4.86×10^{-5})	(1.10×10^{-4})	(7.46×10^{-5})	(5.51×10^{-5})	(2.08×10^{-4})	(1.20×10^{-4})
$Q4T^2$	8.29×10^{-5}**	2.31×10^{-4}	8.58×10^{-5}	5.76×10^{-6}	-4.48×10^{-4}	3.35×10^{-4}*
	(3.93×10^{-5})	(1.71×10^{-4})	(7.08×10^{-5})	(5.43×10^{-5})	(2.48×10^{-4})	(1.14×10^{-4})
m	-0.366	-0.550	-0.411	0.175	2.469	-2.528
	(0.354)	(0.920)	(0.711)	(0.464)	(1.905)	(1.985)
aw	-0.783*	-0.032	0.141	0.378	-3.406**	-0.724
	(0.209)	(0.832)	(0.427)	(0.261)	(1.531)	(0.451)
tpop	0.205	-0.446	0.509	3.686**	-5.878	3.243*
	(1.038)	(2.291)	(0.859)	(1.767)	(5.076)	(0.911)
unemp	0.034	-0.058***	-0.007	0.043	-0.012	-0.042
	(0.040)	(0.032)	(0.052)	(0.060)	(0.057)	(0.057)
Diagnostic tests						
Q(1)	0.290	0.413	0.050	2.342	0.498	0.973
	[0.590]	[0.521]	[0.822]	[0.125]	[0.480]	[0.324]
Q(4)	6.027	4.780	2.146	5.187	5.743	3.369
	[0.197]	[0.311]	[0.708]	[0.269]	[0.219]	[0.498]
$Q^2(1)$	0.482	1.603	1.592	0.514	0.198	0.289
	[0.487]	[0.205]	[0.207]	[0.473]	[0.656]	[0.591]
$Q^2(4)$	2.097	2.136	3.011	2.276	7.192	1.083
	[0.718]	[0.711]	[0.556]	[0.685]	[0.126]	[0.897]
$\bar{R}^2$	0.775	0.478	0.829	0.772	0.457	0.852

Note: Figures in ( ) and [ ] are robust standard errors and p-values respectively. Q(*k*) and $Q^2(k)$ are Ljung-Box test statistics under the null that the residuals and squared residuals are serially correlated with order *k*, respectively. T = 49. Data are from the ABS Labour Force Survey from 1996 Q1 to 2008 Q1.

As a second robustness check, we also redid our analysis in section 5 using the longer period 1978 Q2 to 2008 Q1. For the part-time employment to population ratio in the Northern Territory, we find structural breaks in 1984 Q2 and 1984 Q3. No other breaks were detected for any of the other series. This reinforces the finding using the shorter sample we report in section 5 that it is unlikely the introduction of the federal minimum wage in 1997 led to any adverse employment outcomes.

7. Conclusion

Despite detailed studies of the effects of minimum wages by legions of economists using various alternative approaches, to date, the issue remains highly contentious and politically charged with no clear consensus. Australia formally introduced minimum wage legislation in April 1997. This paper uses tests for structural breaks to determine if there is a significant relationship between minimum wage legislation and employment in the unique institutional setup in Australia. The tentative conclusion is that the seven minimum wage increases in Australia from 1997 to 2003 appear to not have had any significant negative employment effects for teenagers. A possible explanation is that the increases have generally been moderate and predictable, closely tracking the general rise in price levels. Furthermore, for all three states, the initial relatively high values of the Chow statistics in 1997 (but insignificant) and the subsequent downward trend from that point onwards are suggestive of a possible adaptation to the new regime.

More generally, this paper also makes a contribution to the evaluation literature as a whole. Structural break tests are more commonly employed by macroeconomists rather than micro econometricians, but there is no reason why the latter should not be using them more in applied work. Such tests for regime shifts are often conducted when it is basically impossible to create a counterfactual using a comparison group approach. Examples from the macro literature include the analysis of the effects of the abandonment of the Bretton Woods system, and the introduction of the common European currency.

We believe that the techniques employed in this paper are highly applicable to other non-experimental policy scenarios, where relatively long time-series data are available, and where there are no obvious comparison groups because of statewide or nationwide implementation. Importantly, such an evaluation approach might be the only option available in many contexts, where experimental or quasi-experimental designs are impossible to implement.

Appendix

Table A1: Unit Root Tests Results for the Traditional Time-Series Approach

Victoria	*fte*	*ptr*	*tpop*	*unemp*	*aw*	*m*
ADF test						
(a)	-1.538 [2]	-1.338 [3]	-1.558 [2]	-0.299 [4]	-1.401 [0]	-2.673***[4]
(b)	-1.994 [2]	-5.282* [0]	-2.715 [1]	-4.282* [0]	-2.132 [0]	-3.043 [4]
ZA test						
(a)	-3.874 [9] {2005:3}	-7.133* [0] {2004:1}				
(b)	-3.864 [9] {2005:3}	-7.135* [0] {2004:1}				
Northern Territory						
ADF test						
(a)	-3.726* [0]	-4.001* [0]	-0.221 [0]	-3.615*[0]	-2.072 [0]	-2.673***[4]
(b)	-3.976** [0]	-3.976** [0]	-1.596 [0]	-3.575**[0]	-2.064 [0]	-3.043 [4]
ZA test						
(a)	-4.198 [5] {2005:3}	-5.738* [3] {2006:3}				
(b)	-3.935 [5] {2004:3}	-5.018*** [3] {2006:3}				
Australian Capital Territory						
ADF test						
(a)	-1.319 [0]	-4.036* [0]	-0.802 [5]	-0.231 [3]	-1.212 [0]	-2.673***[4]
(b)	-4.447* [0]	-4.527* [0]	-3.699** [1]	-4.503*[0]	-2.930 [0]	-3.043 [4]
ZA test						
(a)	-5.392* [0] {2000:1}	-6.081* [0] {2000:3}				
(b)	-5.256** [0] {2000:1}	-6.071* [0] {2000:3}				

Notes: ADF = Augmented Dickey-Fuller Test. ZA = Zivot and Andrews test. (a) and (b) denote tests are conducted with intercept, and with both trend and intercept respectively. Figures in [.] represent the AIC-selected lag length. *,** and *** denote significance at 1%, 5% and 10% levels. The ZA test has a null hypothesis of a unit root with no break, and an alternative hypothesis of stationarity with a single break. The figure in {} under ZA test denotes break date. The 1%, 5% and 10% critical values for the ZA test with break(s) in intercept (intercept and trend) are -5.34 (-5.57), -4.80 (-5.08), and -4.58 (-4.82) respectively. The sample period for the unit root tests is for 1996 Q1 to 2008 Q1.

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